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Demand for culture in Spain and the 2012 VAT rise

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Abstract This paper analyzes the effects that the 2012 VAT rise in Spain had on household demand for cultural goods and services. Household demands are modeled as a two-stage QUAIDS. After estimating price and expenditure elasticities, and the pass-through parameter associated with the reform, our results show that the individual welfare loss and the increment in the tax bill increase, but less than proportionately, with income. Consequently, the reform can be considered as regressive. Relating the effects of the VAT reform to households incomes also implies a low quantitative effect, because of the low proportion of total household expenditure that cultural expenditure represents. From a social perspective, the size of the induced welfare loss would positively depend on societys inequality aversion. Regardless of the latter, it cannot be concluded whether the reform would have increased or reduced inequality in the distribution of cultural spending. Our results prove qualitatively robust to alternative values of the pass-through parameter.

Keywords Culture · Two-stage budgeting · QUAIDS model · VAT reform · Welfare · Spain

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1 Introduction

The effects that economic policies in general, or in particular fiscal policies, have on individuals' welfare represent a relevant issue in economic research. Given that policies affect economic agents' decisions, a key aspect of policy evaluation is that of the assessment of the welfare implications for societies. Not only at the aggregate and the individual levels, but also noting how the effects are distributed among different segments of the population, what allows one to ascertain the degree of progressivity of the policy at issue.

One particular strand of this literature is that on indirect taxation issues, such as reforms of taxes and subsidies, whether potential or effectively implemented. Since the seminal papers of King (1983a and 1983b), the economic literature has generated a now well-established analysis procedure. First, a household utility-generated demand system is estimated to provide price and income elasticities, so that the researcher can predict households' reactions to policy changes. And, second, welfare effects are quantitatively evaluated on the basis of such reactions.

Despite the fact that the seminal papers were published almost four decades ago, just a few empirical applications have been published since then, among which a sample follows. Banks et al. (1997) simulate the effects of the imposition of a 17.5% sales tax on clothing under the ongoing VAT regime in the U.K. at that time. Urzúa (2001) evaluates the welfare impact of two indirect-tax reforms (changes in VAT and excise taxes) that took place in Mexico in 1995 and 1998. Ramadan and Thomas (2011) estimate the negative welfare change measures of different alternatives suggested to eliminate subsidies on selected food groups in Egypt. Janský (2014) simulates two effective changes in VAT legislation in the Czech Republic implemented in 2012 and 2013, plus a proposed change postponed until 2016. Attanasio and Lechene (2014) analyze the impact on the structure of food consumption of a cash transfer program in rural Mexico.

Focusing on Spain, Labeaga and López (1994) estimate the welfare impact of the 1992 VAT reform, and 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. García-Enríquez and Echevarría (2016), focalizing on food and non-alcoholic beverages, studies the effects that the 2012 VAT reform in Spain had on households' welfare.

In this paper, using data from the 2011 Spanish Household Budget Survey (SHBS), we study the increase in VAT rates introduced in Spain in mid 2012, focusing on cultural goods and services rather than on the whole set of expenditure groups that form the Spanish average household's consumption bundle, and which represented 2.63% of the average household's total expenditure in 2011. Our main motivation lies in the fact that, as a result of the reform, cultural goods and services suffered the most remarkable increment in the tax rates, thereby raising concern and discomfort among both Spanish producers and customers of culture. Thus, *i*) the VAT (weighted) average tax rate applied to cultural goods and services was raised from 13.04% in 2011 to

18.14% in 2013 (39.11% higher) as most sub-items ended up being taxed at a 21% rate; but *ii*) in the particular case of cultural services (shows, museums, internet, radio and TV licenses) the average tax rate was raised from 13.55% in 2011 to 20.89% in 2013 (54.17% higher) [see Table 1]. Not surprisingly, employers and workers of live shows and cinema rejected the measure with greater intensity.[1]

TABLE 1. CULTURE AND VAT RATES IN SPAIN

		VAT rate (%)	
No.	*Cultural items description*	*pre-reform*	*post-reform*
1	*Audio, video and computer equipment, and musical instruments*	**18.00**	**21.00**
	- Audio reception, recording and reproduction	18.00	21.00
	- TV, VCR and DVD	18.00	21.00
	- Photograph and cinema	18.00	21.00
	- Computers and peripherals	18.00	21.00
	- Recordables (CDs, DVDs, cards,...)	18.00	21.00
	- Audio, video and computer equipment repair	18.00	21.00
	- Musical instruments	18.00	21.00
2	*Shows, museums, internet, radio and TV licenses*$^{(*)}$	**13.55**	**20.89**
	- Dances, cinemas, theaters and shows	8.00	21.00
	- Museums, botanical gardens, libraries	8.00	10.00
	- License fees for radio and television	8.00	21.00
	- Rental of TV and video equipment	18.00	21.00
	- Services (private parties, photograph, pets)	8.00	21.00
	- Internet related services	18.00	21.00
3	*Non-text books & periodicals*	**4.00**	**4.00**
	- Non-text books	4.00	4.00
	- Periodicals	4.00	4.00
Average$^{(*)}$		**13.04**	**18.14**

Key to Table 1: VAT rates in percent terms before and after the 2012 reform. $^{(*)}$Weighted average, budget shares used as weights.

To further underline the motivation of our paper, Table 2 gives us a bird's eye view of cultural expenditure per household in Spain for the period 2011-2015. Four features are worth mentioning. First, the household's average real expenditure in cultural goods and services has experienced a cumulative fall of 21%. Second, the household's average share of expenditure in culture has fallen from 2.63% to 2.26%. Third, the number of households with no cultural expenditure has increased from 2.9 millions in 2011 (16.8% total households) to 3.7 millions in 2015 (19.9% of total households). And, fourth, the reasons

[1] As of June 28th 2017, the VAT rates applicable to live shows (and, therefore, excluding cinema) were brought back down to their pre-reform value, i.e. 10%.

for this observed pattern seem to be diverse and not mutually exclusive, but the deep economic recession suffered by the Spanish economy in those years and the VAT policy implemented might have played a role: as shown in column 4, average household real income fell by 7.24%.

Table 2. Cultural Expenditure per household in Spain in 2011-2015

Year	*Households surveyed*	*Households represented*	*ARI*	*ARE*	*AES*	*Households with no cultural expenditure*
2011	22,119	17,342,147	23,351	825	2.63	2,912,958
2012	21,808	18,091,838	21,685	731	2.50	3,498,249
2013	22,057	18,212,214	20,983	662	2.35	3,837,497
2014	22,144	18,301,426	20,997	652	2.32	3,846,116
2015	22,130	18,374,351	21,661	652	2.26	3,663,080

Key to Table 2: ARI: Average real income (in euros). *ARE*: Average real (cultural) expenditure (in euros). *AES*: Average (cultural) expenditure share (%). *SOURCE*: Spanish Household Budget Survey (SHBS), for several, years, conducted by the National Statistics Institute. Base year is 2011. See `http://www.ine.es/en/`

Among the growing body of the literature on cultural economics, and concentrating on the quantitative and econometric studies, one might distinguish two different approaches in, respectively, those papers which consider individual demand equations for some given specific cultural good or service (the vast majority of the references), and those which estimate a complete demand equation system.

Among the former, one particular case of interest is that of books and periodicals. Thus, one could mention, among others, Villarroya and Escardíbul (2010) who study the determinants of the consumption of such goods in Spain in 2006. Palma-Martos *et al.* (2009) study the book market in Spain between 1989 and 2006. Escardíbul and Villarroya (2009) study the consumption of newspapers in Spain. Jaén-García (2012) analyzes the demand for books and periodicals in Spain in 2006-2008. Hjorth-Andersen (2000) studies the market for books in Denmark for the period 1973-1993. Borowiecki and Navarrete (2015) estimate the effect of a drop in the VAT rate upon the price and the expenditure on books in EU-28 countries in the period from 1993 to 2013. Álamo-Cerrillo and Lagos-Rodríguez (2016) focus on the market for electronic books in Europe and the discriminatory fiscal treatment that these suffered, compared to their printed alternatives since electronic books are taxed at higher rates than paper books.

Researchers have also paid attention to other kinds of cultural goods and services apart from books and periodicals. For instance, Fernández-Blanco *et al.* (2013) study the demand for ticket movies in Spain in 2003-2005. Devesa *et al.* (2009) estimates the demand for tickets at a cultural festival, the Valladolid

International Film Festival in 2001. One last example of cultural item demand may be that of performing arts, Seaman (2006) thorough survey being a classical reference. Focusing on econometric studies, it is claimed as a criticism that while income and price elasticities are the usual end-products of empirical demand analysis, a substantial portion of the performing arts demand literature does not derive such elasticities, which suggests as a conclusion that "a notable part of this literature [is] devoted instead to more broadly examining the competing determinants of arts attendance or participation patterns without any formal link to the neoclassical theory of consumer". Along these lines, and highlighting those few studies reporting either own-price or income elasticities, or both, performing arts are really luxury goods. Regarding own-price elasticity, however, no clear conclusion can be reached about whether the demand for performing arts is price elastic or inelastic [see Table 1, p.11 in Seaman (2006)].

Unlike the above numerous cases where the demand for specific or isolated cultural items has been studied, the number of previous works in the literature that have considered the subset of all cultural goods and services as a part of a broader set comprising all the goods and services purchased by the consumption unit (individual consumers or, more generally given the availability of data, households) is relatively scarce. This is the case, of course, with complete demand system studies which is where this paper falls. Ringstad and Løyland (2006) estimate a three-good AIDS demand system (books, other cultural goods and non-cultural goods) of Norwegian households for the period 1986-1999. Ringstad and Løyland (2011) study a six-good AIDS demand system (live performing arts, cinema, printed media, audiovisual media, sports and other goods) of Norwegian households for the period 1986-2002. Finally, Håkonsen and Løyland (2016) estimate a demand equation system for 8 categories of cultural expenditure in Norwegian municipalities for the period 2002-2010.

The closest precedent to our paper, and on which it partially draws, is Prieto-Rodríguez *et al.* (2005) where an AIDS model is estimated for the Spanish economy. The demand system considers 19 different groups of goods, including 3 cultural goods: *i*) cinema, theater, and museum and other events; *ii*) books, magazines and newspapers; and *iii*) film and music on magnetic media. Following the standard procedure, once expenditure and price elasticities are obtained, 3 (hypothetical) alternative cuts in the VAT rate on cultural goods are microsimulated and evaluated in terms of revenue and welfare, concluding that the suggested fiscal reforms would lead to regressive welfare and efficiency gains.

