EERI
Economics and Econometrics Research Institute

Food Prices and Overweight Patterns in Italy

Luca Pieroni, Donatella Lanari and Luca Salmasi

EERI Research Paper Series No 40/2010

ISSN: 2031-4892

EERI
Economics and Econometrics Research Institute
Avenue de Beaulieu
1160 Brussels
Belgium

Tel: +322 299 3523
Fax: +322 299 3523
www.eeri.eu

Food Prices and Overweight Patterns in Italy

Pieroni L.*, Lanari D.**, Salmasi L.***

*University of Perugia (Italy) and University of the West of England (UK)

**University of Roma-La Sapienza (Italy)

***University of Verona (Italy)

Corresponding author: Pieroni Luca, Department of Economics, Finance and Statistics,

Via Pascoli 20, 06123, Perugia, Italy. E-mail: lpieroni@unipg.it

Abstract

In this paper we examine the role of relative food prices in determining the recent increase in body weight in Italy. Cross-price elasticities of unhealthy and healthy foods estimated by a demand system provide a consistent framework to evaluate substitution effects, when a close association is assumed between unhealthy (healthy) foods and more (less) energy-dense foods. We used a dataset constructed from a series of cross-sections of the Italian Household Budget Survey (1997-2005) to obtain the variables of the demand system, which accounts for regional price variability. The relative increase of healthy food prices was found to produce nontrivial elasticities of substitution towards higher relative consumption of unhealthy foods, with effects on weight outcomes. In addition, these changes were unevenly distributed among individuals and were particularly significant for those who were poorer and had less education.

Keywords: Household Survey Data, Healthy and Unhealthy Foods, Overweight and Obesity, Elasticity of Substitution

JEL classification : D12, I10, I18

1 Introduction

In developed countries, the patterns of the excessive weight and obesity of adults and children have become a critical public health concern. Data from the United States show that the prevalence of overweight and obesity began to increase around the mid-1980s, and has continued to increase dramatically, despite the fact that technological advances in medicine have improved the general health of the population in the past half-century. It has been estimated that 3.7 and 5.3 percent of medical spending are attributable to overweight and obesity, respectively (Finkelstein et al., 2003; Daviglus et al., 2004).

In European Union countries, it has been calculated that further extension of the health care system would reach costs of about 7 percent of the total healthcare expenditure (Commission of the European Communities, 2005). Although much later with respect to the United States and some continental European countries, the overweight issue has also become significant in Italy (Costa-Font et al., 2009), where the percentage of adults classified as overweight (including the various forms of obesity) has risen by 4 percent in the last decade (ISTAT, 2007).

However, the complex range of social and economic factors which explain the patterns of obesity are not understood and identified, although the fact that a vast body of studies has recently been attempting to explore its determinants (Kopelman, 2000; Lakdawalla and Philipson, 2009; Powell and Chaloupka, 2009). The research programs quoted in the works of Cutler et al. (2003) and Bleich et al. (2008) explain rising rates of obesity as the result of over-consumption of calories, associated with technological changes which lower the unitary cost of food.

One strand of empirical works focuses on economic determinants to respond to the question of whether changes in food prices translate into changes in the prevalence of overweight and obesity. These tests provide an interesting picture of the long-run association of food prices and overweight changes such as the use of aggregate data cannot identify heterogeneous individual (or household) behaviour and may result in a policy-maker receiving weak or unclear information.

Conversely, it would be interesting to obtain direct estimates of overweight on changes in food consumption by changes in relative food prices, with a focus on the segmented groups of a society (Schroeter et al., 2008). The simultaneous absence in Italy of true panel data and discontinous surveys in measuring individuals weights (and heights, i.e. body mass index) does not enable us to account for this behaviour.

Therefore, we take an alternative approach in evaluating whether changes affect body weight. The growing availability of microdata from household expenditure surveys and its role in allocating food expenditure allows us to model the dynamic of food consumption (or its categories) consistently and to exploit the richness of survey data information. This approach has also recently been used to estimate the relationship between household food consumption and obesity in France (Bonnet et al., 2008). In our case, it is worth noting that the choice to analyse changes in food behaviour based on household consumption also appears to be consistent with the statistical representativeness of food consumption. Indeed, the share of food expenditure of households allocated for food away from home has remained constant during the last decade, and ranges from 14 to 16 percent of total food expenditure[1].

In this paper, we use the monthly consumption data of nine Italian Household Budget surveys (IHBS) (1997-2005), to model a micro-economic demand system for foods (and non-foods). In particular, by separating a composite category of expenditure on healthy food from unhealthy ones (and residual foods and non-durables), we shed light on the latent mechanisms of substitutability, based on the differences between the price patterns of these goods, and the conditions in which total calories rise when the relative price of the unhealthy food category - associated with energy-dense foods - falls.

Our research is consistent with the hypothesis of Cutler et al. (2003), i.e. technological changes have lowered the relative full prices of mass-produced foods and favoured the consumption of energy-dense ones. Our work is also related to the work of Auld and Powell (2009), who model food consumption in such a way that changes in relative prices change their contribution in terms of total calorie intake solely by their compensated (or Hicksian) price elasticities. Our work is methodologically close in spirit to the work of Zheng and Zhen (2008) which, within a demand system, with three categories of consumption for the United States and Japan, comprises the components of healthy and unhealthy foods and tests the influence of substitution effects on these food categories.

This paper makes a number of contributions. First, to answer the important

[1]Recent research programs have also used scanner data from grocery stores to estimate price elasticities, with observation of prices paid for individual transactions. Although within-category substitution of food would be much more meaningful as consumers can substitute more easily towards similar goods, it is doubtful whether this segmentation would lead to more effective policy implications. In a plethora of food products on offer, many items are generally not consumed by households, so that the problem of possible inefficiencies in implementing food policies to reduce the frequency of overweigh, while recognized in the literature, needs to be resolved. One way of doing this should involve modelling households' zero expenditure.

question as to whether shifts observed in relative prices affect consumer behaviour in food choices, we examine whether the estimated cross-price elasticities of unhealthy and healthy foods are significant for Italy.

Second, unlike Zheng and Zhen (2008), who estimate a demand system by an aggregate national price index, we use a monthly regional price index associated with data of household expenditure and demographic and socio-economic characteristics. This choice reflects the fact that the variability of the demand system obtained from regional prices is much larger than that of prices derived at national level. Our empirical strategy addresses the inference problem in estimating demand equations which may arise from insufficient price variations - a problem which is particularly severe when dealing with macro-data.

Third, we test the hypothesis that the rise in unhealthy food consumption is partly responsible for the increased body weight of Italian adults. In turn, this hypothesis assumes, that unhealthy foods are more dense in energy. A nutritional method is used to transform the consumption of healthy and unhealthy food categories and to clarify the relationship with light and more energy-dense foods, respectively.

A final contribution is to extend the estimations of cross-price demand elasticities, explaining how they vary according to gender and socio-economic factors. Collecting time-series of cross-sections allowed us to determine consistently whether changes in food prices alter dietary intake patterns sufficiently to explain weight gains by sub-groups[2]. Previous research, exploring the rise of disparities in overweight and obesity among social groups, supports the hypothesis that, the less education an individual receives, the greater the overweight recorded (Klesges et al. 1998; Dennis et al., 2000; Hardy et al., 2000; de Saint-Pol, 2009). These heterogeneous responses among groups appear to be confirmed according to individual standards of living (de Saint-Pol, 2009), whereas estimations of household income are not conclusive (Ball and Crawford, 2005; Villar and Quintana-Domeque, 2009). The existence of the confounding effects of age and gender in individual body weight identified by Bray (1987) and de Saint-Pol (2009), respectively, justify our choice to test the patterns of substitution elasticities separately.

Our main results are that the estimated elasticities of substitution do have the expected effects. We cannot reject the hypothesis that falling relative unhealthy food prices is one probably cause of the prevalence of overweight in Italy, leading people

[2]For a review of longitudinal studies of socio-economic status and weight, see Ball and Crawford (2005).

to substitute away from healthy towards unhealthy foods even if the magnitude of the cross-price elasticity is not large. These changes appear to be consistent in aggregate although they do not affect individuals equally and turn out to be particularly important for poorer and less educated people. Thus, the results call into question whether the government should intervene in the food market. Section 2 motivates our analysis by the descriptive analyses. It illustrates some descriptive facts regarding the recent increase in overweight of the Italian population, matched with patterns of relative healthy/unhealthy food prices. Section 3 presents the model of interest and shows that an almost ideal demand system (AI) generally leads to a consistent framework, even when it is extended to obtain long-run measures of elasticities. Section 4 derives a long-run demand system from the dynamic system approach and the identification issue concerning price elasticities is argued. Section 5 provides a range of cross-price elasticities and discusses policy implications. Section 6 concludes.

2 Basic facts

We begin this section by illustrating the patterns of the prices and quantities of healthy and unhealthy foods purchased in Italy. The analysis serves to motivate a theoretically consistent demand model, which follows in Section 3 and is based on the classification (other than a residual food category) provided by Gelbach et al. (2007) and adapted for features of Italian food consumption. We report the details of this classification in Appendix A. This classification is built to highlight meaningful substitution effects between food with high contents of fat or sugar (i.e., unhealthy food) with respect to food with fewer calories (i.e., healthy food). It reflects, for example, a strategy addressed to assess policy implications of taxes in some unhealthy categories as their consumption grows when the prices of this category food decrease. Although bread, pasta and olive oil are foods with high calorie contents, we exclude these categories because they do not reveal a clear link with unhealthy food. Rather, they are highly characteristic of the Mediterranean diet, in which the habits and traditional aspects of Italian consumers preserve their consumption behaviour. In addition, as discussed in Mazzocchi et al. (2008), medical studies have shown that the balanced composition of these foods within the Mediterranean diet leads to better conditions in terms of morbidity and general health.

As a benchmark of our model, we include these foods in a residual category, together with other non-durable goods, so that the link between healthy - energy-light foods and unhealthy - more energy-dense ones is unambiguous The empirical section then provides a sensitivity analysis of the estimates of substitution effects by first including bread, pasta and olive oil in the unhealty food category and then in healthy food one.

Figure 1 shows the overlapping paths of the fraction of overweight and obese adults and the relative changes in healthy and unhealthy food prices. For this purpose, we match two sources of data. First, we use the annual multipurpose survey of Italian households (2002-2005) conducted by the Italian Institute of Statistics (ISTAT) to obtain an average of the Italian body mass index (BMI). We then integrate these measures of individual BMI with data from two more general ''multipurpose surveys" of 1994-1995 and 1999-2000, and interpolate the values for the missing annual surveys. Note that these breaks do not allow us to use the BMI indicator as a direct (long-run) body weight response of changes in food price categories [3].

Instead, the annual mean of expenditures and prices for these two categories of consumption are obtained by extracting microdata from the IHBS, released annually by ISTAT. Because the latter surveys are the main sources for the empirical section, we present detailed data below, merely noting that the relative prices of unhealthy and healthy foods are obtained by the ratios between nominal and real expenditures, and that real expenditure, as a consumption index, aggregates the ratios between the current expenditure for each individual item and its index price (at national level). The graph, shows a constant increase in BMI over time, although upward trends are not recorded. As an aggregation of the repeated Italian household surveys, the picture which emerges shows that the prices of healthy foods rise more quickly than those of unhealthy ones. Descriptively, the increase in healthy food versus unhealthy food prices is, on average, 1 percent per year.

