

2012 VAT Reform in Spain: The Food and Non-Alcoholic Beverages Case. Welfare Consequences*

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Abstract

This paper analyzes the effects that the 2012 VAT reform in Spain had on the households' welfare, focusing on one of the main expenditure groups: food and non-alcoholic beverages. To that end, households' demands are modeled as a QUAIDS, which is then estimated by means of a consistent two-step estimator introduced in Tauchmann (2010) and never before applied to studies of this type. This procedure allows to properly impose the traditional parameter restrictions that utility maximization requires on candidate demand functions in a context of a censored model. The results indicate that the higher the income level of a household, the lower the welfare loss of that household relative to its annual income. As a conclusion, the increment in the VAT tax rates that took place in Spain in 2012 could be considered as regressive. Additionally, the tax reform would have caused a lower social loss the higher the inequality aversion of Spanish society as a whole.

JEL classification: D12; D63; H25

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1 Introduction

A relevant issue in economic research is that of the likely effects of economic policies on welfare. In particular, in the case of tax and subsidy policies that affect economic agents' decisions, the aim of the researchers is to quantify the consequences that simulated or effective tax reforms might have exerted on households' demands and, ultimately, on welfare.

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Focusing on one particular strand of this literature as that on indirect taxation, the seminal papers along this line of research date back, at least, to King (1983a) and King (1983b), where the consequences, in terms of welfare, of the potential abolition of housing subsidies in England and Wales in 1973 were simulated. Since then, the economic literature has provided us with a list of works for different countries with a by now well established analysis procedure. In a first step, an as detailed as possible household utility generated demand system is estimated, thereby obtaining the corresponding price and income elasticities. In a second step, once one is able to reasonably predict households' responses to price changes induced by some given tax reform policy (most notably, changes in Value Added Tax, VAT, or in some subsidies) regarding demanded quantities, one is allowed to quantitatively assess the effects of the tax reform in terms of welfare.

Despite the above mentioned seminal papers were published more than thirty years ago, only a few empirical applications have been published since then. Starting for non-Spanish economies, Banks *et al.* (1997) simulate the effects of the imposition of a 17.5% sales tax on clothing as children's clothing and footwear were among the goods untaxed at such a tax rate under the ongoing VAT regime in the U.K. at that time. Urzúa (2001) evaluates the impact of two indirect-tax reforms (changes in VAT and excise taxes) that took place in Mexico in 1995 and 1998, analyzing their impact on welfare at the household and social levels. Salti and Chaaban (2010) study the impact of a rise in the VAT on poverty and inequality in Lebanon. Ramadan and Thomas (2011) estimate the negative welfare change measures of different alternatives suggested to eliminate subsidies on selected food items in Egypt. Janský (2014) simulates two effective changes in VAT legislation in the Czech Republic that took place in 2012 and 2013, plus a proposed change postponed until 2016.

Regarding the Spanish case, Labeaga and López (1994) estimate the welfare impact of the 1992 VAT reform, stressing the role played by demographics across households. Prieto-Rodríguez *et al.* (2005) simulate and evaluate three alternative (potential) cuts in the ongoing VAT rate on cultural goods at that time in terms of revenue and welfare, concluding that they would have led to welfare and efficiency gains that could be described as regressive.

In this paper, using data from the 2011 release of the Spanish Household Budget Survey, we study the VAT reform (an increase, in short) introduced in Spain in mid-year 2012 as a part of a major economic reform aimed at guaranteeing the budget stability and promoting economic competitiveness. In particular, we focus our study on one particular set of consumption goods (namely, food and non-alcoholic beverages) rather than the whole set of expenditure items that form the Spanish average household's consumption bundle, and which represented a 14.39% of the average household's total expenditure in that year. This choice obeys a threefold purpose: keep the magnitude of the econometric estimation process tractable but, also, aggregating

expenditure items reasonably homogeneous. Additionally, food and non-alcoholic beverages would represent what academic economists and a vast majority of society would consider as consumers' basic needs, thereby featuring a reduced room for substitution among expenditure items.

On the estimation side, two features present in this paper are worth mentioning. First, we adopt the Quadratic Almost Ideal Demand System (QUAIDS) model by Banks *et al.* (1997), which extends the Almost Ideal Demand System (AIDS) model introduced in Deaton and Muellbauer (1980) by including an additional quadratic term of the log-expenditure in the budget share equations. And, second, the treatment of zero observations. Despite the fact that we deal with aggregate items, the sample suffers from a serious problem, usual in consumption habit surveys, such as the non-negligible presence of zero expenditure observations.¹ This would not be a serious issue were it not for our need to do welfare analysis. This is so because one must first assure that the estimated demand system is utility generated and, consequently, the corresponding parameter restrictions hold. As in the AIDS model, negative semi-definiteness of the Slutsky substitution term matrix can be neither imposed nor tested.² And, also as in the AIDS model, homogeneity of degree zero and symmetry pose no problem. Adding-up, however, as is explained later in the paper, is hard to reconcile with a proper treatment of zero expenditure observations. Thus, the consistent estimator proposed by Shonkwiler and Yen (1999), which is probably the most extended alternative to estimate censored demand systems since its introduction and which represents an improvement over the Heien and Wesells (1990) inconsistent estimator, presents a patent flaw. While it allows to deal with the homogeneity and the symmetry restrictions, it fails doing so with the adding-up restriction as explained in Drichoutis *et al.* (2008). In an attempt to overcome this shortcoming, in this paper we implement the two-step estimator introduced in Tauchmann (2010) which, to the best of our knowledge, has not been previously applied to the estimation of a demand system. The main advantage of this estimator is that, while dealing with censored data, in addition to being consistent, it allows one to properly impose the adding-up restriction in an easy way.

Once the QUAIDS model is properly estimated, obtained price and expenditure elasticities are used for predicting the expected reaction of the Spanish households to the 2012 VAT reform. After this, welfare effects can be easily assessed. Regarding individual effects, the equivalent variation and the compensating variation are used to know to what extent households have been affected by the tax increase. Meanwhile, to evaluate social welfare effects, a measure

¹As it will be detailed below, the criterion followed to obtain the aggregates is a common VAT rate both before and after the reform.

²At most, one could always check, for instance, the sign of the eigenvalues of the estimated Slutsky substitution term matrices for each of individual in the sample and obtain, say, the proportion of negative semi-definite matrix cases, or check the sign of the eigenvalues of the Hicksian term matrix evaluated at some centered moment, usually the mean or the median.

such as the one proposed by King (1983a, 1983b) is computed. Efficiency of the reform is also studied by means of two alternative measures of the deadweight loss: one based on the equivalent variation, and the other one based on the compensating variation.

The paper is organized as follows. Section 2 introduces the consumer demand theoretical model. Section 3 deals with the estimation strategy. Section 4 describes the 2012 Spanish VAT reform and the data set. Section 5 shows the estimation results of the demand system. Section 6 is devoted to welfare analysis. Finally, Section 7 concludes. A formal Appendix contains some theoretical concepts used for the welfare analysis exercise.

2 The consumer demand model: the QUAIDS model

The consumer demand model that we adopt in this paper is the QUAIDS model introduced by Banks *et al.* (1997), who estimated a five commodity group demand system for the UK economy for a panel data set running from 1970 to 1986. The QUAIDS model can be considered as a generalization of Deaton and Muellbauer (1980) popular AIDS model as it includes the square of the logarithm of expenditure as an additional regressor. Therefore, any given good is allowed to be, say, a luxury at one level of expenditure, but a necessity at another, and Engel curves feature the maximum 3-rank condition that the theory predicts.³

Banks *et al.* (1997) start by assuming an indirect utility given by

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

where V denotes the indirect utility function, m denotes (nominal) expenditure, and \mathbf{p} an n -dimension price vector. Assuming further that $a(\mathbf{p})$, $b(\mathbf{p})$ and $\lambda(\mathbf{p})$ are flexible enough functions of \mathbf{p} such as

$$\ln a(\mathbf{p}) = \alpha_0 + \sum_{i=1}^n \alpha_i \ln p_i + \frac{1}{2} \sum_{i=1}^n \sum_{j=1}^n \gamma_{ij} \ln p_i \ln p_j, \quad (2)$$

$$b(\mathbf{p}) = \prod_{i=1}^n p_i^{\beta_i}, \quad (3)$$

and

$$\lambda(\mathbf{p}) = \sum_{i=1}^n \lambda_i \ln p_i, \quad (4)$$

³Since the QUAIDS model was introduced to the profession, it has been widely used in applied micro-econometric works and become a standard, the literature being plenty of references. Just to name some recent ones, see Abdulai and Aubert (2004), Dong *et al.* (2004), Kumar *et al.* (2005), Lambert *et al.* (2006), Mittal (2010), Zheng and Henneberry (2010), Davis *et al.* (2011) and Kumar *et al.* (2011).