Our model departs from Prieto-Rodríguez *et al.* (2005) in several aspects. First, we adopt the Quadratic Almost Ideal Demand System (QUAIDS) model introduced by Banks *et al.* (1997), an extension to the AIDS model introduced in Deaton and Muellbauer (1980a). Second, we estimate a two-stage demand system, so that expenditure and price elasticities are properly estimated, *i.e.* we obtain unconditional elasticities. In particular, cultural goods and services, one of the five expenditure groups included in the first stage, is split into three sub-groups: *i*) audio, video and computer equipment and musical instruments; *ii*) shows, museums, internet, radio and TV licenses; and *iii*) non-text books and periodicals. Third, we correct for the potential bias introduced by endogeneity of both cultural expenditure and total expenditure. Fourth, we estimate the pass-through parameter of the change in VAT tax rates, thereby abandoning the usual assumption that producers shift 100% of the changes in VAT rates to customers. This allows us to carry out some sensitivity analysis exercises concerning the incidence of VAT changes and their effect on producers' and consumers' prices. Fifth, we include a consistent treatment of zero observations, which are not uncommon in consumption and expenditure data, especially when dealing with microeconomic data. The nature of a null consumption of a certain good can be multiple: corner solutions, households genuine non-consumers of the good, absence of consumption during the survey period, etc. Unfortunately, the Spanish Household Budget Survey does not detail the origin of the zeros. Thus, without prior information, it is not possible to be certain about the precise cause of the observed zeros [Meghir and Robin (1992)], so that the choice of empirical specification is typically based on assumptions. In this work we assume that households are true non-consumers of the good, so that a standard sample selection model is proposed, and a generalized Heckman estimator by Tauchmann (2010) is applied.[2]

Regarding the last point, notice that in order to analyze the welfare effects of tax reforms such as the one considered here, one first needs to ensure that the demand system is utility-generated and, consequently, that the corresponding parameter restrictions hold. As in the AIDS model, homogeneity of degree zero and symmetry are not considered to be a problem.[3] Adding-up, however, as is explained later in the paper, is hard to reconcile with a proper treatment of zero expenditure observations [Drichoutis et al., (2008)]. In an attempt to overcome this shortcoming, we implement the consistent two-step estimator introduced in Tauchmann (2010) which, to the best of our knowledge, has been previously applied to the estimation of a demand system in only two works [García-Enríquez and Echevarría (2016) and Gálvez et al. (2016)]. A valuable feature of this procedure in the context of demand system estimation is that it allows one not only to deal with censored data, but also to properly impose the adding-up restriction.

Once the QUAIDS model is properly estimated, we use the price and expenditure elasticities to estimate the expected reaction of Spanish households' demand for culture, following the 2012 VAT reform, in terms of tax revenues, equivalent and compensating variations, and equivalent and compensating initial and final expenditures both in absolute levels and relative to income dis-

[2] If, for instance, the households' optimal consumption decision were zero, then we could propose a double hurdle model [Cragg (1971)]; if the reason were that the households have not purchased the good due to the short time of the survey (in Spain the households cooperate in the survey for two weeks a year), a good alternative would be a purchase infrequency model [Blundell and Meghir (1987)].

[3] Nevertheless, negative semi-definiteness of the Slutsky substitution term matrix can be neither imposed nor tested.

tribution. Thus, in this paper we focus on the consumers' welfare, therefore leaving aside the effect on producers' side. Our main results follow. First, estimates of the own-price elasticities have the expected negative sign and are inelastic. Second, the cultural goods and services considered in the second stage are normal goods; and, in particular, non-text books and periodicals are luxury goods. But, third, considering the set of all cultural goods and services, we show that expenditure on culture grows less than proportionately with income. Fourth, consequently, households with lower income experience a greater welfare loss relative to their income levels, so that the reform, as regards this expenditure group, can be considered as regressive as it induced less than proportional welfare losses for higher income households. Fifth, relating the effects of the VAT reform to households' incomes also implies, however, a very low quantitative effect, the reasons being *i*) the low proportion that cultural expenditure represents in Spanish households' total expenditure, with a mean of 2.63% and a median of 1.96%, and *ii*) the low value of the pass-through parameter associated with VAT changes. Sixth, the size of social welfare loss caused by the reform positively depends on the size of the inequality aversion of Spanish society as a whole, while it is unclear whether the VAT reform would have diminished or raised the cultural spending distribution inequality among Spanish households. And, seventh, the results prove qualitatively robust to alternative assumptions about the pass-through parameter.

The paper is organized as follows. Section 2 introduces the 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 deals with our welfare analysis. Section 7 concludes. An Appendix section including formal definitions and tables with results of some robustness check exercise goes at the end.

2 Consumer demand: the QUAIDS model

The consumer demand system is modeled by using the QUAIDS model introduced by Banks *et al.* (1997). This can be considered as a generalization of the popular Deaton and Muellbauer (1980a) AIDS model as it includes the square of the logarithm of expenditure as an additional regressor. Therefore, any given good is allowed to be 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.

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

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

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

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

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

and

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

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

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

for $i = 1, 2, ..., n$.

It can be shown that the adding-up, homogeneity and Slutsky term matrix symmetry restrictions that utility maximization imposes on candidate demand functions satisfying Eq. (5) are given by the following set of conditions:

$$\sum_{i=1}^{n} \alpha_i = 1,\ \sum_{i=1}^{n} \beta_i = 0,\ \sum_{i=1}^{n} \gamma_{ij} = 0,\ \sum_{i=1}^{n} \lambda_i = 0,\ \gamma_{ij} = \gamma_{ji} \tag{6}$$

for all $i, j = 1, 2, ..., n$.[4]

Further algebraic manipulation yields the expressions for expenditure, Marshallian (uncompensated) and Hicksian (compensated) price elasticities, E_i^m, $E_{i,j}^U$ and $E_{i,j}^C$ respectively as

$$E_i^m = 1 + \frac{1}{w_i} \left[\beta_i + \frac{2\lambda_i}{b(\mathbf{p})} \ln \left\{ \frac{m}{a(\mathbf{p})} \right\} \right], \tag{7}$$

$$\begin{aligned} E_{i,j}^U = -\delta_{ij} + \frac{1}{w_i} \Bigg\{ \gamma_{ij} - \left[\beta_i + \frac{2\lambda_i}{b(\mathbf{p})} \ln \left\{ \frac{m}{a(\mathbf{p})} \right\} \right] \times \\ \left(\alpha_j + \sum_{k=1}^{n} \gamma_{jk} \ln p_k \right) - \frac{\beta_j \lambda_i}{b(\mathbf{p})} \left[\ln \left\{ \frac{m}{a(\mathbf{p})} \right\} \right]^2 \Bigg\}, \end{aligned} \tag{8}$$

and $E_{i,j}^C = E_{i,j}^U + E_i^m \times w_j$, for $i, j = 1, 2, ..., n$, and where $\delta_{ij} = 1$ if $i = j$, and $\delta_{ij} = 0$ otherwise.

Even though our focus in this paper is on a specific group of consumption items, we will follow a by now well-established procedure, theoretically justified

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

on the basis of so-called two-stage budgeting. In the first stage, households (or individual consumers) allocate total expenditure to different groups, thereby obtaining optimal expenditure levels for each group. In the second stage, expenditure on each of the groups is optimally assigned to each of the individual sub-groups within each group [see, among others, Menezes *et al.* (2008), Mittal (2010)]. In principle, restrictive assumptions (involving separability of preferences) are needed to guarantee the optimality of this budgeting procedure [see, *e.g.* Blackorby *et al.* (1998) and references therein.] For instance, *weak separability* is both necessary and sufficient for the *second* stage of two–stage budgeting [see Deaton and Muellbauer (1980b), p. 124]. Therefore, properly estimated expenditure and price elasticities require the estimation of the two stages: conditional elasticities estimated in the second stage must be adjusted after taking into account the elasticities obtained in the first stage as explained in detail below.

Along these lines Edgerton (1997) showed that "*...restricting the analysis to the last stage of a multistage budgeting process [leads] to considerable errors, which could well have policy consequences*", and found simple relationships between expenditure and price elasticities in different budgeting levels. Carpentier and Guyomard (2001) later found that the formulae for price elasticities obtained by Edgerton (1997) violate the symmetry condition except in the particular case of homothetic sub-utility functions. Thus, assuming that *i*) the direct utility function is weakly separable, and that *ii*) the price indices for each of the groups hardly vary with corresponding sub-utility levels, Carpentier and Guyomard (2001) obtained formulae for expenditure and own- and cross-price elasticities consistent with the demand properties of homogeneity, adding-up, and symmetry.

In particular, they showed that total (or unconditional) expenditure elasticity for good i in expenditure group G, E_i^M, is given by

$$E_i^M = E_i^{m_G} \times E_{m_G}^M, \tag{9}$$

where $E_i^{m_G} \equiv \partial \ln q_i / \partial \ln m_G$, (*i.e.* the partial or conditional second-stage elasticity of good i with respect to expenditure in group G, m_G), and $E_{m_G}^M \equiv \partial \ln m_G / \partial \ln M$ (*i.e.* the elasticity of expenditure on group G with respect to total expenditure, M), which coincides with the corresponding formula obtained by Edgerton (1997).

Concerning total (or unconditional) Marshallian price elasticity of, say, good i in expenditure group G with respect to price j in expenditure group H, $E_{i,j}^u$, this is given by

$$\begin{aligned} E_{i,j}^u = \hat{E}_{i,j}^u + w_{j,H} \times \left(\frac{\delta_{G,H}}{E_j^{m_H}} + E_{G,H}^U \right) \times E_i^{m_G} \times E_j^{m_H} \\ + w_{j,H} \times W_H \times E_{m_G}^M \times E_i^{m_G} \times (E_j^{m_H} - 1), \end{aligned} \tag{10}$$

where $\hat{E}_{i,j}^u$ denotes the conditional second-stage Marshallian price elasticity of good i with respect to price j, $w_{j,H}$ denotes the budget share of good j in expenditure group H (*i.e.* $w_{j,H} \equiv p_j q_j / \sum_{k=1}^{n_H} p_k q_k$), $E_{G,H}^U \equiv \partial \ln Q_G / \partial \ln P_H$

(*i.e.* the conditional first-stage Marshallian price elasticity of composite good G with respect to the price index of composite good H), W_H denotes the budget share of expenditure group H (*i.e.* $W_H \equiv P_H Q_H / \sum_{k=1}^{n_H} P_k Q_k$), and $\delta_{G,H}$ is the Kronecker delta (*i.e.* $\delta_{G,H} = 1$ when $G = H$ -both good i and good j belong to the same group- and $\delta_{G,H} = 0$ when $G \neq H$ -good i and good j belong to different groups).

3 The estimation strategy

The model introduced in Eq. (5) is modified to deal with estimation issues. First, the model admits the existence of additional regressors, other than the price vector and total expenditure, which potentially help to explain the household's consumption decision, such as the demographic variables. The standard procedure followed is to include these additional regressors in an additive manner [see, among others, Banks *et al.* (1997) and Attanasio and Lechene (2014)].

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 Blundell and Robin (1999), the QUAIDS model in Eq. (5) is augmented by introducing a correction term.

Third, we need to deal with the issue of zero consumption of one or more goods and services. In this case of a censored dependent variable Ordinary Least Squares (OLS) estimation of Eq. (5) will be biased and inconsistent.

Heien and Wessells (1990) was one of the first studies to address this point in the context of the estimation of an equation system. Suppose that the following system of equations characterizes the *latent* model

$$w_{ih}^* = f(\mathbf{x}_h, \theta_i) + \varepsilon_{ih}, \tag{11}$$

$$d_{ih}^* = \mathbf{z}'_h \pi_i + v_{ih}, \tag{12}$$

for $i = 1, 2, ..., n$, $h = 1, 2, ..., H$, 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; θ_i and π_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:

$$d_{ih} = \begin{cases} 1 & \text{if } d_{ih}^* > 0 \\ 0 & \text{if } d_{ih}^* \leq 0 \end{cases}, \tag{13}$$

$$w_{ih} = d_{ih} w_{ih}^*. \tag{14}$$

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. (12)] is estimated. In the second step, the inverse Mills ratio for each household and each good is computed and introduced as an additional regressor in the system in Eq. (11).

Shonkwiler and Yen (1999) find, however, that the Heien and Wessells (1990) estimation procedure is not consistent and, after rewriting the system in Eq. (11), they propose a new estimation procedure. 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)], which is often ignored [see, *e.g.* Zheng and Henneberry (2010) or Bakhshoodeh (2010)].