According to the hypothesis that, as unhealthy food becomes relatively cheaper, people are expected to substitute away from healthy foods towards unhealthy ones, with greater consumption in the latter category. Relative consumption, expressed as a constant expense, is reported in the continuous line of Figure 2 (*i.e.*, vice versa with respect to relative prices). The predictions of the theory are supported by the slight upward positive trend in the growth of unhealthy food consumption during

[3] The body mass index (BMI) is a measure of body fat based on height and weight, which applies to both adult men and women. Four categories are generally used to classify adults: i) Underweight BMI ≤ 18.5; ii) Normal weight = 18.5-24.9; iii) Overweight = 25-29.9; iv) Obesity = 30 or over. It is known that BMI is not the most accurate measure of body fat and that self-reported weight produces measurement errors for young and adult people. For a critical discussion of this indicator, see Burkauser and Cawley (2008).

the period 1997-2005.

Although we cannot directly retrieve the effects of food prices on weight outcomes, we follow the schemes used by Chesher (1998), transforming the categories of food consumption into energy (*i.e.*, calories). By retrieving information from the Italian Institute of Nutrition (INN, 1997), we can aggregate food categories, construct relative patterns, and associate then patterns of body weight. In particular, this source provides data for over 1000 food items, which were first aggregated into 62 items - the same source employed by ISTAT in its surveys - and further aggregated into the categories of healthy and unhealthy foods considered in this study. The dotted line of Figure 2, highlights the evolution of the energy changes (expressed in calories) of unhealthy with respect to healthy foods. These patterns are close to those of unhealthy versus healthy food consumption, indicating that changes in food prices translate into changes in quantities of calories consumed by individuals, which may contribute to explaining the increase in the share of overweight individuals, in line with the argument of Drewnowski and Darmon (2005).

Although the patterns of the figures seem to be consistent with the working hypothesis, our approach differs conceptually from that of works which use representative time-series data to measure the elasticity of substitution (*e.g.*, Zheng and Zhen, 2008), assuming the existence of several effects of a unitary change in relative prices on food budget share allocation across groups of consumers. For this reason, the rest of this study focuses on measuring the elasticity of substitution by grouping the household expenditures of the Italian surveys and estimating them according to socio-economic and demographic characteristics.

At descriptive level we use the annual multipurpose surveys of Italian households (2007) to discuss average changes in the prevalence of obesity and overweight for the selected socio-economic and demographic groups. It is argued from this source that the problem of overweight affects males specifically and increases with age, although it should be noted that it is important at all ages. The rise in BMI disparities among social groups is also reflected at educational level. In recent period, the only Italian adult people who did not increase in the share of overweight and obesity were those with higher education, whereas less educated adults showed a greater risk of obesity.

It is worth noting that, since the patterns of BMI shared by household income or expenditure levels are not reported in the annual multipurpose survey of households, we can only assume the hypothesis of food substitution in these subsamples, but cannot compare the patterns of relative prices of consumed quantities and the relative transformed energy extracted by IHBS. Thus, the assumption that individuals of lower economic status cover calorie requirements more easily by purchasing high-calorie products, which are often cheaper than low-calorie ones, is at work. (Drewnowski and Specter, 2004)[4].

Figure 3 illustrates these points by dividing the sample into several socio-economic characteristics - gender, age, education and economic status - and collecting household data for each subsample. Naturally, gender is shared by males and females, education by household heads with a university degree, Master or PhD with respect to a lower level of education, age by four classes of adults (18-34; 35-49; 50-64; ≥ 65) and economic status by households which are below or above the relative poverty line.

Panel a) shows that shifts in relative prices (healthy/unhealthy) for males are related to an increase in unhealthy food consumption and, on average, of total calories; the influence on energy is less important for females. These results also find confirmation in educational and economic status. Greater sensitivity to healthy/unhealthy price changes regarding the consumption of unhealthy foods is shown for households with lower education (panel b) and lower income (panel c), with quicker growth of the sub-samples in the last few years. Panel d) shows the patterns of price changes by age. Although all the classes respond quickly to changes in healthy food prices, the age classes in which the potential substitution of healthy foods with unhealthy ones is more evident are those between 35-49 and 50-65.

These observations impose testable extensions on demand models: if the data predict that significant price substitution effects of healthy versus unhealthy food consumption partly justify the pattern increase in Italian overweight, then: (a) the use of a regional price index for the categories of healthy and unhealthy foods increases the variability of food response in estimating demand equations, which are particularly severe when dealing with the aggregate price index; (b) the existence of specific patterns of the elasticity of substitution may be assessed within demographic and socio-economic groups as well as comparative estimations with respect to those of the whole economy.

This extension therefore provides a general framework for testing the hypotheses

[4]The subsample for approximating low income households was obtained following the ISTAT procedure of calculating the relative poverty threshold, according to which, a household is considered under the threshold if its total per-capita expenditure (total expenditure over the number of components) is lower than the average per-capita expenditure. An equivalence scale was used to determine the relative poverty threshold for households with a number of components other than two.

of the paper empirically and can support policy interventions selectively. The rest of the paper is devoted to applying this framework to one of the major demand systems proposed in the literature, the *AI* model of Deaton and Muellbauer (1980).

3 Theoretical background

We briefly review the static *AI* model. The specification of this demand system arises from a class of preferences in the logarithm of total expenditure, known as 'price independent generalized logarithmic'. It satisfies the necessary and sufficient conditions for consistent aggregation across consumers (Muellbauer, 1976; Deaton and Muellbauer, 1980) and allows the estimation of demand elasticities with limited restrictions (Deaton, 1986).

It is assumed that there are n goods which can be purchased by consumers and which can potentially be included in the demand system. We index these categories of goods (and services) by $i = 1.., n$. We note that choosing the goods to be included in the demand system depends on the purpose of the paper. Because we are interested in the reaction of relative consumption between two macrocategories of foods, *i.e.*, healthy and unhealthy, this classification determines what sub-categories of goods are included in the *AI*; its dimensionality is of negligible interest.

Analytically, the budget share of a certain good is equal to the expenditure generated by the good divided by the total expenditure of the categories of goods included in the demand system. We use w_i to represent the budget share of a good i, $i = 1, ..., n$. Under the *AI* specification, w_i takes the form of:

$$w_i = \alpha_i + \sum_{j=1}^{n} \gamma_{ij} \log p_j + \beta_i \log (Y/P) + \nu_i \tag{1}$$

where p_j are $j = 1, ..., n$, are the prices of goods, Y is the total expenditure in the demand system, P is an overall price index for the goods, ν_i is a stochastic error, and α_i, β_i, and γ_{ij} are parameters to be estimated. Note that the last term of (1) is based on the real expenditure ($Y/P = y^*$) devoted to category w_i. The budget share of product i increases as the total real expenditure of the category increases if β_i is positive, and decreases if β_i is negative. The second term is based on the price effects of the various goods. We will return to this point later, after introducing all the ingredients of this flexible demand system to estimate the (cross) price elasticities of demand and to identify the patterns of substitution between goods.

We then define w_i^* as the "optimal" level of the observable expenditure budget share w_i for commodity i and $\log P = \alpha_0 + \sum_{k=1}^{n} \alpha_k \log p_k + \frac{1}{2} \sum_{k=1}^{n} \sum_{j=1}^{n} \gamma_{kj} \log p_k \log p_j$. As commonly done in empirical papers, we also employ a linear approximation in this price index, *i.e.*, Stone's price index, defined as $\log P = \sum_{i=1}^{n} w_i \log p_i$.

As discussed in the review by Barnett and Serletis (2008), the *AI* flexible model has a number of desirable properties. It derives an expenditure function from a second-order approximation to any expenditure function and provides the possibility of including the theoretical restrictions of adding up, homogeneity and symmetry in order to respect the predictions of the demand theory. Because the expenditure function must be linearly homogeneous and strictly increasing in p, adding up and homogeneity can be explained as $\sum_{i=1}^{n} \alpha_i = 1$ and $\sum_{i=1}^{n} \gamma_{ij} = \sum_{j=1}^{n} \gamma_{ij} = \sum_{i=1}^{n} \beta_i = 0$, respectively, while symmetry requires $\gamma_{ij} = \gamma_{ji}$ for all i, j.

Another property of a robust expenditure function (and demand system) is that it must be concave in prices. This means that the matrix of the second-cross partial derivatives must be negative semi-definite. In turn, this property gives rise to the matrix of the substitution effects of Slustky, $S_{ij} = \partial h_i(p, u)/\partial p_j$, with non-positive own-price effects , where $h_i(.)$ is Hicksian demand. Formally, it can be shown that the Slutsky substitution coefficients of model (1) are given as:

$$S_{ij} = \frac{y}{p_i p_j}[\gamma_{ij} + w_i w_j - \delta_{ij} w_i] \tag{2}$$

where δ_{ij} is the Kronecker parameter ($\delta_{ij} = 1$ if $i = j$, and $\delta_{ij} = 0$ if $i \neq j$). The matrix of substitution effects for the *AI* model varying with data determines that negativity conditions must be evaluated (and eventually imposed) locally at a specific point in the sample[5]. That is, by scaling the data at a representative point (e.g. the mean of the sample) in which $P = y^* = 1$, we can obtain the local substitution term $\theta_{ij} = S_{ij}(P = y^* = 1)$[6].

Equation (1) is singular by construction, as the expenditure shares sum to 1. A frequently employed procedure to avoid econometric problems consists of dropping one equation from the system and, although the budget share demand system with a $n - 1$ rank must be empirically confirmed, it provides complete characterisation of consumer preferences. Consequently, it can be used to estimate the income, own- and cross-price elasticities as well as the elasticities of substitution.

[5]See Cranfield and Pellow (2004) for a more thorough discussion of the role of global and local negativity in functional form selection.

[6]In order to remark the properties of the demand system, the *AI* provide a reasonablly accurate approximation at any set of prices not too far from the point of approximation.

4 Econometric framework, data issues and elasticities

This subsection provides econometric support for modelling a long-run demand system which includes a gradual adjustment over time of consumption in response to shifts in relative prices. However, when time-series have a significant dimension, empirical demand system studies suffer from severe econometric flaws, because the time-series of budget shares, prices and real income are non-stationary.

One way of solving these issues is to use linear model cointegration methods (Attfield, 1997, 2004) although they may not be completely consistent, since errors in demand systems tend to be autocorrelated (Lewbel and Ng, 2005). Standard asymptotic theory may provide a poor guide to finite-sample inference when the errors are persistent in a cointegrated demand system. As one aim of the empirical strategy, we complement statistical analysis by investigating and testing the non-stationarity behaviours of the time-series residuals.