where $\sum_{i=1}^n \lambda_i = 0$, applying Roy's identity yields the following demand system in terms of budget shares, w_i , after some algebra

$$w_i = \alpha_i + \sum_{j=1}^n \gamma_{ij} \ln p_j + \beta_i \ln \left\{ \frac{m}{a(\mathbf{p})} \right\} + \frac{\lambda_i}{b(\mathbf{p})} \left[\ln \left\{ \frac{m}{a(\mathbf{p})} \right\} \right]^2, \quad (5)$$

for $i = 1, 2, \dots, n$. Once again, mild but tedious algebra yields the expressions for expenditure, Marshallian and Hicksian price elasticities, E_i^m , $E_{i,j}^M$ and $E_{i,j}^H$ respectively as

$$E_i^m = 1 + \frac{1}{w_i} \left[\beta_i + \frac{2\lambda_i}{b(\mathbf{p})} \ln \left\{ \frac{m}{a(\mathbf{p})} \right\} \right], \quad (6)$$

$$E_{i,j}^M = -\delta_{ij} + \frac{1}{w_i} \left\{ \gamma_{ij} - \left[\beta_i + \frac{2\lambda_i}{b(\mathbf{p})} \ln \left\{ \frac{m}{a(\mathbf{p})} \right\} \right] \times \right. \\ \left. \left(\alpha_j + \sum_{k=1}^n \gamma_{jk} \ln p_k \right) - \frac{\beta_j \lambda_i}{b(\mathbf{p})} \left[\ln \left\{ \frac{m}{a(\mathbf{p})} \right\} \right]^2 \right\}, \quad (7)$$

and $E_{i,j}^H = E_{i,j}^M + E_i^m \times w_j$, for $i, j = 1, 2, \dots, n$, and where $\delta_{ij} = 1$ if $i = j$, and $\delta_{ij} = 0$ otherwise. Finally, it can be shown that the adding up, homogeneity and Slutsky term matrix symmetry restrictions that utility maximization imposes on candidate demand functions are given by the following set of conditions:

$$\sum_{i=1}^n \alpha_i = 1, \quad \sum_{i=1}^n \beta_i = 0, \quad \sum_{i=1}^n \gamma_{ij} = 0, \quad \sum_{i=1}^n \lambda_i = 0 \quad \text{and} \quad \gamma_{ij} = \gamma_{ji}$$

for all $i, j = 1, 2, \dots, n$.⁴

3 The estimation strategy

The model introduced in Eq. (5) is modified as some estimation issues must first be accounted for. First, the model admits the existence of additional regressors other than the (log of the) price vector and total expenditure which might potentially help to explain the household's consumption decision such as the demographics variables (in our case only age composition, for simplicity). The standard procedure followed is to include these additional regressors in an additive manner [see, among others, Banks *et al.* (1997)].

Second, as total expenditure is usually considered as a non-exogenous variable, introducing it in Eq. (5) could cause an endogeneity problem. Thus, and following the strategy proposed by

⁴As mentioned in the introduction, negative semi-definiteness of the Slutsky substitution term matrix can be neither imposed nor tested.

Blundell and Robin (1999), the QUAIDS model is augmented by introducing a correction term. To construct this new variable we regress households' log of total expenditure on demographics variables (age composition in our case), log of total income and log of prices. The residuals are then introduced in Eq. (5) as a new regressor.

Third, a usual difficulty when a researcher faces a demand system estimation is the zero quantity problem, that is, households that given market prices and total expenditure report a zero consumption of one or more goods and services. If that is the case, censored dependent variables pose a problem, and a model such as the one specified in Eq. (5) should not be estimated by the Ordinary Least Squares (OLS) estimator because this is biased and inconsistent. Thus, specially when zero observations represent a relevant sample proportion, as in the case here analyzed, they should be properly treated.

Heien and Wessells (1990) is one of the first studies that addressed this point in the context of the estimation of an equation system. Suppose that the following system of equations characterize the *latent* model

$$w_{ih}^* = f(\mathbf{x}_h, \boldsymbol{\theta}_i) + \varepsilon_{ih}, \quad (8)$$

$$d_{ih}^* = \mathbf{z}'_h \boldsymbol{\pi}_i + v_{ih}, \quad (9)$$

for $i = 1, 2, \dots, n$, $h = 1, 2, \dots, H$, and where, for the i -th equation (good) and h -th observation (household), w_{ih}^* and d_{ih}^* are the latent variables; \mathbf{x}_h and \mathbf{z}_h are vectors of exogenous variables for the h -th household; $\boldsymbol{\theta}_i$ and $\boldsymbol{\pi}_i$ are conformable vectors of parameters for good i ; and ε_{ih} and v_{ih} are random errors.

The *observed* counterparts, w_{ih} and d_{ih} , are given by:

$$d_{ih} = \begin{cases} 1 & \text{if } d_{ih}^* > 0 \\ 0 & \text{if } d_{ih}^* \leq 0 \end{cases}, \quad (10)$$

$$w_{ih} = d_{ih} w_{ih}^*. \quad (11)$$

This means that if a positive consumption of good i and for the h -th household is observed (i.e. $w_{ih} > 0$), then d_{ih} equals 1 and $w_{ih} = w_{ih}^*$; whereas if no consumption of good i and for the h -th household is observed (i.e. $w_{ih} = 0$), then d_{ih} equals 0.

The procedure designed by Heien and Wessells (1990) consists of two steps. In the first step, a probit regression that determines the probability that household h consumes good i [that is, system in Eq. (9)] is estimated. In a second step, the inverse Mills ratio for each household and each good, R_{ih} , is computed and introduced as an additional regressor in system in Eq. (8). R_{ih} is defined as $\phi(k_{ih} \mathbf{z}'_h \hat{\boldsymbol{\pi}}_i) / \Phi(k_{ih} \mathbf{z}'_h \hat{\boldsymbol{\pi}}_i)$, where vector $\hat{\boldsymbol{\pi}}_i$ denotes the maximum likelihood estimate of vector $\boldsymbol{\pi}_i$ obtained in the first step, $k_{ih} \equiv 2d_{ih} - 1$, and $\phi(\cdot)$ and $\Phi(\cdot)$

are the univariate standard normal probability density function and cumulative distribution function, respectively.

Shonkwiler and Yen (1999) find, however, that Heien and Wessells (1990) estimation procedure is not consistent. These authors assume that for each equation the error terms $[\varepsilon_{ih}, v_{ih}]'$ are distributed as a bivariate normal with $cov(\varepsilon_{ih}, v_{ih}) = \sigma_i$, and rewrite system in Eq. (8) in the following way:

$$w_{ih} = \Phi(\mathbf{z}'_h \boldsymbol{\pi}_i) f(\mathbf{x}_h, \boldsymbol{\theta}_i) + \sigma_i \phi(\mathbf{z}'_h \boldsymbol{\pi}_i) + \xi_{ih}, \quad (12)$$

where ξ_{ih} are the new errors. The system in Eq. (12) can be also estimated by a two-step procedure. The first step would be the same as in Heien and Wessells (1990) procedure: estimate the probit multiequational system in Eq. (9). In the second step, $\Phi(\mathbf{z}'_h \boldsymbol{\pi}_i)$ and $\phi(\mathbf{z}'_h \boldsymbol{\pi}_i)$ are replaced with $\Phi(\mathbf{z}'_h \hat{\boldsymbol{\pi}}_i)$ and $\phi(\mathbf{z}'_h \hat{\boldsymbol{\pi}}_i)$ and the system in Eq. (12) is estimated.