If the researcher simply pursued consistent estimation of a demand system as in Eq. (5) without caring about the parametric restrictions in Eq. (6), the Shonkwiler and Yen (1999) estimator would be *the* alternative: it is quite easy to implement and, not surprisingly, has become very popular in the applied literature. Nevertheless, one often needs not only estimates with good statistical properties, but *also* requires that the estimated demand system is truly utility-generated, *i.e.* one needs to make sure that the restrictions in Eq. (6) hold. In particular, this is a necessity if the demand system estimation is followed by some type of welfare analysis as here. Otherwise consumer welfare calculations will *not* be valid [Hausman and Leonard (2005)].

Dong *et al.* (2004) introduced a variation of the Amemiya-Tobin estimation procedure [Amemiya (1974)] allowing consistent imposition of the traditional parameter restrictions that utility maximization requires in a context of a censored model. However, it is hard to implement and, consequently, very few applied researchers actually use it.

Tauchmann (2010) later suggested an easier approach based on a two-step estimator that, instead of conditioning on only d_{ih}, conditions on the entire selection pattern, $d_h = [d_{1h}, ..., d_{nh}]'$, thereby obtaining a consistent generalized Heckman-type estimator. Such an estimator is implemented as follows. First, a multivariate probit for the observed version of Eq. (12) is estimated, and its results are used to build the following correction terms:

$$M_{jh} = k_{ih}\phi(\mathbf{z}_h'\widehat{\pi}_i)\frac{\Phi^{n-1}(\widetilde{A}_{jh}, \widetilde{R}_{jh})}{\Phi^{n}(\mathbf{z}_h'\widehat{\pi}_1, ..., \mathbf{z}_h'\widehat{\pi}_n)}, \tag{15}$$

for $j = 1, ..., n$, where $k_{ih} = 2d_{ih} - 1$ and $\phi(\mathbf{z}_h'\pi_i)$ is the univariate standard normal density function, 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. (12) and s_{lj}^{vv} the corresponding (l, j) element, $\widetilde{A}_{jh}$ represents a vector of $n-1$ elements $k_{lh}(\mathbf{z}_h'\widehat{\pi}_l - s_{lj}^{vv}\mathbf{z}_h'\widehat{\pi}_j)/\left(1-\left(s_{lj}^{vv}\right)^2\right)^{1/2}$, $l = 1, ..., n$, $l \neq j$. $\widetilde{R}_{jh}$ 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 in Eq. (15), are incorporated in the observed version of Eq. (11), giving rise to the following system

$$w_{ih} = d_{ih} f(\mathbf{x}_h, \theta_i) + d_{ih} \sum_{j=1}^{n} \rho_{ij} M_{jh} + d_{ih} \widetilde{\varepsilon}_{ih}, \tag{16}$$

where $\widetilde{\varepsilon}_{ih} = \varepsilon_{ih} - E(\varepsilon_{ih}|d_h)$. Importantly, 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_{jh}'s, are the *same* for all equations; this ensures that the estimates will be invariant to the particular equation dropped.

Applying this procedure to the QUAIDS model in Eq. (5), the system in Eq. (16) can be rewritten as

$$\begin{aligned} w_{ih} = d_{ih} & \left[\alpha_i + \sum_{j=1}^{n} \gamma_{ij} \ln p_{jh} + \beta_i \ln \left\{ \frac{m_h}{a(\mathbf{p}_h)} \right\} + \frac{\lambda_i}{b(\mathbf{p}_h)} \left[\ln \left\{ \frac{m_h}{a(\mathbf{p}_h)} \right\} \right]^2 \right] \\ & + d_{ih} \sum_{j=1}^{n} \rho_{ij} M_{jh} + d_{ih} \widetilde{\varepsilon}_{ih}. \end{aligned} \tag{17}$$

As a particular case of Tauchmann's procedure, Gálvez *et al.* (2016) proves that, when only one dependent variable is censored, all the M_{jh} terms tend to zero except the one associated with the censored good. For example, if the censure is only observed in the first good, then $M_{1h} \neq 0$ and $M_{jh} = 0$ $\forall j \neq 1$. Moreover, it can be easily proved that in this case M_{1h} would reduce to the inverse Mills ratio, that is, $M_{1h} = k_{1h} \phi(z_h' \widehat{\pi}_1) / \Phi(z_h' \widehat{\pi}_1)$. Thus, applying this result to the QUAIDS model in Eq. (5), the system in Eq. (16) can be rewritten as

$$\begin{aligned} w_{ih} = d_{ih} & \left[\alpha_i + \sum_{j=1}^{n} \gamma_{ij} \ln p_{jh} + \beta_i \ln \left\{ \frac{m_h}{a(\mathbf{p}_h)} \right\} + \frac{\lambda_i}{b(\mathbf{p}_h)} \left[\ln \left\{ \frac{m_h}{a(\mathbf{p}_h)} \right\} \right]^2 \right] \\ & + d_{ih} \rho_{i1} M_{1h} + d_{ih} \widetilde{\varepsilon}_{ih}, \end{aligned} \tag{18}$$

where in this particular case $d_{ih} \equiv 1$ for $i \neq 1$, $d_{1h} = 1$ if $w_{1h} > 0$, and $d_{1h} = 0$ if $w_{1h} = 0$. The relevance of this result will become apparent later on as it will substantially help in simplifying the estimation procedure of the first-stage demand system.

Following Tauchmann (2010) we estimate the QUAIDS models in Eqs. (17) and (18) as SUR systems. Moreover, as the equations are non-linear in the parameters, a non-linear estimation method must be applied.[5] Additionally, the

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

way in which theoretical restrictions that follow from utility maximization are imposed is trivial and intuitive, and the parameters are consistently estimated.

4 The data set

VAT was first introduced in Spain (in the whole country except the Canary Islands, Ceuta and Melilla) as a requirement for the official integration of Spain into the *European Economic Community* (*EEC*) on January 1, 1986. Since then, Spanish VAT legislation has gone through several reforms, the latest (the one that will be studied here) becoming effective on September 1, 2012 [see European Commission (2013)].

This tax law reform was a part of a major law reform designed as "*(a) set of measures to ensure budgetary stability and promoting competitiveness*". As a result, VAT rates were modified as follows: 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 of 4%, which applies to basic necessities such as vegetables, milk, bread, fruit, pharmaceutical products and books, newspapers and the like, 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 data, obtained from the SHBS, which is collected by the Spanish Statistical Office, correspond to 2011, the last whole year with tax rates prior to the VAT reform. The survey consists of three separate files: the household file, which provides general information about the household (region, municipality size, household size, household head features, main dwelling features, total expenditure or total income among others); the household member file, which provides information about nationality, educational attainment, labor status or revenues on each household member; 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 case.

Each household participates in the survey for two weeks a year, reporting all the goods and services that their members purchase. Information about those purchases with higher periodicity is collected by means of personal interviews during each fortnight. Half of the sample is renewed on a yearly basis, so that each family takes part in the survey for two consecutive years. In 2011 the sample contains data for 22,119 households. Each household is assigned a time and space scaling factor, so that these households represent a total of 17,342,147 households. In the first stage of the estimation process, some observations were eliminated for various reasons. More specifically, 119 observations were dropped for displaying an attributed level of income equal to 0. Additionally, 1,227 observations were dropped as they related to households living in the Canary Islands, Ceuta or Melilla. Two more observations were dropped because they provide no information about the living arrangement of the household head, a feature that will be considered among the regressors in the estimated model. And, finally, 72 additional observations were dropped as such a number of households reported zero expenditure on Group 2 (namely, food and non-alcoholic beverages). As a result, the sample was finally reduced to 20,699 households representing a total of 16,356,756 households.

As already stated above, in this paper we focus on just one broad expenditure group, namely *cultural goods and services*. To this end, we estimate a five composite good demand system in the first stage. More precisely, the 225 expenditure categories referred to above are grouped into five broad expenditure groups: 1) cultural goods and services, 2) food and non-alcoholic beverages, 3) housing (in a broad sense, including water, electricity, gas and other fuels, furniture, equipment and maintenance), 4) transport and communications, and 5) others (a heterogeneous group including, among others, alcoholic beverages, tobacco, clothing, footwear, health, leisure, entertainment, hotels, bars, coffee shops and restaurants, and education).

One remark concerning what is understood in this paper as cultural spending follows. As a principle, and following the criterion set by the Spanish Statistical Office (*INE*), the SHBS contains an expenditure group specifically devoted to cultural items, namely, *Group 09. Leisure, Shows and Culture*. The Spanish Ministry of Education, Culture and Sports, and following the SHBS, annually elaborates the Yearly Report on Cultural Statistics (*Informe Anual de Estadísticas Culturales*) "*with the aim of providing a selection of the most significant statistical results in the field from a variety of sources, to facilitate awareness of the situation and the evolution of culture in Spain, its social value and its character as a source of wealth and economic development in Spanish society*" [Ministerio de Educación, Cultura y Deporte (2011), (2016)]. It turns out, however, that the criteria followed by these two institutions are not coincident. In particular, some of the expenditure items included in Group 09 of the SHBS are excluded from the Ministerial yearly report just referred to, as it would be arguable whether they represent "strictly" cultural expenditure. That is the case, for instance, regarding *Great equipment related to sports and leisure outdoors*, or *Gardening and flowers*, or *Purchase of pets* or *All inclusive holidays*. And, conversely, one particular expenditure item not included in Group 09 (but in *Group 08. Communications*) is included in the Ministry's report as cultural spending: *Internet related services*. In this paper we have followed the same criterion as the Spanish Ministry, as we believe that it better represents what an average consumer would understand as cultural spending.

Some descriptive statistics associated with the dependent variables (*i.e.* budget shares for each of the 5 expenditure groups in the first stage) are reported in Table 3. Two remarkable features of Spanish households' taste for culture are shown in Table 3: *i)* the low expenditure share of cultural consumption, and *ii*) the high proportion of households which do not consume culture at all. Table 4 reports some descriptive statistics for the independent variables. These include, in addition to prices, total expenditure and total income, variables describing household features (age composition, household head's gender, educational attainment, living arrangement) and locational features (municipality size, region and whether the town of residence is a provincial capital or not).

The survey data has a major drawback as it does not include prices. This issue was dealt with by Cox and Wohlgenant (1986) in their seminal work, and since then authors have provided the literature with alternative solutions [see Majumder, Ray and Sinha (2012) for a thorough review of this issue]. In this paper we follow Hoderlein and Mihaleva (2008), who use household-level price indices instead of unit values. Thus, and focusing on this first-stage estimation process, a Stone-Lewbel like price index is computed first [Lewbel (1989)]. More precisely, for any group $I \in \{1, 2, 3, 4, 5\}$ which is made of, say, J_I expenditure sub-groups, we construct a price index at the household level $P_{Ih} \equiv (1/K_I) \prod_{j=1}^{J_I} \left(P_{I,j}/W_{I,j}^h\right)^{W_{I,j}^h}$, where $K_I \equiv \prod_{j=1}^{J_I} \overline{W}_{I,j}^{-\overline{W}_{I,j}}$, $\overline{W}_{I,j}$ being the mean of $W_{I,j}^h$, that is, the mean of the (household-level) expenditure share of expenditure sub-group j relative to total expenditure on group I, with $W_{I,j}^h \geq 0$, $\sum_{j=1}^{J_I} W_{I,j}^h = 1$ for all I and h, and where $P_{I,j}$ denotes the aggregate nation-wide price index for expenditure sub-group j.[6] In those cases in which the budget share $W_{I,j}^h$ happens to equal 0 for *all* j, the household at issue is assigned a price index equal to the mean value of all the other households.