There is also a profound policy interest in obtaining parameter estimations from a cointegration framework. They are inextricably linked with the notion of long-run estimation (Pesaran, 1997). Because we are specifically interested in analysing substitutability effects of healthy foods with respect to unhealthy ones, an important question for policy-makers is whether these trends will continue in the future. Indeed, a measure of the long-run elasticity of substitution is in being a powerful tool in assessing potential government intervention in preventing obesity by taxes imposed on unhealthy foods or subsidies applied to healthy ones (Powell and Chaloupka, 2009). In subsection 4.1, we report the conditions for the identification of a long-run demand system based on the cointegration rank of a vector autoregressive (VAR) model under the theoretical constraint of adding-up. We then show that, in this framework, the other theoretical restrictions of homogeneity and symmetry can be imposed and tested, and the estimated parameters recovered to calculate the price elasticities.

4.1 Methods

We formalise the specification of the equations of the demand system in (1) as a cointegrated demand system. Firstly, we consider the vector autoregressive (VAR) formulation of a demand system and describe the corresponding vector error correction (VECM) representation, following Johansen (1995). Formally, the data-generating process for $X_t = (X_{1t}, X_{2t})$ is assumed to belong to the class of VAR models:

$$X_t = \mu_0 + \mu_1 T + \phi_h D_{th} + \sum_{i=1}^{p} A_i X_{t-i} + \varepsilon_t \qquad (3)$$

where $X_{1t} = (w_{1t},, w_{2t}, ...w_{n,t})$ is the $n * 1$ vector of budget shares and $X_{2t} = (p_{1t}, p_{2t}, ...p_{nt}, y_t)$ is the $n + 1 * 1$ vector containing price indices and real expenditure. μ_0 is a $n * 1$ constant term, μ_1 is a $n * 1$ vector of coefficients related to the deterministic trend T, D_{th} is a vector containing deterministic variables (in our application centred seasonal dummies) and ϕ_h the corresponding $n * h$ matrix of parameters. A_i is a matrix of unknown parameters for the lags of X_t, ε_t is a Gaussian white noise process with covariance matrix Ω and p the lag order of the VAR. Equation (3) may be re-written in a VECM form as:

$$\Delta X_t = \mu_0 + \Pi X^*_{t-1} + \sum_{i=1}^{p-1} \Gamma_i \Delta X_{t-i} + \varepsilon_t \qquad (4)$$

where $\Pi = (\sum_{i=1}^{p} A_i - I_p)$, $\Gamma_i = -\sum_{j=i+1}^{p} A_j$,with $j = 1, ...p-1$. The matrix of parameters Π describes the long-run relationships of the VECM among the variables in vector $X^*_{t-1} = [X_{t-1}; D_t; T]$. Γ_i, with $i = 1, ...k-1$, is a vector of parameters which refers to the short-run dynamics of the system ΔX_{t-i}. In known general conditions, VECM equation (4) is formulated as:

$$\Delta X_t = \mu_0 + \alpha \beta^{*'} X^*_{t-1} + \sum_{i=1}^{p-1} \Gamma_i \Delta X_{t-i} + \varepsilon_t \qquad (5)$$

where α is a $n * r$ matrix, $\beta^* = (n + \Upsilon) * r$ matrix and $r(0 < r < Q = 2n)$ is the cointegration rank of the demand system. Υ is the matrix containing deterministic and seasonal components.

The hypothesis of stochastic trends in budget shares predicts a convergence (in the long-run) towards steady state, and these values are proxied by their intercepts (Pesaran and Shin, 2002).

Pesaran and Shin (2002) also show that, to recover exactly the long-run structural parameters of model (5), r restrictions on cointegrating relationships must be imposed on each non-singular demand equation, expressed in budget shares. In this context, the adding-up theoretical constraint executes a crucial role in identifying the structural model, implying a further implicit restriction of the rank of cointegration of the VAR model, i.e., $r = n - 1$. Formally, disregarding deterministic terms, the matrix of cointegration vectors for the demand system is specified as:

$$\beta' = [-W_{n-1}, B] \tag{6}$$

where W is the $n-1*n-1$ budget share matrix, and B the matrix of the parameters for log prices and real income. The long-run identifying restrictions can be imposed in the W matrix. A necessary and sufficient condition indicates that the number of identifying conditions k, is at least equal to r^2, and that the exact identification for a long-run AI demand system requires $r^2 = (n-1)^2$ restrictions. This implies that, for a demand system with three good categories, a diagonal framework for identification may be imposed as:

$$\beta' = [-I_{n-1}, B] \tag{7}$$

where I_{n-1} is the unit matrix, and $k - r^2$ over-identifying restrictions can be imposed and tested directly on cointegrating vectors. In the demand system context, the restrictions are derived from the theory and concern the hypotheses of homogeneity and symmetry (Deaton and Muellbauer, 1980). Appendix B lists the matrices of the dynamic demand system described in (6) and (7) and those with imposed the theoretical over-identifying restrictions.

The maximum likelihood estimations of cointegration matrix β' are carried out by the ML estimator, which is super-consistent and mixed normal. This allows us to test over-identifying restrictions by a log-likelihood ratio statistic which is asymptotically distributed as an χ^2, with degrees of freedom equal to the number of over-identifying restrictions imposed.

As already observed, the cointegration rank for identifying the patterns of adjustment in each budget share implies $r = n-1$. Thus, the rank condition excludes all the cases in which $r < n-1$. Although in the empirical analysis of demand systems it may happen that the rank of cointegration is higher than the number of cointegration relations, i.e., $r > n-1$, in our study this hypothesis is economically inadmissible. This should imply additional cointegration relations involving price variables in the cointegration space. However, the consequence of this over-dimensionality in equilibrium relationships is in contrast with the causal impact of relative prices on the evolution of budget share consumption and in the measure of effects of substitution.

4.2 Data and elasticities

The IHBS provides information about the socio-demographic characteristics and expenditure levels of Italian households. Although this survey is monitored weekly and published on a monthly basis, it does not provide any information about the quantities purchased and the prices relative to the consumption of each good or service. To obtain estimates of price (substitution) elasticities with a lack of survey information, empirical works include household expenditures with aggregate national price indexes. However, this approach requires a long span of cross sectional data to estimate a demand system with sufficient price variation - features in which are almost never available in empirical applications. Aggregate price indexes are highly correlated due to the restricted number of categories of consumption normally analysed in the demand system, leading to the rejection of theoretical restrictions and making estimated elasticities highly uncertain. This problem was examined and discussed recently by Coondoo et al. (2004).

For these reasons, surveys only gathering expenditure data have limited applicability in modern demand and welfare analysis, unless if researchers are able to include sufficient variability in prices. We believe that, by combining a regional price index (RPI) - instead of a national price index (NPI) - with expenditure recorded in the household surveys, we can respond to the issues discussed above satisfactorily, and consistently estimate the demand system of interest made up of three aggregate categories: a) healthy foods; b) unhealthy foods; and c) other foods and nondurables.

The use of RPI is also important for identifying price elasticities. Above we identify the long AI model from a technical standpoint. However, a demand system also requires that variations in prices should be a result of supply shifts. In this context, it is assumed that regional price index differences reflect supply shifts rather than movement along the demand curve. Following the line of argument of Gelbach et al. (2007) and Zheng and Zhen (2008) we also assume for Italy that marketing healthy foods like fruit and vegetables is more expensive than marketing unhealthy ones like fats and oils, because the former entail higher costs in transport, refrigeration, labour and packaging, and are much more prone to spoilage. But, in addition, the use of a regional price index allows us to account for heterogenous differences in food production and distribution across the Italian regions, and enable us to identify the preference parameters for healthy food separately from unhealthy food.

Data on the consumer price index for the whole collectivity (NIC in ISTAT methodology) is published by ISTAT on a monthly basis. It includes more than 100 goods and services and allows us to aggregate sub-categories of goods consistently. This (chained) price index has been available at regional level since January 1999 and spans until December 2005, the last household survey currently available.

In addition, since the household survey was revised in 1996, we were obliged to use the Italian household budget survey starting from 1997. However, to avoid any restrictions on the time-dimension of the sample, we approximate a monthly consumer price index at regional level for 1997 and 1998. ISTAT has also published the price index for aggregate goods for each province starting from January 1996, referring to the consumption habits of the members of households whose heads are workers or white-collar workers in non-agricultural sectors. We matched the monthly regional elementary price index with the households interviewed in a given month and region, grouping them according to the categories of goods selected in the demand system.

Figure 4 shows the graphs for aggregate budget shares w_{it}, with seasonal adjustments obtained by the X12 census procedure, highlighting the fact that at least some of these shares may also appear non-stationary. This study confirms the systematic evidence of non-stationary in the variables of the demand system (Ng, 1995; Lewbel, 1999; Attfield, 1997, 2004). But, as discussed in Lewbel and Ng (2005), because budget shares must, by construction, lie between 0 and 1, they cannot remain non-stationary forever. The small changes that take place from month to month imply that budget share changes can therefore approximate a non-stationary process for a long time, as is the case of budget shares in the Italian non-durable data.

Although the demographic and socio-economic samples seem to have some specificity, we extend the assumption that the variables of the demand system are non-stationary because the patterns of the variables related to the subsamples are close to those of the full sample and, as noted above, changes in the budget share subsamples are very small[7]. We therefore proceed with the empirical analysis by testing the cointegrating rank of system (5), as a test for identifying the long-run demand system, irrespective of whether it is applied to the data of the whole population or only to subsamples.

In order to confine the analysis of a demand system to non-durables and to

[7]Results of formal tests of non-stationarity with the FDFGLS test of Elliot, Rothenberg and Stock (1996) and KPSS (Kwiatkowsky et al., 1992) for the full sample and subsamples are not presented here, but they are consistent with the hypothesis that budget shares (and also prices and real expenditure) contain unit roots. Both estimations and graphs of budget share subsample patterns are available upon request.

estimate elasticities of demand, we assumed a multistage decision process (Edgerton, 1997). The estimated system is conditional on the choice of purchasing goods that consists of non-durables in a previous non-modelled stage and is determined as part of consumers' overall decision regarding how to allocate expenditure across the full range of goods. In fact as the expenditure is allocated within non-durables, only in the second (conditional) stage consumers decide how to allocate across healthy foods, unhealthy foods, and other foods and non-durables[8]. In this case, income elasticity and Hicksian price elasticities are computed as follows:

$$\eta_i = \frac{\partial q_i}{\partial y}\frac{y}{q_i} = \frac{\partial w_i}{\partial \log y}\frac{1}{w_i} + 1 \tag{8}$$

with $i = 1, 2, 3$. The expression for expenditure elasticity indicates that a good is a luxury if $\eta_i > 1$ and a necessity if $\eta_i < 1$. The compensated price elasticities are given as:

$$\eta_{ij} = \frac{\partial h_i}{\partial p_j}\frac{p_j}{h_i} = \frac{\partial w_i}{\partial \log p_j}\frac{1}{w_i} + w_j - \delta_{ij} \tag{9}$$

where the partial derivatives in (8) and (9) are obtained from equation (7) and i,j=1,2,3. Then, the uncompensated price elasticities are obtained from the Slutsky equation, $\varepsilon_{ij} = \eta_{ij} - w_j\eta_i$.