One of the implications of this procedure is that the adding-up condition cannot be imposed any longer via parametric restrictions [see, e.g. Drichoutis *et al.* (2008)], an often skipped issue [see, e.g. Zheng and Henneberry (2010) or Bakhshoodeh (2010)]. Note that if adding-up were imposed in Eq. (12), this condition would not hold in Eq. (8) because, in general, and using the QAIDS model as an example, $\sum_{i=1}^n \alpha_i \Phi(\mathbf{z}'_h \boldsymbol{\pi}_i) \neq 1$. This can result in a serious difficulty if one is interested in doing some kind of welfare analysis: if the demand system does not satisfy the properties that utility maximization imposes, consumer welfare calculations will *not* be valid [Hausman and Leonard (2005)].

An example of a recent attempt to address this issue can be found in Yen *et al.* (2003). In their paper the authors propose dropping the n -th good equation to avoid the singularity of the variance and covariance matrix of the perturbations, treating it as a residual category, and estimating the resulting $n - 1$ equation system along with the identity $\hat{w}_{nh} \equiv 1 - \sum_{i=1}^{n-1} \hat{w}_{ih}$. However, this procedure has an important drawback: the resulting estimates are *not* invariant to the dropped equation, and \hat{w}_{nh} could even be negative. Dong *et al.* (2004) implemented a mapping mechanism of the observed and latent variables that ensures: *i*) the adding-up condition, and *ii*) $w_{ih} \geq 0$ for i and h both in Eqs. (8) and (11), i.e. the latent and the observed shares. This procedure is a variation of the Amemiya-Tobin approach [Amemiya (1974)] for estimating censored systems. Despite the advantages that this procedure presents, it has been hardly used in empirical studies due to its high technical complexity.

As an alternative way to estimate a censored system of equations that allows one to impose the adding-up condition, in this paper we follow the procedure suggested in Tauchmann (2010). Based on the model described in Eqs. (8)-(11), the author suggests a two-step estimator that, instead of conditioning on only d_{ih} , conditions on the entire selection pattern, $d_h = [d_{1h}, \dots, d_{nh}]'$, thereby obtaining a consistent generalized Heckman-type estimator. Such an estimator is implemented as follows. First, a multivariate probit for Eq. (9) is estimated,

and its results are used to build the following correction terms:

$$M_{jh} = k_{ih}\phi(\mathbf{z}'_h\hat{\boldsymbol{\pi}}_i)\frac{\Phi^{n-1}(\tilde{A}_{jh},\tilde{R}_{jh})}{\Phi^n(\mathbf{z}'_h\hat{\boldsymbol{\pi}}_1,\dots,\mathbf{z}'_h\hat{\boldsymbol{\pi}}_n)},$$

for $j = 1, \dots, n$, where k_{ih} and $\phi(\mathbf{z}'_h\hat{\boldsymbol{\pi}}_i)$ are defined in the same way as above, while $\Phi^x(\cdot)$ denotes the cumulative density function on the x -variate standard normal distribution. Call Σ_{vv} the correlation matrix of the errors in Eq. (9) and s_{lj}^{vv} the corresponding (l, j) element, \tilde{A}_{jh} represents a vector of $n - 1$ elements $k_{lh}(\mathbf{z}'_h\hat{\boldsymbol{\pi}}_l - s_{lj}^{vv}\mathbf{z}'_h\hat{\boldsymbol{\pi}}_j) / (1 - (s_{lj}^{vv})^2)^{0.5}$, $l = 1, \dots, n$, $l \neq j$. \tilde{R}_{jh} is defined as $K_{jh}R_{jh}K_{jh}$, where R_{jh} is the partial conditional correlation matrix $Cor(v_h|v_{jh})$, and K_{jh} is a diagonal matrix with diagonal elements k_{lh} , $l \neq j$.

In the second stage, n new regressors (the M_{jh} correction terms) are incorporated to Eq. (8), giving rise to the following system

$$w_{ih} = d_{ih}f(\mathbf{x}_h, \boldsymbol{\theta}_i) + d_{ih}\sum_{j=1}^n\rho_{ij}M_{jh} + d_{ih}\tilde{\varepsilon}_{ih}, \quad (13)$$

where $\tilde{\varepsilon}_{ih} = \varepsilon_{ih} - E(\varepsilon_{ih}|d_h)$. Importantly enough, note that, first, d_{ih} serves as a weighting variable, i.e. censored observations are weighted by zero and are therefore excluded from the regression; this means that households reporting a zero consumption in *at least one* of the expenditure categories are dropped. And, second, the set of additional regressors, the M_j 's, are the *same* for all equations; this assures that the estimates will be invariant to the particular equation dropped.

Thus, using QUAIDS model in Eq. (5), the system in Eq. (13) can be rewritten as

$$w_{ih} = d_{ih}\left[\alpha_i + \sum_{j=1}^n\gamma_{ij}\ln p_{jh} + \beta_i\ln\left\{\frac{m_h}{a(\mathbf{p}_h)}\right\} + \frac{\lambda_i}{b(\mathbf{p}_h)}\left[\ln\left\{\frac{m_h}{a(\mathbf{p}_h)}\right\}\right]^2\right] + d_{ih}\sum_{j=1}^n\rho_{ij}M_{jh} + d_{ih}\tilde{\varepsilon}_{ih}. \quad (14)$$

As Tauchmann (2010) proves, Eq. (13) can be easily estimated by using OLS on an equation-by-equation basis. However, this does not allow to impose the parameter restrictions needed. Alternatively, one can estimate Eq. (13) as a Seemingly Unrelated Regressions (SUR) system, and this is the way followed in this paper. Moreover, for the particular case of the QUAIDS model in Eq. (14), as the equations are non-linear in the parameters, a non-linear estimation method must be applied.⁵

⁵With this aim the STATA[©] *nlsur* (for non-linear seemingly unrelated regressions) algorithm and option *ifgnls* are made use of. The *ifgnls* option estimates a system of equations by Iterated Feasible Generalized Non-Linear Least Squares, which converges to maximum likelihood.

4 The data set

The data have been obtained from the Spanish Household Budget Survey collected by the Spanish Statistical Office. Data correspond to 2011, the last whole year with the tax rates previous to the VAT reform.

The VAT was first introduced in Spain (in the whole country except Canary Islands, Ceuta and Melilla) replacing the old General Tax on Business Traffic on August 2nd in 1985 as a requirement for the official integration of Spain into the *European Economic Community (EEC)* on 1 January 1986. The tax initially contemplated a general 12% rate along with a reduced rate of 6% and an increased rate of 33%. The Spanish VAT legislation has experienced several reforms since that date, the last one (and the one that will be studied here) becoming effective on September 1, 2012. [See European Commission (2014).]

This tax law reform was a part of a major law reform designed as “(a) set of measures to ensure the budgetary stability and promoting competitiveness”. As a result, VAT rates were modified the following way: the general tax rate, which applies by default unless another specific rate is applied, was raised from 18% to 21%. The reduced tax rate, which mainly applies to some types of food and drinks, hotels, coffee shops and restaurants, transport of passengers and new house building among others, was raised from 8% to 10%. The so-called super reduced rate, which applies to basic necessities such as vegetables, milk, bread, fruit, pharmaceutical products and books, newspapers and the like, of 4% was not changed. Finally, some goods and services such as some cultural services, e.g. public shows, hairdressing services, funeral services, or recreational and sports services among others, taxed at an 8% rate before the reform, became taxed at a 21% rate afterwards. [See Real Decreto-ley 20/2012 in BOE (2012) for details].

The survey is made of three complementary files. The household file, with general information about the household such as region, municipality size, household size, household head features, main dwelling features, total expenditure or total income. The household member file, with information about the nationality, educational attainment, labor status or revenues on each household member. And, finally, the expenditure file, providing information on nominal expenditure and quantities purchased. Depending on the 4 levels of aggregation featured in the survey, this third file includes 12 broad groups of expenditure in the least disaggregated case and 225 in the most disaggregated one.