Table 3. First Stage.
Descriptive Statistics: Dependent Variables (W_I)

No.	*Concept*	*Mean*	*Median*	*SD*	*Min*	*Max*	*zeros*
1	Cultural goods & services	0.026	0.020	0.029	0.0000	0.534	16.329
2	Food & non-alcoholic beverages	0.160	0.147	0.088	0.0003	0.766	0.000
3	Housing	0.397	0.376	0.154	0.0084	0.991	0.000
4	Transport & communications	0.125	0.096	0.109	0.0000	0.852	1.372
5	Others	0.292	0.283	0.136	0.000	0.879	0.169

No. of Observations: 20,699
Households Represented: 16,356,756

Key to Table 3: Main descriptive statistics for the budget shares, W_I, of groups shown in column 1. The proportion of zeros in column 8 is denoted in percent terms.

Regarding the second-stage estimation process, the set of fifteen expenditure sub-groups available within expenditure group 1 (cultural goods and services) have been reduced following two principles: *i*) homogeneity of the

[6] The Spanish Bureau of Statistics provides nation-wide average price indices for a total of 126 expenditure sub-groups.

goods grouped, and *ii*) common VAT tax rates both before and after the reform. Here one caveat is in order. One ideally would need to have exactly the same pre- and post-reform tax rates for all items included in a given sub-group, as is the case for sub-groups 1 and 3. This requirement does not strictly hold, however, for the shows, museums, internet, radio and TV licenses sub-group. This is due to the fact that only two nation-wide specific price indices are available to compute the household-level price indices for the six items in the subgroup: one for telephone services (which has been imputed to Internet related services), and another for cultural services (imputed to each of the other five items). Finally, given that tax rates differ across these items, average tax rates for this sub-group have been obtained by using budget shares as weights. The resulting three expenditure sub-groups and their corresponding pre- and post-reform VAT rates are those already shown in the Introduction section, Table 1. For the sake of comparison with other European Union member countries, the reader is referred to European Commission (2012 and 2013).

In this second stage, and as already pointed out above, a Stone-Lewbel like price index is computed for each sub-group to circumvent the unit value vs. price problem, once again following Hoderlein and Mihaleva (2008). Thus, for sub-group $i \in \{1,2,3\}$ this time we construct a price index at the household level $p_{ih} \equiv (1/k_i)\times \prod_{j=1}^{J} \left(p_{i,j}/w_{i,j}^{h}\right)^{w_{i,j}^{h}}$, where J is the number of individual goods in the sub-group, $k_i \equiv \prod_{j=1}^{J} \overline{w}_{i,j}^{-\overline{w}_{i,j}}$, $\overline{w}_{i,j}$ is the mean of $w_{i,j}^{h}$, at the household-level, of the expenditure share of good j relative to total expenditure on sub-group i, and where $w_{i,j}^{h} \geq 0$, $\sum_{j=1}^{J} w_{i,j}^{h} = 1$ for all i and h, $p_{i,j}$ denoting the aggregate nation-wide price index for good j. As in the first stage, households with budget shares $w_{i,j}^{h}$ equal to 0 for *all* j are assigned a price index equal to the mean value of all the other households. This explains why means and medians of the price indices shown in Table 6 may coincide. Regarding the quantities, for any sub-group i, we obtain a quantity index (at the household level) given by $q_{ih} \equiv \sum_{j=1}^{J} w_{i,j}^{h} m_1/p_{ih}$, where m_1 is the expenditure in group 1. In this stage 3,380 households who reported a zero expenditure in the 3 sub-groups were eliminated. As a result, the sample in this second stage was reduced to 17,319 households representing a total of 13,663,984 households. Therefore, descriptive statistics of common regressors to both stages need not (and will not) coincide.

The main descriptive statistics for budget shares and independent variables are shown in Tables 5 and 6 respectively. Note that Table 5 also shows [in column 8] that there is a substantial proportion of non-consuming families in all the expenditure sub-groups, ranging from 18.88% in the case of shows, museums, internet, radio and TV licenses, to 43.96% in the case of periodicals and non-text books.

Table 4. First Stage.
Descriptive Statistics: Independent Variables

Variable	*Mean*	*Median*	*SD*	*Min*	*Max*
P_1 (Cultural goods & services)	28.738	28.609	14.559	5.531	112.868
P_2 (Food & non-alcoholic beverages)	61.746	63.473	19.170	4.132	116.484
P_3 (Housing)	83.690	78.312	29.881	25.091	306.192
P_4 (Transport & communications)	42.584	41.453	18.489	5.108	132.540
P_5 (Others)	32.969	31.772	14.684	0.850	93.386
M (Total expenditure)	29,785	25,777	18,011	1,754	410,676
$Annual\ Income$	23,643	20,400	15,351	84	199,608
mem_1 (0-4 years)	0.145	0.000	0.408	0.000	3.000
mem_2 (5-15 years)	0.284	0.000	0.609	0.000	6.000
mem_3 (16-24 years)	0.245	0.000	0.547	0.000	5.000
mem_4 (25-34 years)	0.401	0.000	0.683	0.000	5.000
mem_5 (35-64 years)	1.121	1.000	0.882	0.000	5.000
mem_6 (65-84 years)	0.408	0.000	0.695	0.000	4.000
mem_7 (85 or more years)	0.048	0.000	0.230	0.000	3.000
D_cap_1 (province capitals)	35.638	-	-	-	-
D_cap_2 (non province capitals)	64.362	-	-	-	-
D_size_1 ($\geq$ 100,000)	42.809	-	-	-	-
D_size_2 (50,000 - 100,000)	12.060	-	-	-	-
D_size_3 (20,000 - 50,000)	14.940	-	-	-	-
D_size_4 (10,000 - 20,000)	9.509	-	-	-	-
D_size_5 ($<$ 10,000)	20.683	-	-	-	-
D_region_1 (North-West)	10.129	-	-	-	-
D_region_2 (North-East)	10.507	-	-	-	-
D_region_3 (Madrid)	14.322	-	-	-	-
D_region_4 (Central)	13.058	-	-	-	-
D_region_5 (East)	31.103	-	-	-	-
D_region_6 (South)	20.881	-	-	-	-
D_gender_1 (male)	69.999	-	-	-	-
D_gender_2 (female)	30.001	-	-	-	-
D_spouse_1 (with spouse)	68.386	-	-	-	-
D_spouse_2 (without spouse)	31.614	-	-	-	-
D_ed_1 (no studies or prim. education)	21.028	-	-	-	-
D_ed_2 (sec. education, 1st cycle)	32.325	-	-	-	-
D_ed_3 (sec. education, 2nd cycle)	18.321	-	-	-	-
D_ed_4 (higher education)	28.327	-	-	-	-
D_oc_1 (occupied)	57.744	-	-	-	-
D_oc_2 (non-occupied)	42.256	-	-	-	-

Key to Table 4: main descriptive statistics for the independent variables: price indices, P_i, nominal expenditure, M, and household age composition, mem_j, shown in column 1. The table also shows the average values for qualitative locational variables (capital of province, municipality size and region) and qualitative variables corresponding to the household head in per cent terms in both cases. See Table 3 for observations and households represented.

Table 5. Second Stage.
Descriptive Statistics: Dependent Variables (w_i)

No.	*Concept*	*Mean*	*Median*	*SD*	*Min*	*Max*	*zeros*
1	Audio, video & computer equipment and musical instruments	0.300	0.147	0.343	0	1	36.850
2	Shows, museums, internet, radio & TV licenses	0.507	0.507	0.361	0	1	18.875
3	Non-text books & periodicals	0.193	0.041	0.286	0	1	43.963

No. of Observations: 17,319

Households Represented: 13,663,984

Key to Table 5: main descriptive statistics for the budget shares, w_i, of sub-groups shown in column 1. The proportion of zeros in column 8 is denoted in percent terms.

Table 6. Second Stage.
Descriptive Statistics: Independent Variables

Variable	*Mean*	*Median*	*SD*	*Min*	*Max*
p_1 (Audio, video & computer equipment, and musical instruments)	24.774	24.774	8.873	13.451	105.601
p_2 (Shows, museums, internet, radio & TV licenses)	71.193	71.193	20.588	48.662	103.633
p_3 (Non-text books & periodicals)	73.323	73.323	16.051	57.124	116.191
m_1 (Expenditure in cultural goods & services)	995	693	1,136	2	35,367
Annual Income	25,597	21,444	15,653	1,200	199,608
mem_1 (0-4 years)	0.155	0.000	0.421	0.000	3.000
mem_2 (5-15 years)	0.315	0.000	0.634	0.000	6.000
mem_3 (16-24 years)	0.279	0.000	0.576	0.000	5.000
mem_4 (25-34 years)	0.434	0.000	0.702	0.000	5.000
mem_5 (35-64 years)	1.217	1.000	0.867	0.000	5.000
mem_6 (65-84 years)	0.355	0.000	0.669	0.000	4.000
mem_7 (85 or more years)	0.037	0.000	0.205	0.000	3.000
D_cap_1 (province capitals)	37.313	-	-	-	-
D_cap_2 (non province capitals)	62.687	-	-	-	-
D_size_1 ($\geq$ 100,000)	44.829	-	-	-	-
D_size_2 (50,000 - 100,000)	12.068	-	-	-	-
D_size_3 (20,000 - 50,000)	14.846	-	-	-	-
D_size_4 (10,000 - 20,000)	9.278	-	-	-	-
D_size_5 ($<$ 10,000)	18.979	-	-	-	-
D_region_1 (North-West)	9.520	-	-	-	-
D_region_2 (North-East)	10.763	-	-	-	-
D_region_3 (Madrid)	15.926	-	-	-	-
D_region_4 (Central)	12.203	-	-	-	-
D_region_5 (East)	31.361	-	-	-	-
D_region_6 (South)	20.227				
D_gender_1 (male)	72.373	-	-	-	-
D_gender_2 (female)	27.627				
D_spouse_1 (with spouse)	72.821	-	-	-	-
D_spouse_2 (without spouse)	27.179	-	-	-	-
D_ed_1 (no studies or prim. education)	16.254	-	-	-	-
D_ed_2 (sec. education, 1st cycle)	31.717	-	-	-	-
D_ed_3 (sec. education, 2nd cycle)	19.832	-	-	-	-
D_ed_4 (higher education)	32.197	-	-	-	-
D_oc_1 (occupied)	63.617	-	-	-	-
D_oc_2 (non-occupied)	36.383	-	-	-	-

Key to Table 6: main descriptive statistics for the independent variables [prices, p_i, expenditure in group 1 (cultural goods and services, m_1) age composition, mem_j] in column 1. The table also shows the average values for qualitative locational variables (capital of province, municipality size and region) and qualitative variables corresponding to the household head, in per cent terms in both cases. See Table 5 for observations and households represented.