One formal way of testing the substitution effects between healthy and unhealthy foods is to calculate their cross-price elasticities. According to Hicks (1936), $\eta_{ij} > 0$ indicates substitutability among goods, $\eta_{ij} < 0$ complementarity, and $\eta_{ij} = 0$ independence.

One important property of the Slutsky equation is that the estimated parameters of the matrix are symmetric. But unlike Allen's (1938) elasticity of substitution, the Hicksian framework does not impose symmetry restrictions in elasticity terms. In line with the aims of this paper, we can verify how changes in the prices of unhealthy versus healthy foods simultaneously affect the relative cost of purchasing quantities of these categories of foods. The framework is therefore close to that of Auld and Powell (2009). If, as shown in section 2, changes in healthy food prices are higher than in unhealthy food prices during the last decade in Italy, we can predict greater total calorie intake, generated by a substitution towards unhealthy foods, which determines an increase in the consumption of energy-dense foods. In addition,

[8]In a complete demand system, we should consider previously at least the choice of how to allocate total expenditure between goods and services for consumption (Gorman, 1995).

because the elasticity of substitution is the percentage change in the budget share allocated to good w_i, divided by the percentage change in price p_j, we can compute the net effects of healthy and unhealthy budget share responses to changes in prices. We use η_{12} to represent the elasticity of substitution of the healthy food to changes in unhealthy food prices; η_{21} represents the response of the unhealthy food budget share to changes in healthy food prices. Recalling that the food prices of both categories increased in the period 1997-2005, the net effect of this elasticity on the sample mean is given as:

$$\Delta\eta = \eta_{21} - \eta_{12} \tag{10}$$

If the cointegrated AI is exactly identified, when we extend estimations at each (monthly) point of the sample, we can use the estimations of the elasticity of substitution to evaluate the dynamics of the response of healthy and unhealthy food consumption to relative price changes by the scaling procedure of estimation discussed above. In the condition that the concavity condition is satisfied, we consistently reconstruct the elasticities of substitution (η_{21} and η_{12}) for both the full sample and for the subsamples and their confidence intervals by bootstrapping the standard errors of elasticities.

5 Results and discussion

The long-run AI model considers the share equations in equation (1) as the long-run equilibrium relationships of a VAR model, $X_t = (X_{1t}, X_{2t})$. Following the discussion in Section 4, cointegration test procedures were used to identify the number of long-run empirically important relationships in our data. We only note that exact identification requires two cointegrating relations associated with n-1 (3-1) budget shares among the variables of the model, so that we specify a VAR(3) to evaluate it[9].

The test results of the trace statistic of Johansen (1995) and Saikonnen and Lütkepohl test (SL) (2000) are listed in Table 1. We report both tests from data obtained by aggregations of good categories across households of national or regional indices. At the five percent significance level, neither the trace statistic nor the SL test reject the hypothesis that there are two cointegrating vectors among the

[9] To select the order of the VAR, we used the sequential modified likelihood ratio (LR) test as in Lütkepohl (1991), while estimations are carried out by including centred seasonal dummies.

variables of the demand system, irrespective of the price index used. Thus, by assuming that r=2, the long-run model (5) provides a consistent representation for assessing the significance of the effects of substitution between healthy and unhealthy food expenditure categories.

The parameter estimates of the demand model, which aggregates non-durable goods according to the regional price index (RPI), are listed in Table 2a; those obtained from the data aggregated by the national price index (NPI) are shown in Table 2b[10].

Although almost all the parameters of the two models estimated with imposed homogeneity and symmetry restrictions are significant with similar size after their imposition, we find that the theoretical restrictions are not jointly rejected at 1 percent only for the data obtained by RPI[11]. Thus we proceed below to the estimation of elasticities of the long-run demand system by using variables obtained from the regional price index.

From the stationary and error serial correlation criticisms of Lewbel and Ng (2005), generally found in static demand models estimated with aggregate data, show the patterns of estimated error vectors and related residual serial correlation tests of VECM. Figure 5 plots the resulting estimated disequilibrium errors (w_i^*-w_i), with $i = 1, 2$. Consistent with the assumptions of stationarity of errors, a stable dynamic is found. Furthermore, both Q-statistics, adjusted Q-statistics and multivariate LM statistics, Table 3, indicate the absence of any significant autocorrelation in the vector of the errors.

Table 4 lists the estimated compensated price and expenditure elasticities computed at the sample means[12]. A few aspects of these estimations should be noted. We estimate negative and large own price-elasticities. These results show that there are no violations of concavity and that consumers' demand for food responds to price changes. It is worth noting that the smaller compensated price elasticity in the residual component is strongly biased downwards by the inclusion of the expenditure categories for bread, pasta and olive oil (although the size is reduced when we compare it with the uncompensated price elasticities). We will return to the

[10]The third cointegrating vector for other foods and non-durables is then recovered by the adding-up constraint.

[11]The differences in the results of the theoretical restriction tests are emphasized when small sample statistics are performed. In this case, although the model estimated with a national price index is still rejected at one percent, data which use a regional price index do not reject homogeneity and symmetry at five percent. These results are available upon request.

[12]Typically, one chooses this point to hold concavity because it is the point with the highest sample ''information" and the data are scaled consequently. Asymptotic standard errors of elasticities and confidence intervals are derived from bootstrap replications of the estimated parameters and their standard errors.

robustness of results below, to assess the sensitivity of these estimations. Demand for healthy food is a luxury, whereas unhealthy food is a necessity. The estimated expenditure elasticities are in line with the findings of Zheng and Zhen (2008) in the United States although a different classification (and habits) were responsible for differences in impact measures. Lastly, the cross-price elasticities show that shifts in healthy or unhealthy food prices enter consumers' choices to substitute the relatively expensive food category for the cheaper one. According to estimated cross-elasticities, they are statistically significant and have well-defined sizes at the estimation point.

However, we cannot make direct inferences regarding substitution effects on patterns of prevalent obesity in Italy because, as reported in Section 2, the prices of both unhealthy and healthy foods rose in the sample period. The asymmetric responses of the elasticities evaluated at the sample means show that a 1 percent increase in the price of healthy food increases the budget share in the unhealthy food category by 0.536 percent, whereas those of unhealthy food increases the budget share of healthy food by 0.402 percent. The implications are threefold. The net effects of changes in food consumption, given shifts in relative prices, indicate a slight but significant impact on the growth of unhealthy food consumption. According to the net estimated elasticity of substitution, we can also reproduce the data reported in Figures 2 and 4, in which it was shown that quicker changes in healthy versus unhealthy food prices increased relative unhealthy food consumption and its expenditure share. Lastly, because the unhealthy category is more energy-dense, the demonstrated increases in total calorie intake and overweight patterns are therefore partly modelled by the channel of convenient food purchases.

If our attention is concentrates on the dynamics of the elasticities of substitution, η_{21} and η_{12}, an interesting implication of the non-stationarity of prices is that elasticities may change over time. This point is illustrated in Figure 6. The cross-price elasticity for healthy foods, η_{12}, appears to change little over time, whereas a slight recent increase in η_{21} is recorded since the end of 2001, when unhealthy foods became a much larger share of total spending (see Figure 4) or a rising category of food expressed in terms of relative quantities (see Figure 2).

The robustness of our estimates are shown by moving bread, pasta and olive oil into the sectors of healthy (specification I) and unhealthy foods (specification II), respectively. The implicit price index was then used to estimate the parameters of the long-run demand system. Appendix D shows the results. Although these food categories are very demanding in terms of quantity, in views of their cultural importance in the Mediterranean area, estimations are close to the benchmark model. Note that, in both estimations, the dimension of the cross-price elasticities are reduced, revealing how these "new aggregate" food categories show small changes in own or cross-prices. This result is in line with the low consumption responses to changes in prices of bread, pasta and olive oil founded by Conforti et al. (2001).

Clearly, however, the impact of changes in relative prices is not equally spread over the individuals of a society. As described above, inequality may arise from a heterogeneous consumption response among individuals as relative healthy and unhealthy food prices evolve. This implies a growing disparity in food access and calls into question the indiscriminate public health policies generally used to prevent obesity and overweight in Italy.

Table 5 lists the results of the long-run demand systems specified according to gender and socio-economic group, obtained by aggregating household expenditures for food and residual categories by RPI.

Focusing attention on the dimension of elasticities of substitution, η_{21k} and η_{12k}, only for individuals who belong to the high education group, we find that the cointegration rank is not exactly identified. Appendix C, reports the results of the cointegration tests analytically reported for each sub-group. It should also be noted that the computed elasticities of substitution for people above the relative poverty threshold and for the younger age class (age less 35) appear to be not statistically significant.

Besides these exceptions, our estimations contain several points of interest. First, cross-price elasticities for female household heads (0.08) indicate that the (net) effect of substitution of unhealthy foods for healthy foods, given changes in relative prices, is less than half in basis points of the effect for male household heads (0.18). This result partly emerges in Figure 3 (first panel), in which the sharp increase in relative healthy versus unhealthy food prices for female household heads does not affect changes in consumption or total calorie intake. Women are therefore able to keep their previous habits of eating. These findings are confirmed by the annual report of the multipurpose survey, ISTAT (2007), in which an increase in body weight is mainly found in males[13].

[13] The results of a greater propensity towards healthy food purchases find indirect confirmation by the greater control of women's weight with respect to those recorded for men, and by the low perception of being underweight of Italian women. These findings are in line with those obtained in France (Etile, 2007; de Saint Pol, 2009).

Second, the non-exact identification of the long-run demand system for the group comprising the behaviours of highly educated individuals does not allow us to discuss the net substitution effects of changes in healthy over unhealthy food prices. However, those who achieved, at maximum, the level of secondary school are significant and of the expected (higher) dimensional effect with respect to the average population. Individuals in this subgroup respond to changes in the rise of relative prices of healthy foods by consuming more unhealthy foods, 10 basis points more sensitive with respect to the whole sample estimated at sample average.

Thus, in line with the results of Ball and Crawford's review (2005), less educated Italians tend to prefer eating unhealthy foods, with higher calorie intake. The estimated elasticity of substitution also supports the remarkable rate of growth of the share of overweight and obesity prevalence for less educated individuals (2.8) with respect to those with higher levels of education (0.8)[14].

Third, among individuals grouped according to age, the insignificant cross-price responses of the youngest adults were in line with data shown in Figure 3, in which individual responses of quantities are very little sensitive to changes in relative food prices. As an approximate explanation, this age group spends least on healthy foods such as consumers tend to substitute this category of food less.

Bearing in mind that healthy food prices rise more than those of unhealthy ones, the other age groups were found to make a key contribution to the (net) elasticity of substitution by favouring the consumption of unhealthy foods. Higher elasticities of substitution for middle and advanced ages with respect to those estimated for the general population, leads to the fact that this group contributes more to body weight increases. This is particularly interesting from a policy perspective, because it suggests that the food consumption bundle may partly change with age, if consumers believe that their substitution costs less though in line with Baum (2007) age changes are not projected to increase obesity substantially.