Given the computational demands of the estimation procedure, in this paper we focus on just one broad expenditure group, namely *food and non-alcoholic beverages*.⁶ And, in

⁶After reducing the number of expenditure items to 11 as will be explained below, estimation of the multivariate normal probit in the first step proved extraordinarily time consuming. And even though we only estimate the restricted QUAIDS model in the second step, it requires to estimate 275 parameters.

addition to keeping the magnitude of the problem tractable, concentrating on just one type of expenditures allows one to deal with reasonably homogeneous items. A by now well established procedure in the literature, it can be theoretically justified on the grounds of the so called two-stage budgeting: in the first stage, households (or individual consumers) would allocate total expenditure to different groups of expenditure, thereby obtaining optimal expenditure levels for each of the expenditure groups. And, in the second stage, expenditure on each of the groups would be optimally assigned to each of the individual items within each group. [See, among others, Menezes *et al.* (2008), Mittal (2010)]. The optimality of this budgeting procedure is guaranteed under precise forms of separability of preferences, the original result dating back to Gorman (1959). [See, e.g. Honohan and Neary (2003) n. 7, p. 200, and Blackorby *et al.* (1998) and references therein.]

The survey features two main drawbacks. First, it does not include prices, so that these must be imperfectly proxied by unit values (*i.e.* the expenditure-to-quantity ratios) not controlling, therefore, for likely quality differences. [See, e.g. McKelvey (2011), Lahatte *et al.* (1998), Crawford *et al.* (2003)]. Second, not all expenditure items provide quantities (*e.g.* salt, spices and culinary herbs, or pastry and cooked mass). This leaves us with a total number of usable expenditure items of 60 out of the 70 items originally included in the sub-group, representing a 92.09% of the total expenditure on food and non-alcoholic beverages on average across households. Finally, in order to ensure tractability of the estimation process, these items have been further reduced following two principles: *i*) homogeneity of the goods grouped, and *ii*) common VAT rates. The resulting 11 group expenditure items are shown in Table I. To obtain each of these items, a Stone-like price index has been computed first. Thus, for any group i which is made of, say, M individual goods, we set a price index $p_i \equiv \prod_{j=1}^M p_{i,j}^{w_{i,j}}$, where $w_{i,j} \geq 0$, and $\sum_{j=1}^M w_{i,j} = 1$ for all i , $w_{i,j}$ denoting the expenditure share of good j relative to total expenditure on such M goods, *i.e.* $w_{i,j} \equiv p_{i,j}q_{i,j} / \sum_{j=1}^M p_{i,j}q_{i,j}$. Similarly, for any such a group i , we obtain a quantity index given by $q_i \equiv \sum_{j=1}^M w_{i,j}q_{i,j}$.⁷ Finally, total expenditure on food and non-alcoholic beverages is obtained as $m \equiv \sum_{i=1}^{11} p_i q_i$, so that budget shares, w_i are simply obtained as $w_i = p_i q_i / m$, for $i = 1, 2, \dots, 11$.

[Insert Table I around here]

The sample contains 22,119 households. Each household is assigned a time and space scaling factor so that the households interviewed represent a total of 17,342,147 households.

⁷Trivially note that q_i *cannot* be defined as a geometric mean as, unlike the case of prices, quantities $q_{i,j}$ can be zero.

In the first stage of the estimation process, some observations were eliminated for different reasons. More precisely, 119 observations were dropped for displaying an attributed level of income equal to 0. Additionally, 1,227 observations were dropped as they corresponded to households living in Canary Islands, Ceuta or Melilla. Finally, 72 households whose total expenditure in the 11 expenditure categories here considered equalled 0 were eliminated. As a result, the sample was finally reduced to 20,701 households representing a total of 16,359,412 households. Main descriptive statistics, i.e., mean, median, standard deviation (SD), minimum and maximum of both dependent and independent variables are shown below.

Descriptive statistics associated to the dependent variables (i.e. budget shares) are reported in Table II. Means and medians vary significantly, from 0.028 and 0.009 (expenditure item 6) to 0.187 and 0.167 (expenditure item 3). The sum of the three most important expenditure items (items 3, 1 and 5) represents 50% of the total budget share, whereas the three least important ones (items 6, 7 and 10) only a 10%. Regarding SDs, observed differences are less, the smallest one being 0.044 and biggest one 0.119. Minima are always zero, representing households that do not consume any quantity of the corresponding expenditure item. As Table II shows, there is a substantial proportion of non-consuming families in all the expenditure categories, ranging between 1.560% in the case of expenditure item 1 and 44.278% in the case of expenditure item 6. Concerning the maxima, these are always 1 except for expenditure items 6 and 7.

[Insert Table II around here]

A similar analysis is carried out, but now referring to the independent variables, in Table III. Prices largely change from the cheapest products such as items 8, 11 or 5, with mean and median prices around 2 euros, to the most expensive ones (items 3 and 4), around 4 times bigger. In the same way, SDs vary significantly among different expenditure items, from 0.936 to 6.313, so means are not representative in some cases. Meanwhile, there are also important differences in minimum and maximum values, specially in the latter ones. Regarding the total expenditure, one can observe both a representative mean and a big difference between minimum and maximum values. Finally, there are some demographic variables that are included in the QUAIDS model. They represent the number of members composing a household by age intervals. More precisely, our dataset provides seven age groups: [0,4], [5,15], [16,24], [25,34], [35,64], [65,84] and 85 years or more, which in our notation will be denoted as mem_1 , mem_2 , mem_3 , mem_4 , mem_5 , mem_6 and mem_7 , respectively. Important differences among minimum and maximum values are not noticed, but means differ, in part due to the non homogeneous width of the intervals. The same happens with their SDs, although differences are not so remarkable.

[Insert Table III around here]

5 The estimation results

As mentioned before, the estimation strategy consists of two steps. In the first step Eq. (9) was estimated as a multivariate probit using *mvprobit*, a simulated maximum likelihood estimator included in STATA[©] [see Capellari and Jenkins (2003)]. The system consisted of 11 equations, one for each expenditure item, and the set of regressors, z_h , was defined as $z'_h = (const, mem_1, \dots, mem_7, \ln p_1, \dots, \ln p_{11}, \ln m, \ln^2 m)'$, where *const* is a constant term and $h = 1, \dots, 20,071$. As each observation has a particular weight in the sample, the estimation was done using a time and space scaling factor variable.⁸

Using the multivariate probit results the correction terms M_{jh} were constructed and, based on Eq. (14), in the second stage the following model was estimated:

$$w_{ih} = d_{ih} \left[\alpha_i + \sum_{j=1}^7 \tau_{ij} mem_{jh} + \sum_{j=1}^{11} \gamma_{ij} \ln p_{jh} + \beta_i \ln \left\{ \frac{m_h}{a(\mathbf{p}_h)} \right\} \right. \\ \left. + \frac{\lambda_i}{b(\mathbf{p}_h)} \left[\ln \left\{ \frac{m_h}{a(\mathbf{p}_h)} \right\} \right]^2 \right] + d_{ih} \delta_i e_h + d_{ih} \sum_{j=1}^{11} \rho_{ij} M_{jh} + d_{ih} \tilde{\varepsilon}_{ih}, \quad (15)$$

for $i = 1, \dots, 11$, and $h = 1, \dots, 6,064$, and where the mem_{jh} 's represent the demographic variables and τ_{ij} their associated coefficients, and e_h and δ_i are the term for correcting the expenditure endogeneity and its associated coefficients, respectively. Note that in this second stage, 14,637 additional households were dropped because those households reported a zero consumption in *at least one* of the 11 expenditure categories considered. As a result, the sample size for this second stage was reduced to 6,064 households representing a total of 5,137,446 households.

Adding-up, symmetry and homogeneity conditions were imposed, so that the last equation was dropped in order to avoid the singularity of the variance and covariance matrix of the perturbations.⁹ Parameters for this omitted equation were retrieved by using the restrictions imposed. A system of 10 equations and 275 parameters was then estimated taking into

⁸Results, although not shown in the paper for the sake of space saving, are available from the authors upon request.

⁹After computing the median of the Hicksian substitution term matrices for all the households in the sample, it turned out that the corresponding eleven eigenvalues were both real and negative. In other words, at least the matrix of medians of substitution terms was negative semi-definite.

account, as in the first step, the different weight of each observation. Among these parameters, 110 were statistically significant at 5%, and 130 at 10%. As the parameters have no direct economic interpretation, they are of little interest, so that they are not shown in the paper.¹⁰ However, these estimated coefficients are necessary to compute Marshallian own and cross price-elasticities and expenditure elasticities.

Elasticities were computed for each household. Again, all calculations were done taking into account the different weight of each household in the sample. As a large variability due to some extreme values is observed, the means of the elasticities are not representative in many cases. In order to avoid this, the medians of the elasticities are reported instead in Table IV.