5 The estimation results

5.1 First stage

In our first step a probit model was estimated for the composite consumption good in group 1 (cultural goods and services), the only one with a non-negligible proportion of zero observations. The set of regressors, z_h, was defined as $z'_h = (const,\ mem_1, ..., mem_7,\ D\_cap_1,\ D\_size_1, ..., D\_size_4,\ D\_region_1, ..., D\_region_5,\ D\_gender_1,\ D\_spouse_1,\ D\_ed_2, ..., D\_ed_4,\ D\_oc_1,\ \ln P_1, ..., \ln P_5, \ln M,\ \ln^2 M)'$, where *const* is a constant term and all the variables are defined in Table 4, and $h = 1, ..., 20, 699$.[7] As each observation has a particular weight in the sample, the estimation was done using a time and space scaling factor variable.[8] Then, based on Eq. (18), the following model was estimated in the second step:

$$\begin{aligned} W_{ih} = d_{ih} \Bigg[& \alpha_i + \sum_{j=1}^{7} \tau_{ij} mem_{jh} + \mu_i D\_gender_{1h} + \phi_i D\_spouse_{1h} \\ & + \sum_{j=2}^{4} \theta_{ij} D\_ed_{jh} + \eta_i D\_oc_{1h} + \sum_{j=1}^{5} \gamma_{ij} \ln P_{jh} + \beta_i \ln \left\{ \frac{M_h}{a(\mathbf{P}_h)} \right\} \\ & + \frac{\lambda_i}{b(\mathbf{P}_h)} \left[\ln \left\{ \frac{M_h}{a(\mathbf{P}_h)} \right\} \right]^2 \Bigg] + d_{ih}\delta_i E_h + d_{ih}\rho_{i1} M_{1h} + d_{ih}\widehat{\varepsilon}_{ih}, \end{aligned} \tag{19}$$

for $i = 1, ..., 5$, and $h = 1, ..., 17, 319$, and where $d_{ih} \equiv 1$ for $i \neq 1$, $d_{1h} = 1$ if $W_{1h} > 0$, and $d_{1h} = 0$ if $W_{1h} = 0$; M_{1h} is the inverse Mills ratio, that is, $k_{1h}\phi(z'_h\widehat{\pi}_1)/\Phi(z'_h\widehat{\pi}_1)$; and E_h is the term for correcting the expenditure endogeneity. To construct this term we keep the residuals obtained by regressing households' log of total expenditure on demographic variables, log of total income and log of prices. Note that in this second step, 3,380 additional households were dropped because they reported zero consumption expenditure in cultural goods and services. As a result, the sample size for this second stage was reduced to 17,319 households representing a total of 13,663,984 households. Note also that in this step the number of regressors has been reduced. The reason for this is the differing nature of the models estimated in each step: whereas the probit model of the first step can be seen as a participation equation system, the QUAIDS model of the second step identifies which factors affect the demanded quantity of culture among those households that previously decided to participate.

Adding-up, symmetry and homogeneity conditions were imposed. In addition to those restrictions in Eq. (6), the following were also imposed: $\sum_{i=1}^{5} \tau_{ij} = 0$ for $j = 1, 2, ..., 7$, $\sum_{i=1}^{5} \mu_i = 0$, $\sum_{i=1}^{5} \phi_i = 0$, $\sum_{i=1}^{5} \theta_{ij} = 0$ for $j = 2, 3, 4$,

[7] Note that D_cap_2, D_size_5, D_region_6, D_gender_2, D_spouse_2, D_ed_1 and D_oc_2 are excluded as they represent the reference categories.

[8] Results, although not shown in the paper in order to save space, are available from the authors upon request.

$\sum_{i=1}^{5} \eta_i = 0$, $\sum_{i=1}^{5} \delta_i = 0$ and $\sum_{i=1}^{5} \rho_{i1} = 0$. Moreover, 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 4 equations and 82 parameters was then estimated. Among these parameters, 64 were statistically significant at 10% (62 of which were also significant at 5%). The estimated coefficients have no direct economic interpretation, but are necessary to compute Marshallian own- and cross-price-elasticities and expenditure elasticities, which are shown in Table 7.

Two remarks are in order. Our estimates are reasonably close to those found in the literature centered on the Spanish economy. That said, direct comparisons are neither easy nor recommended as the precise composition of expenditure groups differs across published works. Concerning own-price elasticities, all estimates are negative. The median estimate for group 1 (cultural goods and services) equals -0.41. However, although there are several references where price-elasticities are calculated for specific cultural goods and services (e.g. books, newspapers, movies...), as far as the authors know, none of them reports elasticities for cultural goods and services as a whole. Thus, and taking into account that in the second stage this group is divided into three sub-groups, the discussion about the price-elasticities is relegated to this second stage. As for food and non-alcoholic beverages (group 2), Molina (2002) and Prieto-Rodríguez *et al.* (2005) report means of -0.34 and -0.24, respectively, slightly lower than our median, -0.49. Regarding group 3 (housing), we obtain a median estimate of -1.04, substantially higher than the mean of -0.83 reported in Prieto-Rodríguez *et al.* (2005). Finally, considering group 4 (transport and communications), we obtain a median estimate of -1.09. Previous results largely vary: for instance, Molina (2002) obtains a mean estimate of -0.60; and Prieto-Rodríguez *et al.* (2005) reports mean estimates of -0.83 and -2.05 for public and private transport respectively.

Regarding expenditure elasticities, our results are also in line with previous works in the literature. The median estimate for group 1 (cultural goods and services) is 0.70, but as was the case with the price-elasticity, there is no reference dealing with cultural goods and services as a whole, so the discussion about the expenditure-elasticity of this group is postponed to the second stage. Mean expenditure elasticity estimates for food and non-alcoholic drinks (group 2) range from 0.63 in Prieto-Rodríguez *et al.* (2005) to 0.76 in Labeaga and López (1994). Interestingly enough, the latter also report a median estimate, 0.57, substantially lower than the mean, and closer to ours, 0.42. The median expenditure elasticity of housing (group 3), 1.08, falls between the mean estimate in Prieto-Rodríguez *et al.* (2005), 0.79, and the median in Labeaga and López (1994), 1.29. Finally, the median expenditure elasticity of group 4 (transport and communications), 1.19, is in line with elasticity estimates

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

reported for the transport expenditure group. For instance, Prieto-Rodríguez *et al.* (2005) report a mean of 1.20 for private transport and a mean of 1.01 for public transport, and Molina (1997) obtained a mean of 1.70 for total transport and a mean of 1.13 for public transport.

Table 7. First-Stage Marshallian and expenditure elasticities

	j				
Elasticities	*1*	*2*	*3*	*4*	*5*
$E^U_{1,j}$ (Cultural goods and services)	-0.41	-0.14	0.02	-0.12	-0.06
	(2.84)	(0.66)	(0.28)	(0.55)	(0.31)
$E^U_{2,j}$ (Food and non-alcoholic beverages)	-0.01	-0.49	0.18	-0.08	-0.01
	(0.05)	(2.40)	(1.10)	(0.30)	(0.19)
$E^U_{3,j}$ (Housing)	-0.01	-0.06	-1.04	0.03	0.01
	(0.01)	(0.03)	(0.07)	(0.02)	(0.01)
$E^U_{4,j}$ (Transport and communications)	-0.04	-0.19	0.01	-1.09	0.10
	(0.16)	(0.87)	(0.66)	(0.39)	(0.59)
$E^U_{5,j}$ (Others)	-0.02	-0.09	-0.04	0.04	-1.03
	(0.03)	(0.17)	(0.07)	(0.09)	(0.03)
E^M_j	0.70	0.42	1.08	1.19	1.13
	(1.47)	(3.31)	(0.09)	(0.67)	(0.18)

No. of Observations: 17,227

Households Represented: 13,585,240

Key to Table 7: $E^U_{i,j}$ denotes Marshallian price elasticity of the i-th expenditure group with respect to the j-th price [see Eq.(8)]; E^M_j denotes expenditure elasticity of good j [see Eq. (7)]. In order to avoid the extreme values effect, the table reports the median elasticities across households rather than the more often used means. Numbers in parentheses denote standard deviations.

5.2 Second stage

Following the same procedure as in the first-stage estimation, the estimation strategy consists of two steps. In the first step the observed version of Eq. (12) was estimated as a multivariate probit using *mvprobit*, a simulated maximum likelihood estimator included in Stata©. The system consisted of 3 equations, one for each expenditure sub-group, and the set of regressors, z_h, was defined as $z'_h = (const,\ mem_1, ...,\ mem_7,\ D\_cap_1,\ D\_size_1, ...,\ D\_size_4,\ D\_region_1,\ ...,\ D\_region_5,\ D\_gender_1,\ D\_spouse_1,\ D\_ed_2, ...,\ D\_ed_4,\ D\_oc_1,\ \ln p_1, ..., \ln p_3,\ \ln m_1,\ \ln^2 m_1)'$, where $\ln p_j$ stands for the log-price index of the j-th consumption good, and m_1 denotes expenditure on composite good 1, for $h = 1, ..., 17,319$, whereas all the other variables are the same as in the first stage [see Table 6].[10] Once again, given that each observation has a particu-

[10] Note that D_cap_2, D_size_5, D_region_6, D_gender_2, D_spouse_2, D_ed_1 and D_oc_2 are excluded as they represent the reference categories.

lar weight in the sample, estimation was done using a time and space scaling factor variable.[11]

Using the multivariate probit results, the correction terms M_{jh} in Eq. (15) were constructed and, based on Eq. (17), the following model was estimated in the second stage:

$$w_{ih} = d_{ih} \left[\alpha_i + \sum_{j=1}^{7} \tau_{ij} mem_{jh} + \sum_{j=1}^{3} \gamma_{ij} \ln p_{jh} + \beta_i \ln \left\{ \frac{m_{1h}}{a(\mathbf{p}_h)} \right\} + \frac{\lambda_i}{b(\mathbf{p}_h)} \left[\ln \left\{ \frac{m_{1h}}{a(\mathbf{p}_h)} \right\} \right]^2 \right] + d_{ih}\delta_i e_h + d_{ih} \sum_{j=1}^{3} \rho_{ij} M_{jh} + d_{ih}\widetilde{\varepsilon}_{ih}, \tag{20}$$

for $i = 1, 2, 3$, and $h = 1, ..., 5, 361$, and where e_h is the term for correcting the expenditure endogeneity. To construct this term we keep the residuals obtained by regressing households' log of total expenditure in cultural goods and services on all the demographic variables, log of total income and log of prices. Note that in this second step, 11,866 additional households were dropped because those households reported zero consumption in at least one of the 3 expenditure categories considered. As a result, the sample size for this second step was reduced to 5,361 households representing a total of 4,383,195 households. Additionally, comparing the estimated QUAIDS model in (20) to the one in (19) it is observed that some demographic variables $(D\_gender_1, D\_spouse_1, D\_ed_2, ..., D\_ed_4, D\_oc_1)$ were dropped. The reason was the lack of significance at 5%.

Ours is a static demand model in the sense that past consumption levels do not affect the utility of current consumption. There is a vast literature, however, that explains the demand for culture by means of the so-called rational addiction models. Thus, past consumption of cultural goods gives rise to a cultural capital stock which positively affects the utility derived from current cultural consumption, the idea dating back to Stigler and Becker (1977). As pointed out by Seaman (2006), even habit formation must be distinguished from learning-by-consuming and rational addiction in examining dynamic determinants [see Ateca-Amestoy (2007) and references therein]. The SHBS database that we use, however, does not provide any variable that could be used even as a proxy for acquired consumption habit.

Along these lines, it should also be noted that each year half of the SHBS sample is renewed, so that only half of households can be followed, which would further greatly reduce the sample size if one were tempted to consider a first-difference estimation procedure. But, in this case, an additional problem would rise from the fact that the VAT reform was introduced on September 1, 2012. This means that some households would have been interviewed before the price increase and others afterwards. And if we wanted the sample to extend backwards, we would find a similar problem as another partial VAT

[11] As in the first stage, results are not shown in the paper in order to save space, but are available from the authors upon request.

reform came into force on July 1, 2010. Considering all these limitations, it seems appropriate not to introduce habits in the model.