Fourth, the substitution effects towards unhealthy foods which interest people living below the relative poverty threshold are, remarkably, 6 basis points higher than that of the mean of the population. As found in other countries by Komlos and Baur (2004), the relative price mechanism for adults of lower economic status passes to specific allocation of their disposable income on food that covers calorie requirements more easily by substituting low-calorie products with cheaper high-

[14] It is worth noting that these results obtained for the sample of less educated individuals, were computed by averaging the growth of each group from 1999/2000 to 2005.

calorie foods.

These results have a direct consequence on policy-makers' actions. As advocated by some, taxing unhealthy foods to arrest the dynamics of the prevalence of obesity may not only be of limited impact due to low price responses but may also increase the inequality of groups with lower economic or social status. Chouinard et al. (2007) show that a ''fat tax" instrument may be extremely regressive at lower income levels. Thus, the cost of a percentage increase in price by buying an extra unit of unhealthy food is borne to a greater extent by lower income groups (and lower educated groups).

Conversely, subsidies for healthy foods are likely to be more successful, especially for the disadvantaged subpopulation (Lin and Guthrie 2007)[15]. That is, because more indigent groups react more to changes in relative food prices, by targeting subsidies for each Euro spent on healthy foods, participants may not only have monetary benefits but also, as an externality, make an improvement by reducing the prevalence of overweight.

In addition, estimations suggest that consumption income responses for individuals below the relative poverty line is higher for healthy foods by 10 basis points than otherwise assessed so far for representative individuals. Instead, we can project a demand response for unhealthy foods, other foods and non-durable goods which is 4 and 5 basis points lower, respectively, than previously estimated in the full sample model.

These results are crucial when we attempt to address complementary health and fiscal policies. let us consider the introduction, with the Budget Act for 2009, approved at the end of 2008 by the Italian Parliament, of the ''social card" for indigent people aged over 65 and poor families with children under the age of three. Although the benefits for health are constrained by the limited numbers of potential beneficiaries and the dimension of the subsidy of the program which, in intent, follows that of the US food stamps, this welfare program could be used as a strategy to trigger policies to prevent the rise in the prevalence of overweight and obesity. This social program, which provides 40 euro a month and potentially involves less than the 20 percent of indigent people, may be extended, with very small administrative costs, to individuals below the poverty line, by targeting subsidies for purchasing healthy

[15] An implicit reason exists for the ineffectiveness of unhealthy taxation of energy-dense foods, determined by market competition in developed countries. As discussed by Powell and Chaloupka (2009), the presence of large quasi-competitive non-taxed high-calorie foods sold by groceries can potentially substitute taxed foods, making the impact on individual or aggregate body weight limited or irrelevant.

foods within the social card program - as proposed, for example, by the pilot project of the state of California (Guthrie et al., 2007)[16].

6 Conclusions

Cross-price elasticities estimated by demand systems provide a consistent framework to evaluate substitution effects among goods. In this paper, we present estimates of the long-run substitution effects of the categories of unhealthy and healthy foods for Italy.

By showing the close pattern linking unhealthy foods with more energy-dense foods, our findings suggest that the largest rise in healthy food prices, versus those of unhealthy ones, have favoured the consumption of high-calorie foods. This result matches the increase in body weight recently recorded in Italy.

In addition, to deal with the heterogeneous impact of food price changes on food expenditure allocation, we extracted the advantages of using data surveys as repeated cross-sections of time-series to assess the uneven substitution effects in demographic and socio-economic groups. One peculiarity of our results is that, as relative healthy food prices rise, individuals who are male, below the relative poverty threshold, and those with lower education tend to substitute energy-light foods for more energy-dense.

To sum up, the changes in relative prices, which have caused healthy foods to become 10 basis points more expensive in Italy in ten years, have generated a mechanism of substitution towards foods which are cheaper and high in calories, affecting mainly some disadvantaged groups. This calls into question the actions of policy-makers and the strategies that should be used to reduce such inequalities. If combating obesity is one objective of the government, then our results indicates that a lack of intervention regarding the loss of power to purchase by some segments of society may actually maintain the rate of overweight growing. Thus, among the tools of policy-makers, it would seem potentially more efficient to use subsidies in favour of the purchase of healthy foods, associated with an active welfare program or more specific food programs.

An open question is to reconcile our estimated cross-price elasticities, according

[16]The Italian Minister for Economy and Finance stated that in 2007, there were about 1,300,000 indigent people which were eligible for the ''social card", and over 7,400,000 poor individuals were estimated to be below the poverty line in the same year. A first ex-post evaluation reports that only 42 percent of indigent people complied with the ''social card" program.

to age group, with findings in the recent literature that changes in body weight are also a significant problem for younger generations. In this paper, age groups served to distinguish the food behaviour of individuals of different ages, using cohorts may be an appealing alternative for future research, when we want to assess intergenerational shifts in consumption. By studying for the changes in demographic structures and tastes, cohort characterisation may provide a better means of determining whether changes in relative prices further explain the recent overweight trends in Italy, or whether they will continue in the near future.

References

Allen RGD (1938) Mathematical analysis for economists Macmillan London

Attfield C (1997) Estimating a cointegrated demand system European Economic Review 41:61–73

Attfield C (2004) Stochastic trends demographics and demand systems University of Bristol mimeographs

Auld MC, Powell LM (2009) Economics of food energy density and adolescent body weight Economica forthcoming

Ball K, Crawford D (2005) Socioeconomic status and weight change in adults: a review Social Science & Medicine 60:1987–2010

Baum CL (2007) The effects of race ethnicity and age on obesity Journal of Population Economics 20:687–705

Barnett WA, Serletis A (2008) Consumer preferences and demand systems Journal of Econometrics 147:210–224

Bleich S, Cutler D, Murray C, Adams A (2008) Why Is the Developed World Obese? Annual Review of Public Health 29:273–95

Bonnet C, Dubois P, Orozco V (2008) Food consumption and obesity in France Paper presented at INRA-IDEI Workshop on Economics of Obesity Toulouse 12–13 December

Bray GA (1987) Overweight is risking fate: Definition classification prevalence and risks Annals of the New York Academy of Sciences 499:14–28

Burkhauser RV, Cawley J 2008 Beyond BMI: The value of more accurate measures of fatness and obesity in social science research Journal of Health Economics 27(2) 519–529

Chesher A (1998) Individual demands from household aggregates: Time and Age

variation in the composition of diet Journal of Applied Econometrics 13:505–524
Chou S, Grossman M, Saffer H 2004 'An economic analysis of adult obesity: results from the behavioral risk factor surveillance system' Journal of Health Economics 23:565–587
Chouinard H, Davis D, LaFrance J, Perloff J (2007) Fat Taxes: Big Money for Small Change Forum for health Economics and Policy 10:2
Commission of the European Communities (2005) Green Paper "Promoting healthy diets and physical activity: a European dimension for the prevention of overweight obesity and chronic diseases" COM (2005) 637
Costa-Font J, Fabbri D, Gil J, (2009) Decomposing body mass index gaps between Mediterranean countries: A counterfactual quantile regression analysis Economics and Human Biology 7:351–365
Coondoo D, Majumder A, Ray R (2004) Method of Calculating Regional Consumer Price Differentials with Illustrative Evidence from India Review of Income and Wealth 50(1):51–68
Conforti P, Pierani P, Rizzi PL (2001) Food and Nutrient Demands in Italy Actual Behaviour and Forecast through a Multistage Quadratic System with Heterogeneous Preferences Working Paper 303 Univerity of Siena
Cranfield JAL, Pellow S (2004) The role of global vs local negativity in functional form selection: an application to Canadian consumer demands Economic Modelling 21:345–360
Cutler DM, Glaeser EL, Shapiro JM (2003) Why Have Americans Become More Obese? Journal of Economic Perspectives 17(3):93–118
Daviglus ML, Liu K, Yan LL, Pirzada A, Manheim L, Manning W (2004) Relation of body mass index in young adulthood and middle age to Medicare expenditures in older age JAMA 292(22):2743–2749
Deaton AS, Muellbauer J (1980) An Almost Ideal Demand System American Economic Review 70:312–326
Deaton A (1986) Demand Analysis In Z Griliches and MD Intriligator (eds) Handbook of Econometrics 3:1767–1839
Dennis BH, Pajak A, Pardo B, Davis CE, Williams O, D Piotrowski W (2000) Weight gain and its correlates in Poland between 1983 and 1993 International Journal of Obesity 24:1507-1513
de Saint Pol T (2009) Evolution of obesity by social status in France 19812003

Economics and Human Biology 7:398–404

Drewnowski A, Darmon N, Briend A (2004) Replacing fats and sweets with vegetables and fruits–a question of cost American Journal of Public Health 94(9):1555–1559

Drewnowski A, Specter S (2004) Poverty and obesity: the role of energy density and energy costs American Journal of Clinical Nutrition 79:6–16

Drewnowski A, Darmon N (2005) Food choices and diet costs: an economic analysis The American Society for Nutritional Sciences J Nutrition 135:900–904

Edgerton DL (1997) Weak separability and the estimation of elasticities in multistage demand systems American Journal of Agricultural Economics 79:62–79

Etile F (2007) Social norms ideal body weight and food attitudes Health Economics 16:945–966

Elliott G, Rothenberg TJ, Stock JH (1996) Efficient Tests for an Autoregressive Unit Root Econometrica 64:813–836

Finkelstein EA, Fiebelkorn IC, Wang G (2003) 'National medical spending attributable to overweight and obesity: how much and who's paying?' Health Affair 219–26

Frisch R (1959) A complete scheme for computing all direct and cross demand elasticities in a model with many sectors Econometrica 27:177–196

Gelbach J,B Klick J, Stratmann T (2007) Cheap donuts and expensive broccoli: relative price effects on body mass index Working Paper Florida State University

Gorman WM (1995) Separability and Aggregation in: C Blackorby and A Shorrocks Editors Collected Works of W M Gorman Vol 1 Clarendon Press Oxford (1995)

Guthrie J, Frazao E, Andrews M, Smallwood D (2007) Improving Food Choices-Can Food Stamps Do More? Amber Waves 5(2):22–28

Hardy R, Wadsworth M, Kuh D (2000) The influence of childhood weight and socioeconomic status on change in adult body mass index in a British national birth cohort International Journal of Obesity 24:725–734

Hicks JR (1936) A revision of demand theory Oxford University Press Oxford

Klesges R, Ward K, Ray A Cutter G, Wagenknecht L (1998) The prospective relationships between smoking and weight in a young biracial cohort: the coronary artery risk development in young adults study Journal of Consulting and Clinical Psychology 66:987–993