Own price elasticities are always negative, ranging between -1.784 for item 7 and -0.315 for item 1. This means that all the items are considered ordinary goods. However, regarding cross price elasticities there is not a clear pattern. As both positive and negative signs are observed, expenditure items are substitutes or complements, respectively. Item 3 has the highest number of complementary relationships of all the goods, being complement of 9 items, while item 7 has the highest number of substitutability relationships, being substitute of 8 items. Looking at the expenditure elasticities, which are always positive with a range between 0.194 for item 1 and 2.014 for item 7, it is concluded that all the items represent normal goods. Elasticities lower (higher) than 1 in absolute value express that the expenditure items they refer to are necessary (luxury) goods.

[Insert Table IV around here]

6 Distributive and welfare impact of the tax reform

As mentioned in the introduction, the final aim of this paper is to quantify the distributive and welfare effects of the 2012 Spanish VAT reform and make comparisons across households. Once the price elasticities of our QUAIDS model have been conveniently estimated, the next step consists in predicting households' expected reaction after the tax reform. To this end, of course, we first need to assess the induced price changes. Here we follow the often made assumption that producer prices remain invariant to VAT changes or, in other words, we assume infinitely elastic supplies. [See, e.g. Labeaga and López (1994) or Prieto-Rodríguez *et al.* (2005)]. Thus, elements in the post-reform h -th household's price vector, \mathbf{p}_h^1 , are given by $p_{h,i}^1 = (1 + t_i^1)p_{h,i}^0 / (1 + t_i^0)$, where $p_{h,i}^0$ trivially denotes the i -th element of pre-reform h -th

¹⁰As before, results are available from the authors upon request.

household's price vector, \mathbf{p}_h^0 , while t_i^0 and t_i^1 denote the corresponding pre- and after-reform tax rates, respectively, for good i .

Putting together these two ingredients (after tax reform prices and price elasticities) the usually followed procedure in the literature requires next to predict the effects on tax proceeds and welfare, in other words, the "orthodox methodology" referred to in Urzúa (2001) and first settled in King (1983b). Note that the assumption of constant producer prices has direct consequences to obtain the tax revenues of h -th household before and after the tax reform, R_h^0 and R_h^1 , respectively. Thus, one trivially obtains that $R_h^0 = \sum_{i=1}^n t_i^0 p_{h,i}^0 q_{i,h}^0 / (1 + t_i^0)$ and $R_h^1 = \sum_{i=1}^n t_i^1 p_{h,i}^0 q_{i,h}^1 / (1 + t_i^0)$, where $q_{i,h}^0$ and $q_{i,h}^1$ denote the h -th household's quantity demanded of the i -th good at the old and the new price vector respectively. Assuming further as usual that the h -th household's total expenditure on food and non-alcoholic beverages were invariant to the VAT reform, one can easily compute $q_{i,h}^1$ as

$$q_{i,h}^1 = q_{i,h}^0 \left[1 + \sum_{j=1}^n E_{i,j}^M \frac{(t_j^1 - t_j^0)}{1 + t_j^0} \right], \quad (16)$$

for $i = 1, 2, \dots, n$, and where $E_{i,j}^M$ denotes the (Marshallian) demand cross price elasticity of good i with respect to price j , which was obtained in Eq. (7). Regarding welfare analysis, the reader is referred to the formal Appendix at the end of the paper for definitions and formal expressions. Our results follow next.

6.1 Welfare effects

Concerning the effects on tax revenues, the results are shown in Table V. According to our calculations, the VAT reform implied an estimated median increment in the household's tax bill of 56.1 euros from the initial (i.e. pre-reform) tax bill of 266.2 euros. [See Table V, columns 1 and 2]. And the VAT rate on goods and non-alcoholic beverages evaluated at their median values before and after the reform were 6.5% and 7.9%, respectively. [See Table V, columns 4 and 5] Regarding standard measures of individual welfare changes such as the equivalent or the compensating variations, Eqs. (A.1) and (A.3), they turn out to be pretty close to each other and display a negative sign as expected: since the reform studied here is not revenue neutral, *all* households lost after the increase in the VAT rates. More precisely, the estimated medians for the equivalent and compensating variations equal -57.6 euros and -58.5 euros respectively. [See Table V, columns 6 and 7.] As a graphical illustration, Figure 1.A shows the plot of the equivalent variation against the income. The corresponding plot for the compensating variation is quite similar and, therefore, omitted.

From an efficiency standpoint, however, more relevant measures of the welfare change induced by the reform should jointly consider the equivalent or the compensating variations

and the change in tax revenues, that is the deadweight losses. This way, previous results allow us to compute two alternative measures of the excess burden: one obtained after the equivalent variation, the other one based on the compensating variation. [See Eqs. (A.5) and (A.6).] Both magnitudes, of course, must take on similar values: the estimates of the median values are -125.6 euros and -126.7 euros respectively. [See Table V, columns 8 and 9.] Graphically, Figure 1.B shows the plot of the excess burden computed after the equivalent variation against the income. Once again, the plot of the excess burden computed after the compensating variation is also omitted. Finally, complementary to the above shown individual welfare change measures, we compute the initial and final equivalent expenditures, whose estimated median values equal 4429.2 euros and 4315.7 euros respectively. [See Eqs. (A.4) and (A.2) and Table V, columns 10 and 11.]

[Insert Table V around here]

[Insert Figures 1.A and 1.B around here]

Previous figures (medians) have been obtained as aggregate measures. From a distributional point of view, however, the relevant point is how those medians are split among households with different incomes. Table VIa provides us with the answer. Thus, and concerning the distribution of the increment in the tax bill, and as one would expect, this follows an increasing pattern across income levels. For instance, the median increment of the tax bill for the upper income decile equals almost 2 times the increment for the lower income decile. [See Table VIa, column 3]. Along these same lines, a natural question arising here is that of which income groups would suffer a higher welfare loss as a consequence of the reform. It turns out that an expected increasing pattern between welfare loss and income appears. For instance, households in the upper income decile experienced a welfare loss around 2.5 higher than households in the lower income decile. [See Table VIa, columns 4 and 5.] As for the two deadweight loss measures considered, these can be similarly broken down across households within different income deciles: just comparing the lower and the upper ones, the estimated median for the latter equals around 2 times that of the former. [See Table VIa, columns 6 and 7]. Last, columns 8 and 9 in Table VIa show how both equivalent expenditures are distributed across households' income distribution. Once again, comparing the upper and the lower income deciles, the median equivalent expenditure in the former income bracket is around 2.4 times the median equivalent expenditure in the latter.

[Insert Table Va around here]

Figures in Table VIa are expressed in euros, i. e. in absolute terms. A still open question, however, is how households in different income brackets have been affected by the VAT increase in food and non-alcoholic beverages, so that the measure can be termed as progressive or regressive. Or, more precisely, which pattern follows *the proportion of income* that, say, the change in the tax revenues represents over annual income for families with different levels of income. The reason is simple: higher levels of income are associated to higher levels of expenditure on food and non-alcoholic beverages, but this kind of expenditure represents itself a *higher* proportion of the annual income for lower income brackets. [See Table VIb, column 2]. Thus, while households in the first income decile spent a 36.72% of their annual income, households in the upper income decile spent a 10.58%. Looking at the change in tax proceeds, the pattern appears clear: the higher income level, the lower the proportion that the increment in tax revenues represents over total income. [See Table VIb, column 3]. Similar patterns are obtained for all the various welfare losses considered: lower levels of annual income are associated to higher (relative) welfare losses. [See Table VIb, columns 4-7]. Finally, given these results, it is no surprise that both the (relative) initial and final equivalent expenditures display a decreasing pattern relative to income. [See Table VIb, columns 8 and 9] The conclusion seems clear: the increment in the VAT tax rates that took place in Spain in 2012, and focusing on just one particular set of consumption goods such as food and non-alcoholic beverages, could be labeled as regressive.

[Insert Table VIb around here]

To finish up the welfare analysis, we next consider the *social welfare* effect implied by the tax reform. Results for King's proportional increase in equivalent income (λ) are shown for alternative values of the inequality aversion parameter (ε) in Table VII. The conclusion seems to patent: regardless of the inequality parameter, the VAT reform induced a loss in social welfare ($\lambda < 1$). Additionally, note the increasing pattern between ε and λ . This leads us to conclude the tax reform at issue would have caused a lower social loss the higher the inequality aversion of Spanish society as a whole.