The adding-up, symmetry and homogeneity conditions were imposed as in the first stage. In addition to those in Eq. (6), the following restrictions were also imposed: $\sum_{i=1}^{3} \tau_{ij} = 0$ for $j = 1, 2, ..., 7$, $\sum_{i=1}^{3} \delta_i = 0$ and $\sum_{i=1}^{3} \rho_{ij} = 0$ for $j = 1, 2, 3$. As in the first stage, the last equation was dropped in order to avoid the singularity of the variance and covariance matrix of the perturbations, then obtaining the parameter estimates for this equation using the restrictions imposed.[12] The result is a system of 2 equations and 31 free parameters. Among these parameters, 18 were statistically significant at 10% (13 of which were also significant at 5%). Once again, the estimated coefficients have no direct economic interpretation, but are necessary to compute Marshallian own- and cross-price-elasticities and expenditure elasticities.

As in the first stage, in this second stage partial (conditional) elasticities were computed for each household. As discussed in Section 2, however, the formulae for conditional expenditure and price elasticities in equations (7) and (8) cannot be directly applied as one needs *total* or *unconditional* elasticities (*i.e.* modified for the first-stage estimation). Thus, Table 8 shows the medians of the total elasticities.

The available alternatives to compare our results with are very few and, as in the first stage, direct comparisons are neither easy nor recommended as the precise composition of expenditure sub-groups differs across published works. Regarding sub-group 1 (audio, video & computer equipment and musical instruments), the only reference is Prieto-Rodríguez *et al.* (2005). These authors obtain a mean estimate of own-price elasticity 0.65 for music and film on magnetic media, very far from our median estimate, -0.46. Note, however, that if we reported the mean estimate (non representative in this case and, therefore, inappropriate), we would obtain a value of 0.11, much lower than 0.65, although with the same sign. As for sub-group 2 (shows, museums, internet, radio & TV licenses), whereas our median estimate of the own-price elasticity is -0.12, Prieto-Rodríguez *et al.* (2005) report a mean estimate of -1.23 for cinema, theater, museum and other shows. More recently Fernández-Blanco *et al.* (2013) report a lower mean estimate (-1.07) for the Spanish movie market, but still bigger than ours in absolute terms. Finally, concerning sub-group 3 (non-text books and periodicals), we obtain an own-price elasticity of -0.17, while the results in previous works are somewhat mixed. Thus, while Prieto *et al.* (2005) obtain a mean estimate of -1.65, *i.e.* an elastic demand, Palma-Martos *et al.* (2009) report mean estimates between -0.75 and -0.61 for the Spanish book market, both figures representing inelastic demand, and Jaén-García (2012) finds that the demand has unit elasticity with respect to price. Obtaining such inelastic demand estimates might appear somewhat strange. Results like this, however, are quite common in the literature [see, for example, Frey and Pommerenhe (1989) and Bonato *et al.* (1990) for the case

[12] After computing the median of the Hicksian substitution term matrices for all the households in the sample, it emerged that two out of the three eigenvalues were negative, the other one being positive but close to 0: 0.01.

of performing arts; Gapinski (1984) for the theater; or Lange and Luksetich (1984) for the classic music concerts]. Furthermore, as Seaman (2006) highlights for the case of performing arts "regardless of technical sophistication, the price inelasticity result is much more prominent in those studies that used very aggregative data...", the absence of substitutes being the main reason.

Regarding cross-price elasticities, both positive and negative signs are observed, so expenditure sub-groups can be substitutes or complements respectively. Sub-group 1 (audio, video & computer equipment and musical instruments), sub-group 2 (shows, museums, internet, radio & TV licenses) and sub-group 3 (non-text books & periodicals) can be considered as complements, the former relationship being weaker than the latter (elasticities are -0.10 and -0.09, noticeably smaller than -0.17 and -0.46). As for the relationship between sub-group 1 (audio, video & computer equipment and musical instruments) and 3 (non-text books and periodicals), we find a negligible substitution relationship, with values near zero (0.01 and 0.03), so in practice these sub-groups could be considered as independent. Prieto *et al.* (2005) also find that cinema, theater & shows and records & films are substitutes (with mean estimates of 0.711 and 1.329), whereas the relationships between each of these sub-groups and books, newspapers and magazines are negative (between -0.33 and -0.84), as in the case of the complements.

Finally, looking at the expenditure elasticities, the median estimates are always positive, ranging from 0.67 (shows, museums, internet, radio & TV licenses) to 1.03 (non-text books & periodicals), so that it can be concluded that sub-group 3 (non-text books & periodicals) represent a luxury good, whereas the other sub-groups, although they can be considered as necessary goods, are not so far from being luxuries. Prieto *et al.* (2005), where mean estimates are reported, find that all the sub-groups are luxury goods. In particular, they report values of 1.24, 1.75 and 1.37 for records and films, cinema, theater and shows and non-text books and periodicals, respectively. Later, for the case of books, Palma-Martos (2009) calculate a mean estimate of 1.10, whereas Jaén-García (2012) finds a value of 0.8, both of them more in line with our results.[13]

As an overall conclusion it can be said that, in general, our elasticities are smaller in absolute terms than those reported by the other studies. In this connection, apart from the aggregation issue, it is worth remembering that the use of unconditional elasticities (and not the conditional ones as in all the above mentioned papers) affects their values.

[13] For the sake of precision, Palma-Martos (2009) and Jaén-García (2012) refer to income elasticity.

Table 8. Total Marshallian and expenditure elasticities.

	j		
Elasticities	*1*	*2*	*3*
$E^u_{1,j}$ (Audio, video & computer equipment, and musical instruments)	-0.46	-0.10	0.01
	(2.05)	(2.11)	(0.72)
$E^u_{2,j}$ (Shows, museums, internet, radio & TV licenses)	-0.09	-0.12	-0.17
	(0.63)	(1.74)	(0.82)
$E^u_{3,j}$ (Non-text books & periodicals)	0.03	-0.46	-0.17
	(1.42)	(3.84)	(3.40)
E^M_j	0.95	0.67	1.03
	(0.70)	(0.53)	(1.60)

No. of Observations: 5,361

Households Represented: 4,383,195

Key to Table 8: $E^u_{i,j}$ denotes Marshallian price elasticity of the i-th expenditure sub-group with respect to the j-th price and E^M_j denotes expenditure elasticity of sub-group j [see Eqs. (10) and (9)]. In order to avoid the extreme values effect, the table reports the median elasticities across households. Numbers in parenthesis denote standard deviations.

6 Distributive and welfare impact of the tax reform

A key issue when studying the consequences of a tax policy change is that of the tax incidence or, in the specific case of a change in VAT rates, how VAT tax rate changes are split up between consumers and sellers. In other words, how the consumers' and the producer's prices are affected by the change in the tax rates. An frequent assumption when studying issues like this one is that producers' prices remain tax invariant, so that tax changes are completely shifted to buyers' prices [see, e.g. Labeaga and López, (1994), Prieto-Rodríguez *et al.* (2005) or García-Enríquez and Echevarría (2016)]. Economic theory and empirical evidence, however, cast doubt on this assumption. For instance, Besley and Rosen (1999), despite focusing on the particular problem of the incidence of sales taxes on specific commodities in the U.S., finds a "*surprising variety of shifting patterns*".[14]

Focusing on the European economy, Benedek *et al.* (2015) estimates the pass-through of VAT changes to consumer prices for 17 Eurozone countries over 1999-2013, their main findings following: *i*) pass-through is much less than full on average, and differs markedly across types of VAT change. *ii*) Changes in the standard rate lead to a pass-through rate of about 100%. *iii*) Changes in the reduced rates imply a lower pass-through rate, around 30%. *iv*) Reclassifications (movements of some item between rate categories) give rise to an essentially zero pass-through rate. *v*) Pass-through for durables is greater than for non-durables. *vi*) There is no significant difference in pass-through between rate increases and decreases. *vii*) The pass-through dynamics

[14] The reader is referred to Section 8 of the thorough survey in Adam *et al.* (2011) for further details.

depends on the reform type: the largest part of the pass-through of standard rate changes occurs in the months before the actual VAT change (*i.e.* the anticipation effect), in particular in the case of durables. For reduced VAT rates, however, the anticipation effect is weaker. And *viii*) indications of significant anticipation effects together with some evidence of lagged effects in the two years around reform.

Formally, following Benedek *et al.* (2015), the pass-through parameter for the i-th good can be defined as

$$\gamma_i \equiv \frac{(p_i^1 - p_i^0)/p_i^0}{(t_i^1 - t_i^0)/(1 + t_i^0)}, \tag{21}$$

where p_i^0 denotes the i-th element of pre-reform *consumer* price vector, $\mathbf{p}^0$, p_i^1 denotes the i-th element of post-reform *consumer* price vector, $\mathbf{p}^1$, t_i^0 and t_i^1 denote the corresponding pre- and post-reform tax rates, respectively, for good i. From Eq. (21) one obtains that the post-reform consumer price of the i-th good can be rewritten as a function of both the pre- and post-reform tax rates, the pre-reform consumer price and, of course, the pass through parameter γ as

$$p_i^1 = p_i^0 \left[\frac{\gamma_i (t_i^1 - t_i^0)}{1 + t_i^0} + 1 \right], \tag{22}$$

expression that will be used below. Thus, if say the VAT is completely shifted to the consumers (producer prices staying constant), so that $\gamma_i = 1$, one will obtain that $p_i^1 = p_i^0 \left(1 + t_i^1\right)/(1 + t_i^0)$, the standard assumption referred to above. If, conversely, the VAT were completely borne by the producer, *i.e.* $\gamma_i = 0$, one would obtain that $p_i^1 = p_i^0$ (*i.e.* constant or tax invariant consumer prices).

As an approximation to the specific pass-through parameter for the 2012 VAT reform in Spain for cultural expenditure, we set $\gamma = 31.41\%$. We have obtained this figure from Eq. (21) by replacing p_i^1, p_i^0, t_i^1, and t_i^0 with the average price index for cultural items in 2013 and 2011, published by the Spanish Statistical Office, (101.418 and 100 respectively) and the average VAT rates for cultural items in 2013 and 2011 (18.14% and 13.04% respectively) which we have computed, using budget shares to obtain these.[15] Our estimate for γ falls within the range of other estimates found in previous works. Bank of Spain (2012), p. 46, reports that "the estimated pass-through rate [for the Spanish VAT reform in 2012] would be around one third of the total impact potential". Benedek *et al.* (2015), Table 2, p. 14, estimates that, for the 17 Eurozone countries and the period 1999-2013, the pass-trough parameter ranges between 40.0% and 29.0% depending on the econometric specification. Note, however, that the estimates in the last two cases refer to average values obtained for the set of *all* sectors in the economy, *i.e.* not only cultural goods and services.

[15] As a first approximation, we assume the same transfer parameter for the three cultural items under study.

6.1 Welfare analysis: results

In what follows, we show the results that we have obtained concerning the welfare implications for the VAT policy reform focusing exclusively on cultural items. The reader is referred to the formal Appendix at the end of the paper where standard theoretical concepts are formally defined. The results shown, namely, the post-reform tax bill and the equivalent and compensating variations [see Eqs. (A.2) and (A.4)], have been obtained for those households with *positive* consumption of the three cultural sub-group items. This amounts to a sample of $5,361$ observations representing $4,383,195$ households, for whom the median and the average shares on cultural goods and services (conditional on positive spending) are 3.24% and 3.97% respectively.

The aggregate effects of the VAT reform are shown in Table 9. Focusing, first, on the household's VAT tax bill on cultural goods and services, the VAT reform implies an estimated median increment of 32.81 from the initial (*i.e.* pre-reform) tax bill of 125.93, thereby making a final of €164.84. [See last row, columns 3, 4 and 5.][16] The equivalent and the compensating variations display negative signs as expected. For instance, focusing on the median equivalent variation, this equals -13.43 euros [see last row, column 6]. Since the reform studied here is not revenue neutral, all households lose after the increase in VAT rates.[17] For completeness, the table also shows the households' median expenditure on culture, €$1,152.98$ [see last row, column 2].