Komlos J, Baur M (2004) From the tallest to (one of) the fattest: the enigmatic fate of the American population in the 20th century Economics and Human Biology 2 (1):57–74
Kopelman PG (2000) Obesity as a medical problem Nature 404:635–43
Kwiatkowski D, Phillips PCB, Schmidt P, Yongcheol Shin Y (1992) Testing the null hypothesis of stationarity against the alternative of a unit root Journal of Econometrics 54:159–178
ISTAT (2007) Annual multipurpose survey of Households "Aspetti della vita quotidiana" Roma
Johansen S (1995) Identifying restrictions of linear equations with applications to simultaneous equations and cointegration Journal of Econometrics 69:111–132
Lakdawalla D, Philipson T (2009) The growth of obesity and technological change Economics and Human Biology 7:283–293
Lewbel A (1999) Consumer Demand Systems and Household Equivalence Scales in Pesaran MH and Wickens MR Handbook of Applied Econometrics 167–196 Blackwell Oxford
Lewbel A (1996) Aggregation without separability: a generalized composite commodity theorem American Economic Review 86:524–543
Lewbel A, Ng S (2005) Demand systems with nonstationary prices Review of Economics and Statistics 87:479–494
Lin B-H, JF Guthrie (2007) How Do Low-Income Households Respond to Food Prices? In Can Food Stamps Do More to Improve Food Choices? An Economic Perspective Economic Information Bulletin Number 29-5Washington DC: US Department of Agriculture Economic Research Service
Lütkepohl H (1991) Introduction to Multiple Time Series Analysis Springer-Verlag Berlin
Mazzocchi M, Brasili C, Sandri E (2008) Trends in dietary patterns and compliance with World Health Organization recommendations: a cross-country analysis Public Health Nutrition 11(5) 535–540
Muellbauer JN (1976) Community preferences and the representative consumer Econometrica 44:979–999
Ng S (1995) Testing for homogeneity in demand systems when the regressors are nonstationary Journal of Applied Econometrics 10:147–163
Pesaran MH (1997) The Role of Economic Theory in Modelling the Long-run

Economic Journal 107(440):178–91
Pesaran MH, ShinY Smith R J (2000) Structural analysis of vector error correction models with exogenous I(1) variables Journal of Econometrics 97:293–343
Pesaran M, Shin Y (2002) Long-run Structural Modelling Econometrics Reviews 21:49–87
Powell LM, Chaloupka FJ (2009) Food Prices and Obesity: Evidence and Policy Implications for Taxes and Subsidies The Milbank Quarterly 87(1):229–257
Saikkonen P, Lütkepohl H (2000) Trend adjustment prior to testing for the cointegrating rank of a VAR process Journal of Time Series Analysis 21:435–456
Schroeter C, Lusk J, Tyner W (2008) Determining the impact of food price and income changes on body weight Journal of Health Economics 27:45–68
Villar JG, Quintana-Domeque C (2009) Income and body mass index in Europe Economics and Human Biology 7(1):73–83
Zheng X, Zhen C (2008) Healthy food unhealthy food and obesity Economics Letters 100:300–303

Acknowledgements: We would like to thank the participants at the ''8th Annual International Conference on Health Economics, Management and Policy", 27-30 June, 2009, Athens, Greece, at the 50th ''Conference of the Societ Italiana degli Economisti", 22-24 October 2009, Rome at the ''5th Nordic Econometric Meeting", 29-31 October 2009, Lund, Sweden and at the ''24th Annual Conference of the European Society for Population Economics", 9-12 June, 2010 Essen, Germany.

APPENDIX A.

Table A.1: Food classification

Healthy foods	Unhealthy foods
Meat	**Ice creams and Sweets**
Beef (lean)	Sugar
Poultry	Jam, marmelade, chocolate
Other low-fat meats	Ice creams
Fish	**Meat and cold cuts**
Fresh or forzen fish	Cold cuts
Preserved dry smoked fish	Pork
Fresh or frozen shellfish and other seafoods	Other fatty meats
Other fish	
Vegetables and legumes	**Oils and fat**
Dry or tinned legumes	Vegetable oil (except olive oil)
Dry or tinned vegetables	Butter
	Lard
Fruit	
Fresh fruit	**Milk**
Tinned fruit	Whole milk
Preserved fruit	Other whole-milk derivatives

Notes: Monthly price data at national and regional levels are available from ISTAT.

APPENDIX B. Identification of long-run AI

The cointegration relationships in the VECM equation (6), subject to reduced rank restrictions on the $\Pi = \alpha\beta^{*'}$ matrix, are not identified. Following Pesaran and Shin (2002), the identification of the long-run parameters in $\beta^{*'}$ requires the imposition of r restrictions on each cointegrating vector, although a necessary and sufficient condition (order condition) for the identification is that the number of the identifying restrictions, k, should be at least equal to r^2.

In order to explain these fundamental identifying conditions in our demand system with three categories of goods, we note first that adding up reduces the rank to two, i.e. $r = (n-1)$. As a formal extension of the ECM vectors in equation (6), let us consider two non-identified cointegrating vectors made up of the variables w_{1t}, w_{2t}, lnP_{1t}, lnP_{2t}, lnP_{3t}, $ln(y_t/p_t)$ and the intercept. The associated parameters are:

$$\widetilde{\beta}'_U = \left\{ \begin{array}{ccccccc} \beta_{11} & \beta_{21} & \beta_{31} & \beta_{41} & \beta_{51} & \beta_{61} & \beta_{71} \\ \beta_{12} & \beta_{22} & \beta_{32} & \beta_{42} & \beta_{52} & \beta_{62} & \beta_{72} \end{array} \right\} \quad \text{(B.1)}$$

The exact identifying restrictions $r^2 = (n-1)^2 = 4$ assume a diagonal structure because theory suggests that budget shares responds mainly respond to own and cross-price changes and income impulses, but not to (endogenous) changes in other budget shares. Formally,

$$\left\{ \begin{array}{cc} \beta_{11} = -1, & \beta_{12} = 0 \\ \beta_{21} = 0, & \beta_{22} = -1 \end{array} \right\} \quad \text{(B.2)}$$

so that the cointegrating vectors may be written as:

$$\widetilde{\beta}' = \left\{ \begin{array}{ccccccc} -1 & 0 & \beta_{31} & \beta_{41} & \beta_{51} & \beta_{61} & \beta_{71} \\ 0 & -1 & \beta_{32} & \beta_{42} & \beta_{52} & \beta_{62} & \beta_{72} \end{array} \right\} \quad \text{(B.3)}$$

In order to test theoretical restrictions, long-run parameter restrictions should be included. As discussed in the text, the property of symmetry may be imposed as a cross-equation restriction, $\beta_{32} = \beta_{41}$. The cointegrating vectors thus assume the following structure:

$$\widetilde{\beta}'_S = \left\{ \begin{array}{ccccccc} -1 & 0 & \beta_{31} & * & \beta_{51} & \beta_{61} & \beta_{71} \\ 0 & -1 & \beta_{32} & \beta_{42} & \beta_{52} & \beta_{62} & \beta_{72} \end{array} \right\} \quad \text{(B.4)}$$

Estimations of cointegrating vectors subject to symmetry is tested by the LR statistic distributed as a χ^2 with one degree of freedom. This restriction is not rejected when the loglikelihoods of this restricted model is compared with the exact-identified model in equation (B.3), the differences are not significant.

Lastly, as suggested by the demand theory, we impose and test in the cointegration vectors the properties of both symmetry and homogeneity. In addition to the symmetry restriction, $\beta_{32} = \beta_{41}$, the restriction of homogeneity for each equation is added, that is, $(\beta_{31} + \beta_{32} = -\beta_{51})$ and $(\beta_{32} + \beta_{42} = -\beta_{52})$. Thus, the cointegration vector is given as:

$$\widetilde{\beta}'_{SH} = \left\{ \begin{array}{ccccccc} -1 & 0 & \beta_{31} & * & * & \beta_{61} & \beta_{71} \\ 0 & -1 & \beta_{32} & \beta_{42} & * & \beta_{62} & \beta_{72} \end{array} \right\} \tag{B.5}$$

As shown in the text, the LR statistic, distributed as a χ^2 with three degrees of freedom, is then used to test these joint theoretical restrictions.

APPENDIX C.

Table C.1: Johansen's Cointegration Rank Test Statistics for AI system applied to subsamples of Italian survey data

	Gender				Education				Income				
	Male		Female		High		Low		Above pov. thresh.		Below pov. thresh.		95% critical values
H_0	LR		LR		LR		LR		LR		LR		λ trace
r = 0	157.80	[0.000]	136.02	[0.000]	162.69	[0.000]	132.75	[0.000]	136.50	[0.000]	126.97	[0.000]	103.68
r = 1	90.89	[0.002]	75.98	[0.058]	106.86	[0.000]	77.64	[0.043]	76.37	[0.050]	76.20	[0.058]	76.81
r = 2	39.74	[0.488]	41.84	[0.386]	56.88	[0.026]	50.76	[0.095]	49.22	[0.126]	46.74	[0.193]	53.94
r = 3	19.40	[0.761]	21.48	[0.633]	27.62	[0.262]	25.72	[0.362]	27.42	[0.279]	21.91	[0.604]	35.07
r = 4	9.95	[0.649]	9.77	[0.667]	11.56	[0.497]	9.59	[0.683]	9.08	[0.730]	11.61	[0.492]	20.16
r = 5	3.44	[0.512]	3.68	[0.473]	2.27	[0.724]	3.86	[0.445]	3.81	[0.453]	4.21	[0.394]	9.14

	Age								
	<35		35-49		50-65		>65		95% critical values
H_0	LR		LR		LR		LR		λ trace
r = 0	127.01	[0.000]	134.93	[0.000]	167.21	[0.000]	131.27	[0.000]	103.68
r = 1	80.33	[0.025]	81.26	[0.021]	79.98	[0.027]	79.82	[0.028]	76.81
r = 2	45.46	[0.236]	49.86	[0.112]	51.58	[0.081]	45.43	[0.237]	53.94
r = 3	19.85	[0.735]	27.57	[0.264]	30.17	[0.158]	23.84	[0.477]	35.07
r = 4	8.21	[0.802]	12.42	[0.420]	10.16	[0.629]	9.94	[0.650]	20.16
r = 5	2.62	[0.659]	5.05	[0.288]	3.51	[0.500]	3.86	[0.444]	9.14

Notes: r is number of cointegrating relations. We report Johansen trace tests for evaluating the identification condition of AI model. These values are estimated by optimal length lags with restricted intercepts and no trends in six endogenous variables w_1, w_2, $logp_1$, $logp_2$, $logp_3$ and $log(Y/p)$. Seasonal centred dummies are included. P-values of tests are shown in square brackets.

APPENDIX D.

Table D.1: Long run estimated elasticities from the demand system (equation 7)

	Hicksian Price Elasticities			Income
	(1)	(2)	(3)	Elasticities
Specification A				
(1) Healthy foods + bread, pasta and olive oil	-0.547	0.262	0.285	1.091
	(.091)	(.072)	(.0123)	(.601)
(2) Unhealthy foods	0.331	-0.523	0.192	0.852
	(.082)	(.181)	(.085)	(1.502)
(3) Other foods and non-durables goods	0.242	0.086	-0.328	1.022
	(.131)	(.046)	(.179)	(.075)
Specification B				
(1) Healthy foods	-0.657	0.281	0.376	1.381
	(.099)	(.084)	(.018)	(.642)
(2) Unhealthy foods + bread, pasta and olive oil	0.327	-0.253	0.026	0.632
	(.144)	(.076)	(.051)	(.899)
(3) Other foods and non-durables goods	0.187	0.056	-0.243	1.101
	(.102)	(.031)	(.125)	(.080)

Notes: in the specification A, bread, pasta and olive oil is moved from other food and non-durables goods to healthy food while in the specification B, bread, pasta and olive oil is moved to unhealthy food. Standard errors obtained by bootstrap procedure are shown in round brackets.