[Insert Table VII around here]

7 Conclusions

With the aim to get the budgetary stability, and along with other measures, Spanish Government decided to increase VAT rates in 2012. Obviously, this decision has had effects on both households and social welfare. In order to assess to what extent families and society in general have been affected by this decision, we estimate a QUAIDS model focusing on one of the main expenditure groups (food and non-alcoholic beverages), consisting of 11 expenditure items. This type of studies are of great public interest, specially for countries such as Spain, where the current economic crisis has beaten its economy in a particular hard way, triggering an spectacular increase in the unemployment rate (around 26% in 2012) and in the population at risk of poverty (more than 20% in 2012, according to Living Conditions Survey made by the Spanish Statistical Office).

Estimating a demand system when there is a non-negligible presence of zero expenditure observations in the sample used, as is the case here analyzed, is not a trivial task. If one is interested in using the estimation results to carry out a welfare analysis, it is necessary not only to use a estimator with good statistical properties but also to properly introduce the usual conditions that utility maximization imposes on candidate demand functions, i.e. homogeneity of degree zero, symmetry and adding-up. As far as the authors know, and with the exception of the complex procedure of Dong *et al.* (2004), no author has been able to solve this in a satisfactory way. In this paper we have made use of the Tauchmann's proposal [Tauchmann, 2010], who suggests a simple consistent estimation method compatible with the imposition of the above mentioned usual conditions.

Expenditure and price elasticities show that all the expenditure items can be considered ordinary and normal goods. Making use of these elasticities we estimate the welfare loss suffered by Spanish households across income levels. In absolute terms the welfare loss and the increment in the tax bill increase with the income, but not in a linear fashion. The reason is that higher levels of income are associated to higher levels of expenditure on food and non-alcoholic beverages, but this kind of expenditure represents a higher proportion of the annual income for lower income levels. Thus, the 2012 VAT reform in Spain, and focusing on this expenditure group, can be considered as regressive. Finally, concerning social welfare, the King's proportional increase in equivalent income shows that, regardless the inequality aversion degree, the VAT reform induced a social welfare loss, but it would have caused a lower social loss the higher the inequality aversion of Spanish society.

APPENDIX: Individual and Social Welfare Change

For each household, the welfare change arising from the tax reform can be estimated in different ways following standard microeconomics. One possible way is the *equivalent variation*, EV_h , or the amount of money which would have to be given to the h -th household when it faces the initial price vector, \mathbf{p}_h^0 , to make it as well off as it would be facing the new price vector, \mathbf{p}_h^1 , with its initial income, m_h [Gravelle and Rees (2004)]. More formally, upon denoting this household's indirect utility function by V , EV_h is implicitly defined by $V(m_h + EV_h, \mathbf{p}_h^0) \equiv V(m_h, \mathbf{p}_h^1)$, so that $EV_h < 0$ for $\mathbf{p}_h^1 \geq \mathbf{p}_h^0$, $\mathbf{p}_h^1 \neq \mathbf{p}_h^0$. Thus, from Eqs. (1)-(4) one can explicitly solve for EV_h as

$$EV_h = a(\mathbf{p}_h^0) \times \exp \left\{ \frac{b(\mathbf{p}_h^0) \times \ln [m_h/a(\mathbf{p}_h^1)]}{b(\mathbf{p}_h^1) + [\lambda(\mathbf{p}_h^1) - \lambda(\mathbf{p}_h^0)] \times \ln [m_h/a(\mathbf{p}_h^1)]} \right\} - m_h. \quad (\text{A.1})$$

As a closely related concept, one could also define the *final equivalent expenditure*, EE_h^F , as

$$V(EE_h^F, \mathbf{p}_h^0) \equiv V(m_h, \mathbf{p}_h^1), \quad (\text{A.2})$$

(i.e. the expenditure required at the pre-reform prices to attain the same level of utility as with the post-reform prices) so that $EE_h^F \equiv m_h + EV_h < m_h$ for $\mathbf{p}_h^1 \geq \mathbf{p}_h^0$, $\mathbf{p}_h^1 \neq \mathbf{p}_h^0$.

As an alternative to EV_h , one can also consider the *compensating variation*, CV_h , or the amount of money which must be taken from the h -th household's income, m_h , when facing the new price vector, \mathbf{p}_h^1 , in order to make it as well off as it was when it faced the old price vector, \mathbf{p}_h^0 [Gravelle and Rees (2004)]. In other words, CV_h is implicitly defined as $V(m_h - CV_h, \mathbf{p}_h^1) \equiv V(m_h, \mathbf{p}_h^0)$, so that $CV_h < 0$ for $\mathbf{p}_h^1 \geq \mathbf{p}_h^0$, $\mathbf{p}_h^1 \neq \mathbf{p}_h^0$. From Eqs. (1)-(4) one has that CV_h is explicitly solved for as

$$CV_h = m_h - a(\mathbf{p}_h^1) \times \exp \left\{ \frac{b(\mathbf{p}_h^1) \times \ln [m_h/a(\mathbf{p}_h^0)]}{b(\mathbf{p}_h^0) + [\lambda(\mathbf{p}_h^0) - \lambda(\mathbf{p}_h^1)] \times \ln [m_h/a(\mathbf{p}_h^0)]} \right\}. \quad (\text{A.3})$$

As was the case with the equivalent variation, the compensating variation allows one to define the *initial equivalent expenditure*, EE_h^I , as

$$V(EE_h^I, \mathbf{p}_h^1) \equiv V(m_h, \mathbf{p}_h^0), \quad (\text{A.4})$$

(i.e. the expenditure required at the post-reform prices to attain the same level of utility as with the pre-reform prices) so that $EE_h^I \equiv m_h - CV_h > m_h$ for $\mathbf{p}_h^1 \geq \mathbf{p}_h^0$, $\mathbf{p}_h^1 \neq \mathbf{p}_h^0$. Both EE_h^F and EE_h^I are, therefore, monetary measures of the h -th household's welfare after the tax reform which can easily computed after Eqs. (A.1) and (A.3) respectively.¹¹

¹¹Note that $EE_h^I > EE_h^F$ for $\mathbf{p}_h^1 \geq \mathbf{p}_h^0$, $\mathbf{p}_h^1 \neq \mathbf{p}_h^0$.

From an efficiency perspective, however, we should consider not only the compensating (or the equivalent) variations suffered by households, but also the *change* in tax revenues. This leads us, therefore, to two alternative measures of deadweight loss or excess burden given by

$$EB_h^{ev} \equiv EV_h - (R_h^1 - R_h^0), \quad (\text{A.5})$$

and

$$EB_h^{cv} \equiv CV_h - (R_h^1 - R_h^0), \quad (\text{A.6})$$

which, given Eqs. (16), (A.1) and (A.3) and denoting the pre-reform and post-reform household's tax bill as R_h^0 and R_h^1 , respectively, can be easily computed.

As a complement to these individual welfare change measures, we also consider the welfare effects from a social point of view, or the social value of the reform. Borrowing from the tradition in the related literature, we assume the existence of an indirect social welfare function, W , defined in terms of the vector of *equivalent expenditures* (EE_1, EE_2, \dots, EE_H) given by

$$W(\cdot) = \begin{cases} \frac{1}{H} \sum_{h=1}^H \frac{EE_h^{1-\varepsilon}}{1-\varepsilon}, & \text{for } \varepsilon \neq 1 \\ \frac{1}{H} \sum_{h=1}^H \ln EE_h, & \text{for } \varepsilon = 1 \end{cases}, \quad (\text{A.7})$$

where parameter ε captures the degree of aversion to social inequality. [See Atkinson (1970).] From Eq. (A.7) one can derive a measure of social value, *the proportional increment in initial equivalent expenditure*, which we denote by λ , an which is defined as follows: *the proportional increase in initial equivalent expenditure that would make it possible to match the social welfare created by the reform*. Or, more formally,

$$W(\lambda \times EE_1^I, \lambda \times EE_2^I, \dots, \lambda \times EE_H^I) = W(EE_1^F, EE_2^F, \dots, EE_H^F), \quad (\text{A.8})$$

so that, given that $EE_h^I > EE_h^F$ as a result of $\mathbf{p}_h^1 \geq \mathbf{p}_h^0$ and $\mathbf{p}_h^1 \neq \mathbf{p}_h^0$, a value $\lambda < 1$ denotes a social welfare *loss* induced by the reform.