From an efficiency standpoint, more relevant measures of the individual welfare change induced by the reform should jointly consider, first, the equivalent or the compensating variations and, second, the change in tax revenues. In this way, previous results would allow us to compute, for instance, two alternative measures of the deadweight loss or excess burden for household h: one obtained after the equivalent variation, the other based on the compensating variation. Namely, $EB_h^{ev} \equiv EV_h + (R_h^1 - R_h^0)$ and $EB_h^{cv} \equiv CV_h + (R_h^1 - R_h^0)$, where R_h^0 and R_h^1 denote the pre-reform and post-reform household's tax bill respectively. Note, however, that this would make sense only under the assumption that the tax rate change is totally shifted to consumers [in other words, if the parameter γ in Eq. (21) equals 1]. See, for instance, Prieto-Rodríguez *et al.* (2005). Otherwise, a third component should also be considered to com-

[16] This represents a 30.90% increment. To place this figure in context two remarks are in order. First, at least to the best of our knowledge, there are no published data for VAT revenues on, specifically, cultural goods and services. The increment in the *total* VAT real revenue between 2011 and 2013, however, was 8.5% [see *Estadística del año 2013*, available at *http://www.agenciatributaria.es/AEAT/Contenidos_Comunes/La_Agencia_Tributaria/Estadisticas/Publicaciones/sites/iva/2013*]. And, second, one must bear in mind that tax proceeds depend both on the tax policy and the level of economic activity. Along these lines, Spanish real GDP in 2013 was 3.3% lower than in 2011 [see *WEO Data Base, April 2015*, available at *https://www.imf.org/external/pubs/ft/weo/2015*].

[17] Complementary to the individual welfare change measures shown above, we also compute the final and initial equivalent expenditures [Eqs. (A.3) and (A.5)]. As explained in the Appendix, the measures EE^F and EE^I are closely related to EV and CV, respectively. Since EE^F and EE^I enter into the definitions of the social welfare measures shown in Table 11, we report figures for both EV, CV, EE^F and EE^I.

pute the deadweight loss, namely, the change in the producers' surplus. The lack of even an approximated measure of this magnitude prevents any serious attempt to proceed beyond the effects on tax revenues and consumers' welfare change.[18]

Table 9 also shows the *distribution* of the above aggregate effects across households' income levels. As one would expect, they all increase across income levels. For instance, the median increment of the tax bill for the uppermost income decile is almost 4 times the increment for the lowermost income decile [See, column 5]. And, along these same lines, focusing for example on the equivalent variation, one observes that households in the uppermost income decile experienced a welfare loss almost 4 times higher than households in the lowest income decile [see column 6].

TABLE 9. DISTRIBUTIVE ANALYSIS OF WELFARE. ($\hat{\gamma} = 31.41\%$)

Income decile	*Exp*	R^0	R^1	ΔR	EV	CV	EE^I	EE^F
1: 13,632	519.04	58.45	75.02	15.18	−6.07	−6.20	526.15	512.01
2: 15,540	788.68	85.85	115.48	23.68	−9.33	−9.44	793.46	779.60
3: 20,424	960.54	105.77	141.38	28.05	−10.80	−10.89	972.49	944.87
4: 24,480	1,099.46	127.22	164.58	32.05	−12.52	−12.71	1105.94	1088.79
5: 26,988	1,130.99	121.93	160.39	35.64	−14.39	−14.61	1151.33	1117.37
6: 32,340	1,278.45	135.97	187.62	38.47	−16.10	−16.23	1290.25	1269.00
7: 36,000	1,210.43	131.37	172.21	35.40	−14.29	−14.41	1226.85	1198.66
8: 41,532	1,364.72	144.91	195.83	41.89	−17.37	−17.58	1380.69	1348.34
9: 52,116	1,615.84	171.71	228.35	46.99	−20.12	−20.42	1637.56	1600.47
10:199,608	1,808.29	185.78	248.05	57.63	−23.87	−24.19	1835.94	1781.01
Median	1,152.98	125.93	164.84	32.81	−13.43	−13.60	1166.93	1137.09

Key to Table 9. Breakdown of household's tax and welfare change by income deciles. Exp: nominal expenditure in culture; R^0: pre-reform tax revenue; R^1: post-reform tax revenue; ΔR: change in tax revenue; EV: equivalent variation; CV: compensating variation; EE^I: equivalent initial expenditure; EE^F: equivalent final expenditure. All figures are measured at their median values in 2011 euros.

Figures in Table 9 are expressed in 2011 euros, *i.e.* in absolute terms. This raises the need to consider figures in relative terms: Table 10 serves this end, where figures are expressed in per thousand terms. As a norm, higher levels of income are associated with higher levels of expenditure on culture, but this kind of expenditure represents a higher proportion of the annual income for lower income brackets [compare figures in Table 9, column 2, with those in Table 10, column 2]. Following the previous discussion, it is found that the higher the income level, the lower the proportion of total income represented by both the pre- and post-reform tax revenues and also the increase in tax

[18] According to our estimate of the pass-through paramater, producers would have absorbed 68.59% of the average rate increase.

revenues [see column 5]. A similar pattern is also obtained for our measures of the consumers' welfare loss. For instance, lower levels of annual income are associated with higher (relative) equivalent variation levels (with sporadic, negligible exceptions) [see column 6]. The conclusion is clear: considering only its impact on the cultural commodity set, *i*) the 2012 Spanish VAT reform can be labeled regressive, but *ii*) the effects are small, as they all represent low fractions of households' incomes.

Table 10. Distributive analysis of welfare relative to annual income ($\hat{\gamma} = 31.41\%$)

Income decile	*Exp*	R^0	R^1	ΔR	EV	CV	EE^I	EE^F
1: 13,632	56.88	6.23	8.28	1.67	−0.70	−0.71	57.49	56.07
2: 15,540	54.26	5.83	7.97	1.65	−0.64	−0.64	54.87	53.78
3: 20,424	52.04	5.80	7.58	1.54	−0.58	−0.59	52.34	51.50
4: 24,480	50.59	5.94	7.54	1.50	−0.58	−0.59	51.49	49.70
5: 26,988	42.80	4.64	6.02	1.35	−0.55	−0.56	43.53	42.25
6: 32,340	40.88	4.63	6.14	1.25	−0.53	−0.54	41.51	40.55
7: 36,000	36.43	3.95	5.19	1.06	−0.43	−0.44	36.89	36.00
8: 41,532	35.30	3.70	5.03	1.08	−0.45	−0.45	35.75	34.91
9: 52,116	35.08	3.68	5.11	1.01	−0.43	−0.43	35.48	34.61
10:199,608	26.60	2.84	3.72	0.80	−0.33	−0.34	26.97	26.32
Median	41.29	4.47	5.93	1.25	−0.51	−0.51	41.83	40.65

Key to Table 10. Breakdown of household's tax revenues and welfare change by income deciles. See key to Table 9. All figures (except column 1) are expressed in per thousand terms relative to annual income.

Next we discuss social welfare measures of the impact of the VAT rates on cultural expenditures. Results for King's proportional increase in equivalent income (λ) are shown for alternative values of the inequality aversion parameter (ε) in Table 11. The conclusion seems patent: regardless of the inequality parameter, the VAT reform induced a loss in social welfare: in all cases $\lambda < 1$. Additionally, note that (as expected) the increasing pattern between ε and λ: higher aversion to inequality induces higher welfare losses [see Table 11, column 5]. Along the same lines, higher values for the aversion to the inequality in the distribution of the cultural expenditure among Spanish households lead to higher values of the corresponding Atkinson's inequality indices: both for pre-reform expenditure, $A^0(\varepsilon)$, see Table 11, col. 2, and the two post-reform equivalent expenditures, $A^I(\varepsilon)$ and $A^F(\varepsilon)$ [see Table 11, cols. 3-4]. Note also that in all cases (*i.e.* for all ε), $A^I(\varepsilon) < A^0(\varepsilon) < A^F(\varepsilon)$: depending on equivalent expenditure considered (initial or final), the tax reform would have reduced or increased the inequality in the distribution of cultural spending. This leads us to conclude that *i*) the tax reform at issue would have caused a lower social loss the higher the inequality aversion of Spanish society as a whole, and *ii*) it is unclear whether the VAT reform would have diminished or raised the cultural spending distribution inequality among Spanish households.

Table 11. King's proportional increase in initial equivalent income ($\hat{\gamma} = 31.41\%$)

Inequality Aversion, ε	Atkinson's $A^0(\varepsilon)$ index	Atkinson's $A^I(\varepsilon)$ index	Atkinson's $A^F(\varepsilon)$ index	*King's* $\lambda(\varepsilon)$
0	0	0	0	0.97416
0.5	0.14614	0.14602	0.14627	0.97387
1.0	0.28397	0.28356	0.28439	0.97304
1.5	0.42118	0.42020	0.42219	0.97082
2.0	0.55595	0.55420	0.55773	0.96644
2.5	0.67398	0.67171	0.67627	0.96062
3.0	0.76186	0.75955	0.76417	0.95547
3.5	0.82055	0.81848	0.82262	0.95196
4.0	0.85884	0.85704	0.86063	0.94973

Key to Table 11. ε: Inequality aversion index [see Eq. (A.6)]; $A^I(\varepsilon)$ and $A^F(\varepsilon)$ Atkinson's indices for initial and final equivalent expenditures [see Eq.(A.10)]; $\lambda(\varepsilon)$: King's index [see Eq. (A.11)]; $A^0(\varepsilon)$: pre-reform Atkinson's index.

One might argue that a key issue in the previous results, at least from a quantitative viewpoint, is that of the estimated value for the pass-through parameter. To put it in another way, the following question predictably arises: are the results shown above robust to alternative values for $\hat{\gamma}$? We have considered the extreme case of $\hat{\gamma} = 100.00\%$ (*i.e.*, VAT tax rate changes are completely shifted to consumers, so that producers' prices stay constant; as noted above, the usual assumption in this kind work).[19] The reader is referred to the Appendix at the end of the paper for details. The conclusion is clear. When considering the effects of the VAT reform, its regressive nature is confirmed. As expected, the results quantitatively depend on the assumption made for the pass-through parameter, γ, but not from a qualitative point of view: higher values for γ lead to stronger effects of the reform. All in all, the quantitative relevance of the welfare impact of the VAT reform is arguable: when relating the effects to households' incomes, magnitudes are not sizable simply because cultural expenditure in Spanish households represents low proportion of total expenditure.

7 Conclusions

The Spanish Government raised VAT rates in 2012 with the purpose of achieving budgetary stability, which could be expected to affect households' welfare. In particular, the VAT tax rates on cultural goods and services suffered a sizable increment: the (weighted) average tax rate was increased from 13.04%

[19] The results for other (intermediate) cases, namely $\hat{\gamma} = 66\%$ and $\hat{\gamma} = 33\%$, are not shown in the paper for the sake of space saving, but are available from the authors upon request.

in 2011 to 18.14% in 2013. In order to assess to what extent households and society in general have been affected by this reform, we have estimated a two-stage QUAIDS demand model, paying special attention to the treatment of zero expenditure observations.

Estimated expenditure and price elasticities show that all the expenditure groups and sub-groups can be considered as normal. Making use of these elasticities we estimate the effects suffered by Spanish households across income levels in terms of tax revenues, equivalent and compensating variations, and equivalent and compensating initial and final expenditures.