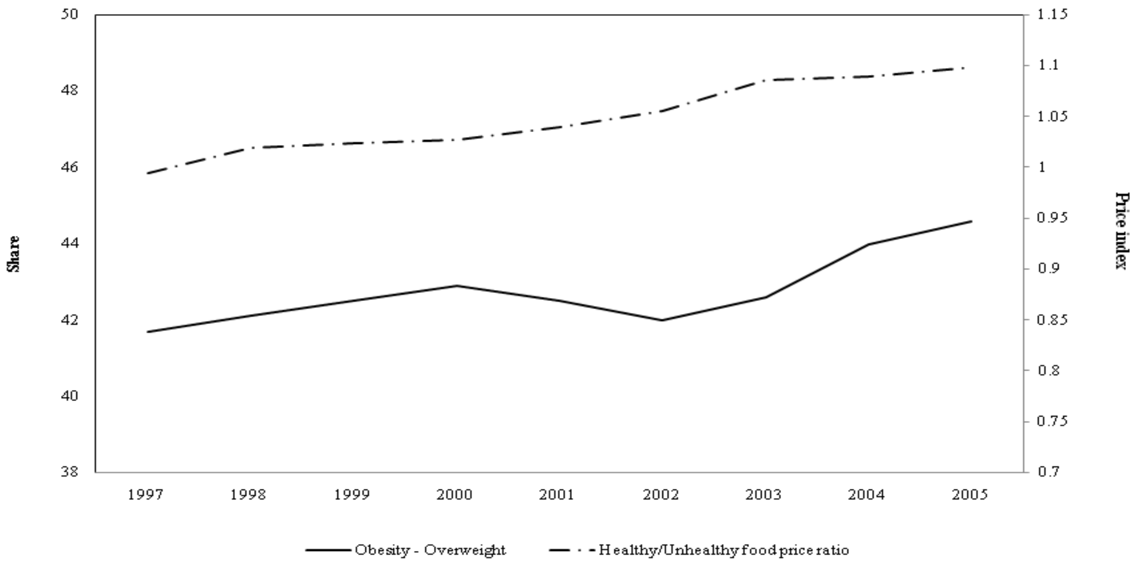


Figure 1: Share of obesity and overweight and price ratio patterns of healthy/unhealthy foods

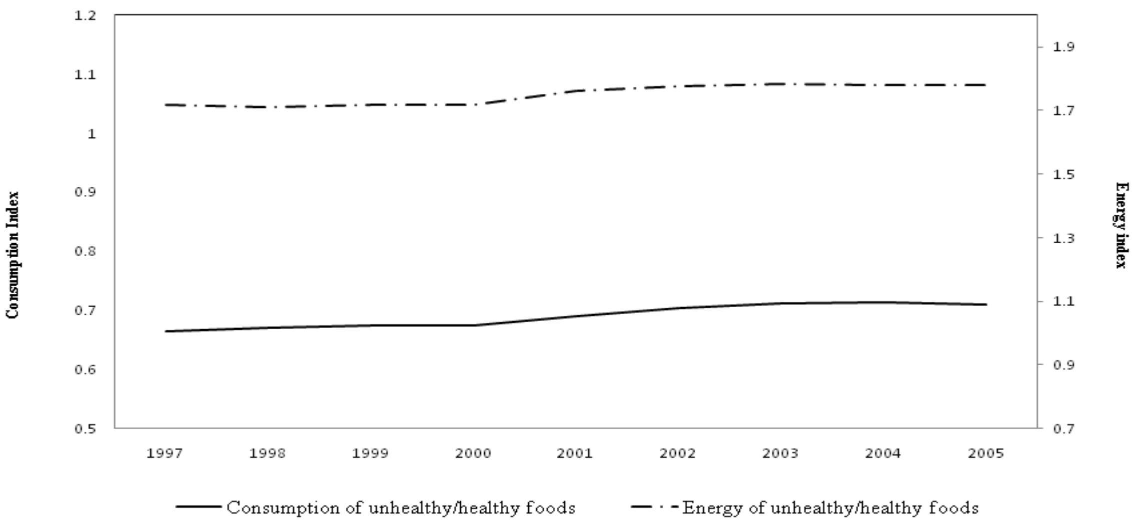
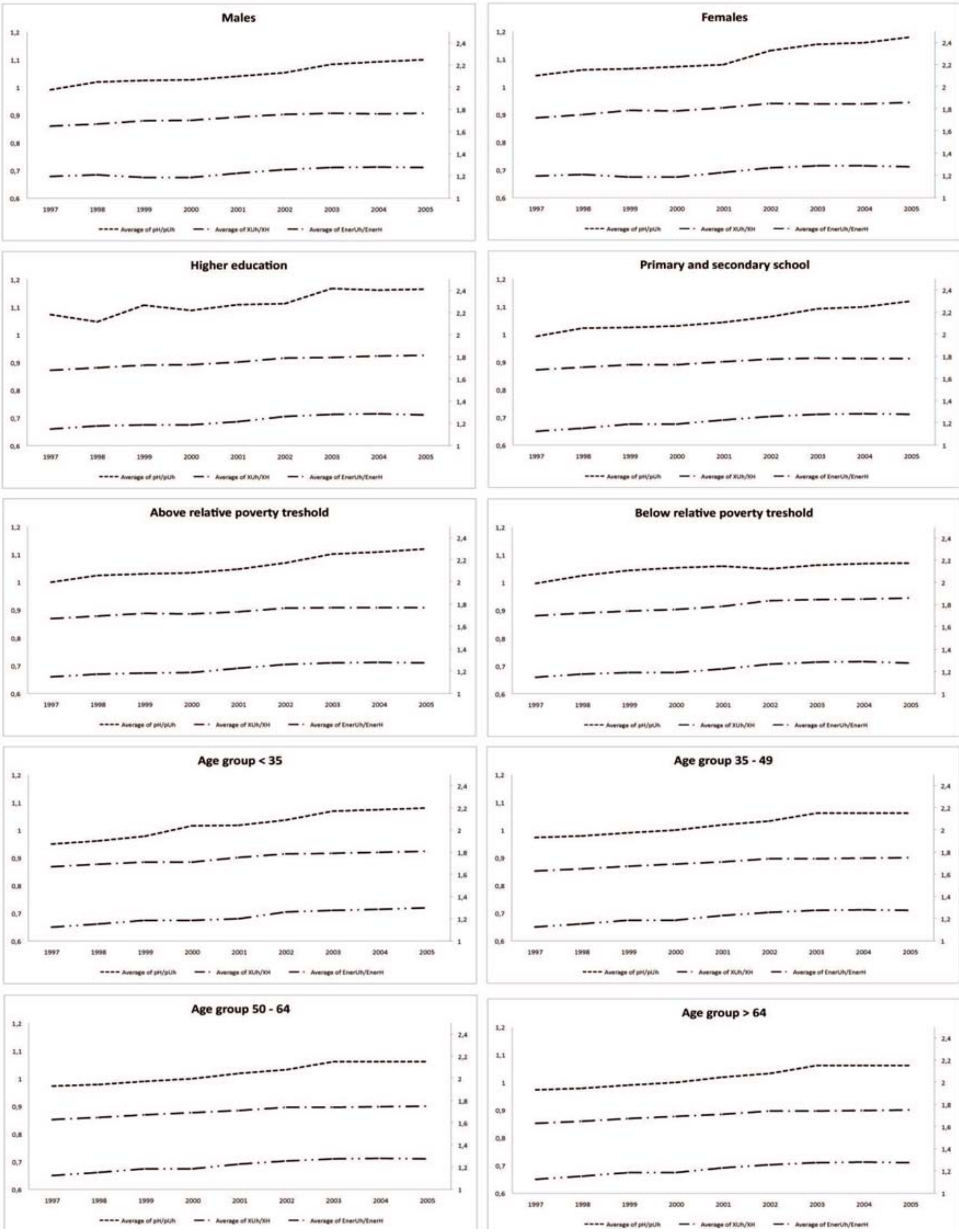
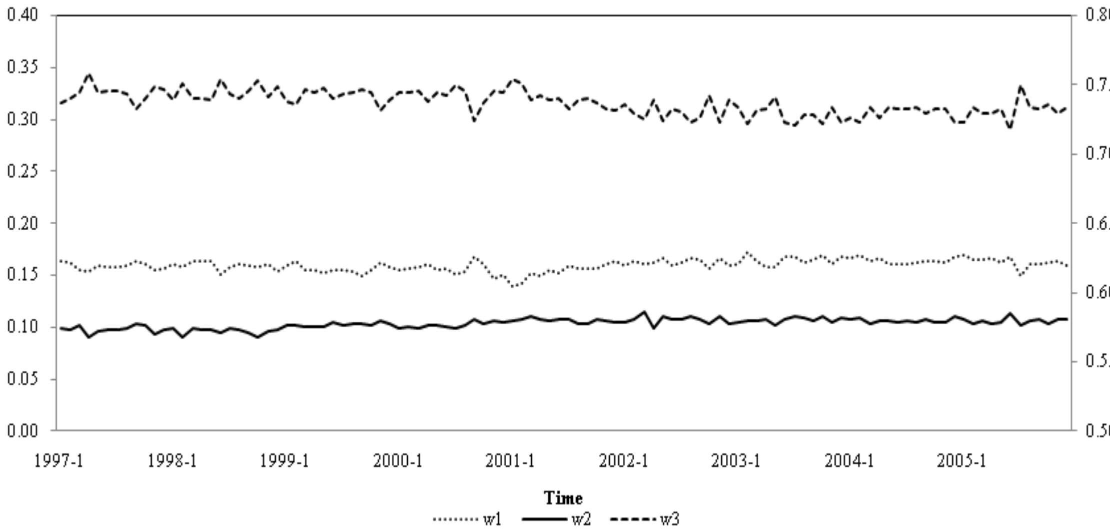


Figure 2: Ratio of unhealthy/helathy food consumption and relative energy index



Notes: Right scale: values of healthy/unhealthy food price ratio (p_h/p_{uh}) and unhealthy/healthy food consumption ratio (x_{uh}/x_h). Left scale: values of unhealthy/healthy food energy ratio ($ener_{uh}/ener_h$).

Figure 3: Patterns of healthy and unhealthy food prices, consumption and relative energy index



Notes: Right scale: budget share of other foods and non-durables (w_3). Left scale: values of budget shares of healthy (w_1) and unhealthy foods (w_2).

Figure 4: Budget shares (deseasonalised data, 1997 - 1 : 2005:12).