Along the same lines, the equivalent expenditure function can also be used to construct *inequality indices* defined on the distribution of equivalent expenditure. Borrowing from Prieto-Rodríguez *et al.* (2005) who, in turn, follow Atkinson (1970) and Sen (1973), we define the *equally distributed equivalent expenditure*, G , as the equivalent expenditure level that, distributed equally among all households, would provide the same level of social welfare as the actual distribution of equivalent expenditure. We can define two alternative expressions for G , depending on whether we consider the initial equivalent expenditure, G^I , or the final equivalent expenditure, G^F , and whose precise definitions are given by

$$W(G^I, G^I, \dots, G^I) \equiv W(EE_1^I, EE_2^I, \dots, EE_H^I) \quad (\text{A.9})$$

and

$$W(G^I, G^F, \dots, G^F) \equiv W(EE_1^I, EE_2^I, \dots, EE_H^I), \quad (\text{A.10})$$

which, given Eq. (A.7) can conveniently be solved to yield G^I and G^F as

$$G^I(\varepsilon) = \left[\frac{1}{H} \sum_{h=1}^H (EE_h^I)^{1-\varepsilon} \right]^{\frac{1}{1-\varepsilon}}, \quad \text{and} \quad G^F(\varepsilon) = \left[\frac{1}{H} \sum_{h=1}^H (EE_h^F)^{1-\varepsilon} \right]^{\frac{1}{1-\varepsilon}},$$

for $\varepsilon \neq 1$, and

$$G^I(\varepsilon) = \exp \left\{ \frac{1}{H} \sum_{h=1}^H \ln EE_h^I \right\}, \quad \text{and} \quad G^F(\varepsilon) = \exp \left\{ \frac{1}{H} \sum_{h=1}^H \ln EE_h^F \right\},$$

for $\varepsilon = 1$. Three remarks follow. First, note that from Eqs. (A.7), (A.9) and (A.10) the welfare change can be easily computed as

$$\lambda(\varepsilon) = \frac{G^F(\varepsilon)}{G^I(\varepsilon)}, \quad (\text{A.11})$$

where $G^I(\varepsilon)$ and $G^F(\varepsilon)$ have just been obtained immediately above, and both G^I and G^F are explicitly dependent on ε , the parameter reflecting the degree of aversion to social inequality for the social welfare function W in Eq. (A.7). Second, denoting the average values of the equivalent expenditures $\overline{EE^I} \equiv H^{-1} \sum_{h=1}^H EE_h^I$ and $\overline{EE^F} \equiv H^{-1} \sum_{h=1}^H EE_h^F$, it is the case that $\overline{EE^I} = G^I(\varepsilon)$ and $\overline{EE^F} = G^F(\varepsilon)$ if and only if $\varepsilon = 0$; that is to say, equally distributed equivalent expenditures equal average equivalent expenditures if and only if there is no inequality aversion. And, third, $\overline{EE^I} > G^I(\varepsilon)$ and $\overline{EE^F} > G^F(\varepsilon)$ if and only if $\varepsilon > 0$.

Tables and figures

TABLE I. EXPENDITURE ITEMS AND VAT RATES

No.	<i>Expenditure Categories</i> <i>Description</i>	<i>VAT rate (%)</i>	
		<i>pre-reform</i>	<i>post-reform</i>
1	Bread & cereals	4	4
2	Other bakery products	8	10
3	Meat products	8	10
4	Fish	8	10
5	Milk & eggs	4	4
6	Other dairy products	8	10
7	Oils and fats	8	10
8	Raw fruits and vegetables	4	4
9	Processed fruit and vegetables	8	10
10	Sugar, jam, honey, chocolate, confectionery and ice creams	8	10
11	Coffee, tea & cocoa	8	10

Key to Table I: Expenditure items in the demand system, and VAT rates in percent terms before and after the 2012 reform.

TABLE II. DESCRIPTIVE STATISTICS:
DEPENDENT VARIABLES

<i>Share</i>	<i>Mean</i>	<i>Median</i>	<i>SD</i>	<i>Min</i>	<i>Max</i>	<i>zeros (%)</i>
w_1	0.165	0.141	0.119	0.000	1.000	1.560
w_2	0.066	0.048	0.074	0.000	1.000	19.419
w_3	0.187	0.167	0.119	0.000	1.000	3.758
w_4	0.124	0.100	0.110	0.000	1.000	12.598
w_5	0.147	0.133	0.091	0.000	1.000	2.609
w_6	0.028	0.009	0.044	0.000	0.558	44.278
w_7	0.037	0.016	0.060	0.000	0.844	36.810
w_8	0.067	0.052	0.061	0.000	1.000	1.995
w_9	0.040	0.029	0.047	0.000	1.000	22.767
w_{10}	0.037	0.023	0.053	0.000	1.000	25.583
w_{11}	0.101	0.082	0.089	0.000	1.000	9.483

Key to Table II: main descriptive statistics for the budget shares, w_i , shown on column 1. See Table I for a description of expenditure items.

TABLE III. DESCRIPTIVE STATISTICS:
INDEPENDENT VARIABLES

<i>Variable</i>	<i>Mean</i>	<i>Median</i>	<i>SD</i>	<i>Min</i>	<i>Max</i>
p_1	2.622	2.533	0.936	0.227	34.142
p_2	4.527	4.400	6.313	0.320	477.813
p_3	8.223	7.926	4.027	0.332	140.649
p_4	8.648	8.485	5.317	1.000	242.210
p_5	2.352	1.839	1.988	0.056	33.993
p_6	2.615	2.382	3.392	0.442	196.742
p_7	2.462	2.310	2.350	0.474	200.874
p_8	1.701	1.502	1.028	0.264	56.499
p_9	4.382	3.906	3.808	0.275	278.167
p_{10}	5.086	4.464	4.855	0.306	190.276
p_{11}	2.144	1.147	3.563	0.060	134.432
m	2026.815	1805.084	1342.304	0.530	62,006.530
mem_1	0.145	0.000	0.408	0.000	3.000
mem_2	0.284	0.000	0.609	0.000	6.000
mem_3	0.245	0.000	0.547	0.000	5.000
mem_4	0.401	0.000	0.683	0.000	5.000
mem_5	1.121	1.000	0.882	0.000	5.000
mem_6	0.408	0.000	0.695	0.000	4.000
mem_7	0.048	0.000	0.230	0.000	3.000

Key to Table III: main descriptive statistics for the dependent variables (prices, nominal expenditure, and household age composition) shown on column 1.

TABLE IV. MARSHALLIAN AND EXPENDITURE ELASTICITIES FOR QUAIDS MODEL

Elasticities	j										
	1	2	3	4	5	6	7	8	9	10	11
$E_{1,j}^M$	-0.315	0.066	-0.096	0.055	-0.044	0.010	-0.033	0.006	0.085	0.030	0.044
$E_{2,j}^M$	-0.062	-1.058	-0.011	0.052	-0.060	0.006	0.022	0.006	-0.022	-0.050	-0.007
$E_{3,j}^M$	-0.160	0.020	-0.762	0.064	0.040	-0.021	-0.033	-0.019	0.021	-0.000	-0.010
$E_{4,j}^M$	-0.096	0.020	0.038	-1.195	-0.044	0.010	0.043	-0.017	0.024	-0.028	0.003
$E_{5,j}^M$	-0.204	-0.045	-0.010	-0.044	-0.954	0.012	0.025	-0.028	-0.002	-0.018	0.001
$E_{6,j}^M$	-0.175	0.008	-0.186	0.048	0.072	-1.409	0.255	0.214	0.105	-0.134	-0.054
$E_{7,j}^M$	-0.390	-0.050	-0.328	0.033	-0.002	0.190	-1.784	0.063	0.051	0.180	0.016
$E_{8,j}^M$	-0.091	0.011	-0.096	-0.022	-0.050	0.138	0.070	-0.991	0.058	-0.049	-0.046
$E_{9,j}^M$	0.030	-0.062	-0.012	0.066	0.002	0.082	0.104	0.067	-1.592	-0.084	-0.060
$E_{10,j}^M$	0.006	-0.064	-0.004	-0.032	0.001	-0.111	0.247	-0.064	-0.058	-0.660	-0.037
$E_{11,j}^M$	-0.087	-0.003	-0.063	0.019	0.027	-0.015	0.044	-0.026	-0.012	-0.026	-0.978
E_i^n	0.194	1.181	0.859	1.241	1.257	1.200	2.014	1.061	1.451	0.790	1.110

Key to Table IV: $E_{i,j}^M$ denotes Marshallian price elasticity of the i -th expenditure item with respect to the j -th price [see Eq. (7)]; E_i^n denotes expenditure elasticity of good i [see Eq. (6)]. In order to avoid the extreme values effect, the table reports the median elasticities across households rather than the more often used means.

TABLE V. MEDIAN VARIATION IN HOUSEHOLD'S TAX BILL AND WELFARE

R^0	R^1	ΔR	t^0	t^1	EV	CV	EB^{ev}	EB^{cv}	EE^I	EE^F
266.2	373.9	56.1	6.5	7.9	-57.6	-58.5	-125.6	-126.7	4429.2	4315.7

Key to Table V. R^0 : pre-reform tax revenue; R^1 : post-reform tax revenue; ΔR : change in tax revenue; t^0 : pre-reform VAT rate (%); t^1 : post-reform VAT rate (%); EV : equivalent variation; CV : compensating variation; EB^{ev} : excess burden for the equivalent variation; EB^{cv} : excess burden for the compensating variation; EE^I : equivalent initial expenditure; EE^F : equivalent final expenditure. All figures are measured at their median values in 2011 euros. See Appendix for formal definitions.

TABLE VIA. DISTRIBUTIVE ANALYSIS OF WELFARE

<i>Income decile</i>	<i>Expe.</i>	ΔR	EV	CV	EB^{ev}	EB^{cv}	EE^I	EE^F
1	2583.7	35.6	-33.0	-33.5	-80.5	-81.2	2614.8	2549.9
2	3319.2	42.9	-43.1	-43.8	-97.8	-98.7	3361.4	3276.9
3	3919.9	49.1	-52.1	-52.8	-114.5	-115.4	3973.9	3866.6
4	4161.8	51.4	-54.8	-55.6	-117.2	-118.3	4219.6	4104.9
5	4395.3	54.7	-57.5	-58.4	-120.9	-122.0	4454.3	4337.2
6	4610.8	56.1	-60.9	-61.8	-129.6	-130.6	4671.1	4551.4
7	4964.2	60.1	-66.2	-67.2	-134.2	-135.5	5030.7	4898.5
8	4994.8	63.3	-66.9	-67.9	-139.2	-140.4	5060.5	4927.9
9	5284.9	64.2	-70.9	-72.1	-144.0	-145.1	5357.2	5211.9
10	6094.8	70.5	-81.9	-83.1	-158.5	-159.8	6181.3	6006.6
<i>Median</i>	4371.8	56.1	-57.6	-58.5	-125.6	-126.7	4429.2	4315.7

Key to Table Via. Breakdown of household's tax and welfare change in Table V by income deciles. Expe. denotes annual expenditure. See key to Table V for the rest.

TABLE VIb. DISTRIBUTIVE ANALYSIS OF WELFARE RELATIVE TO ANNUAL INCOME

<i>Income decile</i>	<i>Expe.</i>	ΔR	<i>EV</i>	<i>CV</i>	EB^{ev}	EB^{cv}	EE^I	EE^F
1	36.72	0.51	-0.47	-0.47	-1.16	-1.17	37.19	36.26
2	32.31	0.40	-0.42	-0.43	-0.92	-0.93	32.74	31.90
3	27.56	0.35	-0.36	-0.37	-0.78	-0.79	27.92	27.21
4	25.30	0.32	-0.34	-0.34	-0.74	-0.74	25.64	24.97
5	21.57	0.27	-0.28	-0.29	-0.59	-0.59	21.87	21.29
6	19.34	0.24	-0.26	-0.26	-0.54	-0.55	19.61	19.08
7	17.99	0.21	-0.24	-0.24	-0.49	-0.50	18.23	17.76
8	15.25	0.19	-0.20	-0.21	-0.43	-0.43	15.46	15.05
9	13.47	0.16	-0.18	-0.18	-0.36	-0.36	13.64	13.29
10	10.58	0.13	-0.14	-0.14	-0.28	-0.28	10.72	10.43
<i>Median</i>	19.89	0.24	-0.26	-0.27	-0.53	-0.53	20.14	20.62

Key to Table VIb. Breakdown of household's tax revenues and welfare change in Table V by income deciles. See key to Table V. All figures are expressed in per cent terms *relative* to annual income.

TABLE VII. KING'S PROPORTIONAL INCREASE IN INITIAL EQUIVALENT INCOME

<i>Inequality Aversion, ε</i>	<i>King's $\lambda(\varepsilon)$</i>
0	0.973
0.5	0.973
1.0	0.974
1.5	0.974
2.0	0.976
2.5	0.979
3.0	0.983
3.5	0.984
4.0	0.985

Key to Table VIII. ε : Inequality aversion index [see Eq. (A.7)]; $\lambda(\varepsilon)$: King's index [See Eq. (A.11).]

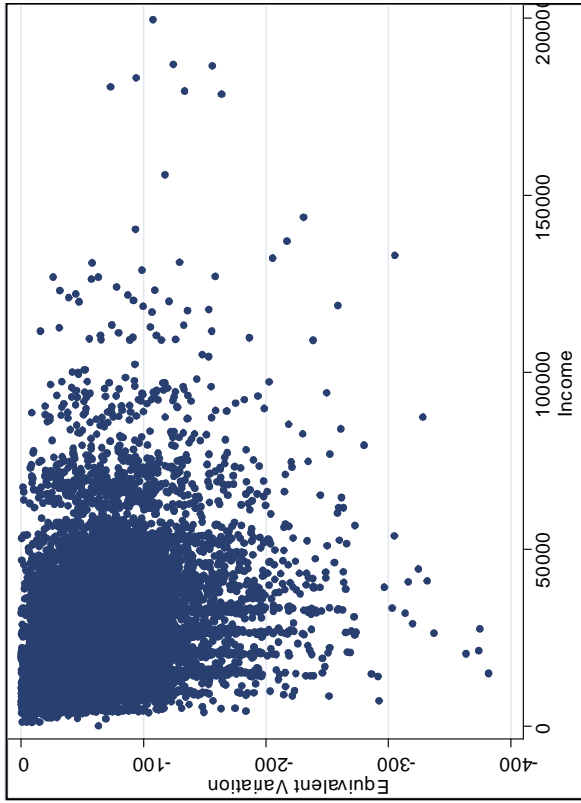


Figure 1.A: Equivalent Variation.

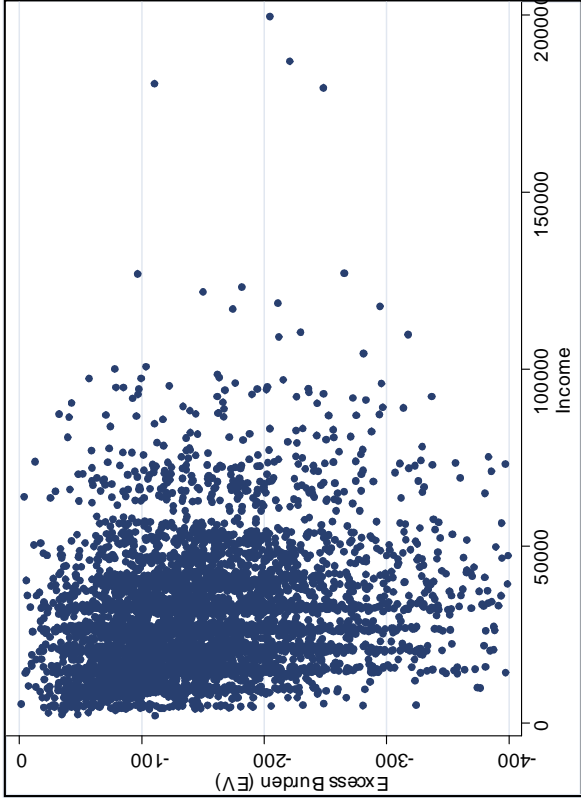


Figure 1.B: Excess Burden (Equivalent Variation).

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