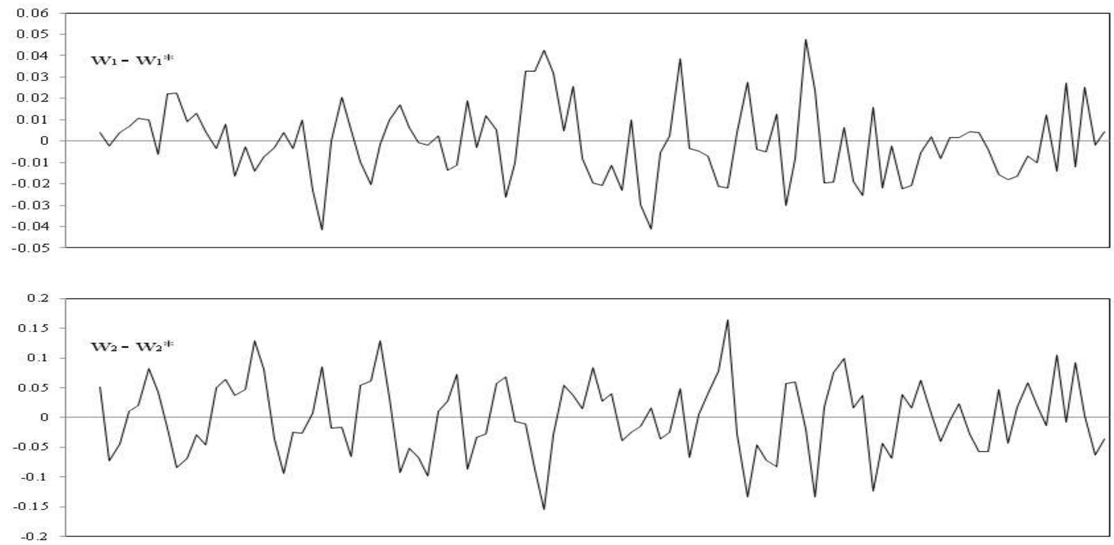


Figure 5: Deviations of observed expenditure shares from long-run equilibrium levels for healthy (w_1) and unhealthy foods (w_2)

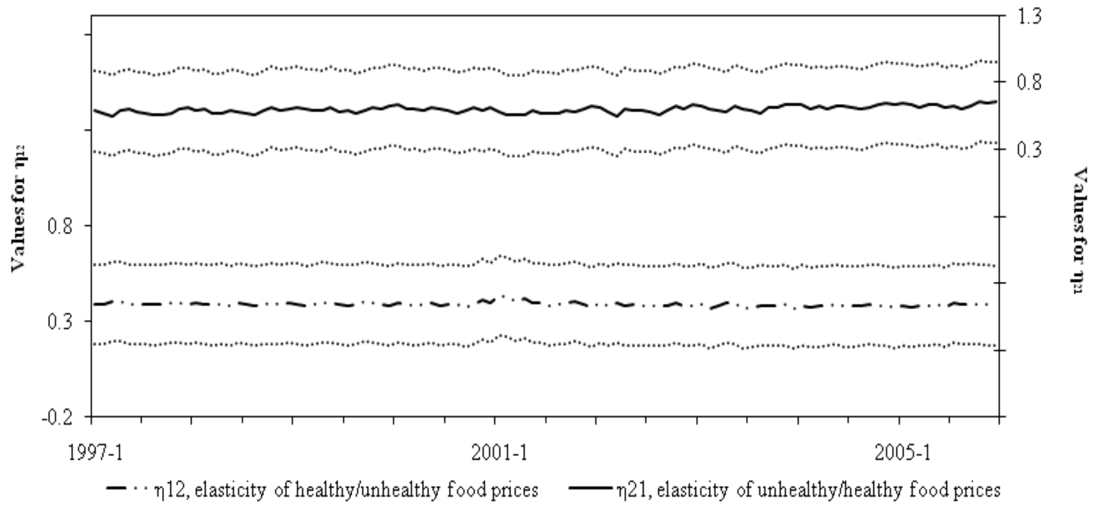


Figure 6: Long-run patterns of substitution elasticities of healthy and unhealthy food categories.

Table 1: Cointegration rank test statistics for AI system applied to Italian survey data with regional (1a) and national price indices (1b)

Specification 1a - Regional price index.

H_0	Johansen Trace Statistic Test		S&L Cointegration Test	
r = 0	137.19	[0.000]	108.43	[0.000]
r = 1	78.41	[0.037]	58.89	[0.061]
r = 2	50.95	[0.092]	34.99	[0.152]
r = 3	27.51	[0.267]	17.62	[0.277]
r = 4	9.97	[0.647]	5.03	[0.566]
r = 5	3.56	[0.493]	0.46	[0.556]

Specification 1b - National price index.

H_0	Johansen Trace Statistic Test		S&L Cointegration Test	
r = 0	144.68	[0.000]	92.14	[0.010]
r = 1	79.75	[0.028]	61.55	[0.036]
r = 2	38.40	[0.557]	33.36	[0.207]
r = 3	18.94	[0.787]	13.96	[0.547]
r = 4	9.96	[0.649]	6.16	[0.421]
r = 5	3.38	[0.523]	0.05	[0.867]

Notes: r is number of cointegrating relations. Johansen trace tests and Saikkonen & Lütkepohl tests are reported for identification of AI model. These values are estimated with a $VAR(3)$ with restricted intercepts and no trends in six endogenous variables w_1, w_2, $logp_1$, $logp_2$, $logp_3$ and $log(Y/p)$. Seasonal centred dummies are included. P-values of the are shown in square brackets.

Table 2: Estimated cointegrating vectors with price indices at regional (2a) and national (2b) levels and theoretical restrictions imposed

Specification 2a - Regional price index.

	w_1	w_2	$logp_1$	$logp_2$	$logp_3$	Income	Intercept
Vector of cointegration (1)	-1	0	0.0106	0.0452	-0.0558	0.0946	0.1456
			(0.0022)	(0.0171)	(0.0155)	(0.0454)	(0.0945)
			[4.938]	[2.644]	[-3.594]	[2.081]	[1.526]
Vector of cointegration (2)	0	-1	0.0452	0.0368	-0.0820	-0.0237	0.0837
			(0.0171)	(0.0061)	(0.0169)	(0.0561)	(0.0629)
			[2.644]	[5.954]	[-4.829]	[-0.422]	[1.331]

Theoretical restrictions:	LR test	d.f.	p-value
Symmetry	1.3972	(1)	[0.2371]
Symmetry and homogeneity	10.39	(3)	[0.015]

Specification 2b - National price index.

	w_1	w_2	$logp_1$	$logp_2$	$logp_3$	Income	Intercept
Vector of cointegration (1)	-1	0	0.0118	0.0503	-0.0621	0.0493	0.1245
			(0.0024)	(0.0108)	(0.0139)	(0.0295)	(0.1045)
			[4.9089]	[4.6567]	[-4.4667]	[1.6712]	[1.1919]
Vector of cointegration (2)	0	-1	0.0503	0.0322	-0.0825	-0.0124	0.0237
			(0.0108)	(0.0056)	(0.0630)	(0.0137)	(0.0562)
			[4.6567]	[5.6929]	[-1.4732]	[-0.9028]	[0.4221]

Theoretical restrictions:	LR test	d.f.	p-value
Symmetry	6.46	(1)	[0.011]
Symmetry and homogeneity	33.754	(3)	[0.000]

Notes: Standard errors in round brackets; student's t-test in square brackets. Degrees of freedom and p-values of LR tests in round and square brackets, respectively.

Table 3: Residual serial correlation of the vector error correction model in equation (7)

	VEC Residual Portmanteau Tests					VEC Residual Serial Correlation LM Tests	
Lags	Q-Stat	Prob.	Adj Q-Stat	Prob.	df	LM-Stat	Prob
1	7.6456	-	7.7191	-	-	46.424	0.1144
2	20.253	-	20.571	-	-	42.455	0.2127
3	42.478	0.2120	43.451	0.1837	36	37.933	0.3812

Note: Q-statistics and adjusted Q-statistics are shown with small sample correction for residual serial correlation up to specified order h (see Lütkepohl, 1991, for details). Also shown: LM test statistics for residual serial correlation up to specified order (see Johansen, 1995, for details). Under null hypothesis of no serial correlation, statistic tests are χ^2 distributed. Degrees of freedom are $k^2(p-h)$ and k^2, respectively, where k is number of endogenous variables of VAR and p is VAR lag order.

Table 4: Long-run estimated elasticities of demand system (equation 7)

	Hicksian Price Elasticities			Income Elasticities
	(1)	(2)	(3)	
(1) Healthy foods	-0.774	0.402	0.371	1.591
	(.134)	(.106)	(.092)	(.801)
	[-5.776]	[3.793]	[4.032]	[1.988]
(2) Unhealthy foods	0.5361	-0.573	0.037	0.802
	(.154)	(.201)	(0.015)	(1.428)
	[3.481]	[-2.855]	[2.466]	[0.560]
(3) Other foods and non-durable goods	0.082	0.006	-0.088	0.902
	(.021)	(.0036)	(.029)	(.824)
	[3.904]	[1.668]	[-3.034]	[1.094]

Notes: Standard errors obtained by bootstrap procedure in round brackets; student's t-test in square brackets.

Table 5: Identification, estimates and residual diagnostics of the VECM and elasticity of substitution by sub-samples

VARIABLES		VAR SPECIFICATION		THEORETICAL RESTRICTIONS		VECM DIAGNOSTIC	ELASTICITY OF SUBSTITUTION		
		Lag order	Rank of cointegration	Symmetry	Homogeneity and symmetry	Residuals serial correlation	η_{21}	η_{12}	Net effect $\Delta(\eta_{21}-\eta_{12})$
Gender	Male	3	2	3.484 (1) [.062]	10.84 (3) [.012]	39.66[.309]	0.533 (.063)	0.354 (.044)	0.178
	Female	3	2	10.723 (1) [.001]	3.517 (3) [.318]	39.280[.325]	0.222 (.021)	0.140 (.013)	0.081
Education	High	2	3				not significant		
	Low	3	2	0.237(1) [.625]	9.693 (3) [.021]	38.828[.343]	0.689 (.201)	0.456 (.127)	0.233
Relative poverty threshold	Above	3	2				not significant		
	Below	4	2	0.170 (1) [.679]	11.91 (3) [.007]	47.711[.092]	0.591(.189)	0.394 (.124)	0.196
Age	< 35	2	2				not significant		
	$35-49$	4	2	0.509 (1) [.475]	12.438 (3) [.006]	76.077[.348]	0.393 (.192)	0.267 (.130)	0.126
	$50-64$	3	2	1.878 (1) [.170]	2.947 (3) [.407]	38.043[.376]	0.543 (.064)	0.359 (.043)	0.184
	> 65	3	2	0.133 (1) [.714]	6.583 (3) [.086]	52.453[.037]	0.321 (.039)	0.201 (.024)	0.119

Note: High education stands for people that achieved a degree, master or PhD. Low education for the others. The criterium used to select the optimal lag order is the sequential modified LR test statistic. The test of the rank of cointegration of the unrestricted VAR use the Johansen's procedure. Cointegration test inferences of the AI model applied to each subsample are reported in Appendix. Degree of freedom and p-values of the LR tests for the theoretical restrictions are reported in round and square brackets, respectively. Diagnostic autocorrelation test and estimations of substitution elasticities are obtained by imposing symmetry and homogeneity in the VECM. Q-statistics distributed as a χ^2 is reported with the p-values in square brackets. The degree of freedom are 36 for VAR lag order of 3 and 72 for lag order of 4. In round brackets are reported bootstrap standard errors of the estiamated elasticities.