Having estimated a pass-through parameter for the VAT on cultural items of 31.41%, we have found that the sizes of all these effects in absolute terms increase with income. Additionally, higher levels of income are also associated with higher levels of expenditure on cultural items, but this expenditure represents a higher proportion of annual income for lower income levels with, consequently, greater impact on household budgets. In short, the 2012 VAT reform in Spain, can be considered regressive. Additionally, the tax reform at issue would have caused a social loss whose magnitude would depend on the level of the inequality aversion regarding the distribution of Spanish households' cultural expenditure: higher inequality aversion levels would imply higher welfare losses. Along these lines, regardless of such inequality aversion, it cannot be determined whether the VAT reform would have diminished or raised inequality in the distribution of cultural spending among Spanish households when compared with the pre-reform distribution.

When relating the VAT reform effects to households' incomes, the relative effects of the reform do not seem sizable, as cultural expenditure in Spanish households represents a quite low proportion of total expenditure: a mean of 2.63% and a median of 1.96%.

Finally, as a means of checking the robustness of our results, we have considered alternative values for the pass-through parameter, concluding that it only plays a negligible quantitative role, as the magnitudes of the effects fall for lower values of the pass-through parameter, thereby corroborating this general conclusion.

Appendix

Individual and social welfare: theory

We follow the methods of Urzúa (2001) and first established in King (1983b). We first obtain the tax revenues of the h-th household before and (expected) after the tax reform, R_h^0 and R_h^1, respectively. Thus, one trivially has that $R_h^0 = \sum_{i=1}^n t_i^0 p_{i,h}^0 q_{i,h}^0/(1+t_i^0)$ and $R_h^1 = \sum_{i=1}^n t_i^1 p_{i,h}^1 q_{i,h}^1/(1+t_i^1)$, 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. We next obtain $q_{i,h}^1$. Assuming a Marshallian demand function $q_{i,h} = q_{i,h}(p_{1,h}, p_{2,h}, ...p_{n,h}, m_h)$, where m_h denotes total expenditure on cultural goods and services, totally differentiating both sides, assuming further as usual that the m_h were invariant to the VAT reform, approximating $dp_{j,h} \approx p^1_{j,h} - p^0_{j,h}$, using Eq. (22) and, finally, approximating $q^1_{i,h} \approx q^0_{i,h} + dq_{i,h}$, it can be shown that

$$q^1_{i,h} = q^0_{i,h} \left[1 + \sum_{j=1}^{n} E^u_{i,j} \frac{(t^1_j - t^0_j) \times \gamma_j}{1 + t^0_j} \right], \tag{A.1}$$

for $i = 1, 2, ..., n$, where $E^u_{i,j}$ denotes the uncompensated (Marshallian) demand cross-price elasticity of good i with respect to price j, which was obtained in Eq. (10). Note that $q^1_{i,h}$ depends on the pass-through parameter, γ_j, and so does the (expected) after-tax reform, R^1_h, above defined.

Once price effects of changing the tax rates have been computed, the individual welfare change arising from the tax reform for each household 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, p^0_h, to make it as well off as it would be facing the new price vector, p^1_h, with its initial cultural expenditure, m_h [Gravelle and Rees (2004)]. More formally, upon denoting this household's indirect utility function by V, EV_h is implicitly defined by $V(m_h + EV_h, p^0_h) \equiv V(m_h, p^1_h)$, so that $EV_h < 0$ for $p^1_h \geq p^0_h$, $p^1_h \neq p^0_h$. Thus, from Eqs. (1)-(4) one can explicitly solve for EV_h as

$$EV_h = a(\mathbf{p}^0_h) \times \exp \left\{ \frac{b(\mathbf{p}^0_h) \times \ln \left[m_h / a(\mathbf{p}^1_h) \right]}{b(\mathbf{p}^1_h) + [\lambda(\mathbf{p}^1_h) - \lambda(\mathbf{p}^0_h)] \times \ln \left[m_h / a(\mathbf{p}^1_h) \right]} \right\} - m_h. \tag{A.2}$$

As a closely related concept, one could also define the final equivalent expenditure, EE^F_h, as

$$V(EE^F_h, \mathbf{p}^0_h) \equiv V(m_h, \mathbf{p}^1_h), \tag{A.3}$$

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

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 cultural expenditure, m_h, when facing the new price vector, p^1_h, in order to make it as well off as it was when it faced the old price vector, p^0_h [Gravelle and Rees (2004)]. In other words, CV_h is implicitly defined as $V(m_h - CV_h, p^1_h) \equiv V(m_h, p^0_h)$, so that $CV_h < 0$ for $p^1_h \geq p^0_h$, $p^1_h \neq p^0_h$. From Eqs. (1)-(4) one has that CV_h is explicitly solved for as

$$CV_h = m_h - a(\mathbf{p}^1_h) \times \exp \left\{ \frac{b(\mathbf{p}^1_h) \times \ln \left[m_h / a(\mathbf{p}^0_h) \right]}{b(\mathbf{p}^0_h) + [\lambda(\mathbf{p}^0_h) - \lambda(\mathbf{p}^1_h)] \times \ln \left[m_h / a(\mathbf{p}^0_h) \right]} \right\}. \tag{A.4}$$

As was the case with the equivalent variation, the compensating variation allows one to define the initial equivalent expenditure, EE^I_h, as

$$V(EE^I_h, \mathbf{p}^1_h) \equiv V(m_h, \mathbf{p}^0_h), \tag{A.5}$$

(i.e. the expenditure required at post-reform prices to attain the same level of utility as with pre-reform prices) so that $EE_h^I \equiv m_h - CV_h > m_h$ for $p_h^1 \geq p_h^0$, $p_h^1 \neq p_h^0$. 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 be easily computed after Eqs. (A.2) and (A.4) respectively. A welfare-enhancing reform (i.e. EV_h, $CV_h > 0$) will imply that $EE_h^I < m_h < EE_h^F$. And, similarly, a reducing welfare reform (i.e. EV_h, $CV_h < 0$) will imply that $EE_h^I > m_h > EE_h^F$. Note that, by construction, $EE_h^F - EE_h^I \equiv EV_h + CV_h$, which, in case of a welfare loss (gain) as a result of the tax policy, will be negative (positive).

As a complement to these individual welfare change measures, we also consider the welfare effects from a social point of view, i.e. 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 $\widehat{EE} = [EE_1, EE_2, ..., EE_H]$ given by

$$W(\cdot) = \begin{cases} \frac{1}{H}\sum_{h=1}^{H} \frac{EE_h^{1-\varepsilon}}{1-\varepsilon}, & \text{for } \varepsilon \neq 1 \\ \\ \frac{1}{H}\sum_{h=1}^{H} \ln EE_h, & \text{for } \varepsilon = 1 \end{cases}, \tag{A.6}$$

where parameter ε captures the degree of aversion to social inequality [see Atkinson (1970)]. From Eq. (A.6) one can derive a measure of social value, the proportional increment in initial equivalent expenditure $\widehat{EE}^I = [EE_1^I, EE_2^I, ..., EE_H^I]$, which we denote by λ, and 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 $\widehat{EE}^F = [EE_1^F, EE_2^F, ..., EE_H^F]$. Or, more formally,

$$\begin{aligned} W(\lambda \times EE_1^I, \lambda \times EE_2^I, ..., \lambda \times EE_H^I) = \\ W(EE_1^F, EE_2^F, ..., EE_H^F), \end{aligned} \tag{A.7}$$

so that, given that $EE_h^I > EE_h^F$ as a result of $p_h^1 \geq p_h^0$ and $p_h^1 \neq p_h^0$, 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 for 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

$$W(G^I, G^I, ..., G^I) \equiv W(EE_1^I, EE_2^I, ..., EE_H^I) \tag{A.8}$$

and

$$W(G^F, G^F, ..., G^F) \equiv W(EE_1^F, EE_2^F, ..., EE_H^F). \tag{A.9}$$

Inequality indices can be easily computed as follows. Denote the average initial and final equivalent expenditures as $\overline{EE^I} \equiv H^{-1}\sum_{h=1}^{H} EE_h^I$ and $\overline{EE^F} \equiv H^{-1}\sum_{h=1}^{H} EE_h^F$ respectively. If W is concave (therefore denoting inequality aversion), then $G^I \leq \overline{EE^I}$ and $G^F \leq \overline{EE^F}$. Under the assumption that $W(\cdot)$ is symmetrical and concave, previous definitions provide two inequality indices (one for the initial equivalent expenditure, the other for the final equivalent expenditure):

$$A^I \equiv 1 - \frac{G^I}{\overline{EE^I}}, \text{ and } \; A^F \equiv 1 - \frac{G^F}{\overline{EE^F}}, \tag{A.10}$$

where $A^I, A^F \in [0,1]$.

Given Eq. (A.6) can conveniently solved to yield G^I and G^F as

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

for $\varepsilon \neq 1$, and

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

for $\varepsilon = 1$. Four remarks follow. First, note that from Eqs. (A.6)-(A.9) the welfare change can be easily computed as

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

where $G^I(\varepsilon)$ and $G^F(\varepsilon)$ have just been obtained immediately above, and both G^I and G^F are expressed as explicitly dependent on ε, the parameter reflecting the degree of aversion to social inequality for the social welfare function W in Eq. (A.6). Second, 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. Third, $\overline{EE^I} > G^I(\varepsilon)$ and $\overline{EE^F} > G^F(\varepsilon)$ if and only if $\varepsilon > 0$. And, fourth, taking into account Eq. (A.10), the social welfare change associated with the reform in Eq. (A.11) can be rewritten as

$$\lambda(\varepsilon) = \frac{\left[1 - A^F(\varepsilon)\right] \times \overline{EE^F}}{\left[1 - A^I(\varepsilon)\right] \times \overline{EE^I}}, \tag{A.12}$$

where the inequality indices in Eq. (A.10) have been explicitly expressed as functions of ε. This means that the proportional social gain equals the increment in the mean equivalent expenditure $\overline{EE^F} \times \overline{EE^I}^{-1}$ times the change in the (equality) indices $\left[1 - A^F(\varepsilon)\right] \times \left[1 - A^I(\varepsilon)\right]^{-1}$.

Robustness check exercise

We show below the results corresponding to one alternative value for the pass-through parameter estimation: the extreme case of $\hat{\gamma} = 100.00\%$ (i.e., VAT tax rate changes are completely shifted to consumers, so that producers' prices stay constant; as noted above, the usual assumption in this kind work) [see Tables A.1-A.3].

Consider the complete shift case whose results are shown in Tables A.1 and A.2 (in absolute and in relative terms, respectively) and in Table A.3. The pattern is clear. First, post-reform tax revenues (and the corresponding increments) are higher for higher pass-through parameter values, which is simply a natural consequence of the price-inelastic demands for cultural goods and services [see own-price elasticities in Table 8]. Second, a higher pass-through parameter value implies a higher welfare loss from the consumers' stand point as, of course, consumers' prices must rise more. For instance, the value that we obtain for the median equivalent variation is closer to those obtained by Prieto-Rodríguez et al. (2005) for the hypothetical reforms that they consider and, also, under the complete shift assumption. The reader can compare the values in the last row of Table A.1 with the values in Table 8, last row, in Prieto-Rodríguez et al. (2005). Third, here we have also computed two alternative measures of the excess burden referred to above (one based on the equivalent variation, the other on the compensating variation). As expected both measures show a negative sign. Higher levels of income are associated with higher levels of excess burden [see Table A.1, columns 8 and 9]. Fourth, and most important, when considering the effects relative to households' incomes, the regressive nature of the VAT policy reform is clearly confirmed on average [see Table A.2].

TABLE A.1. DISTRIBUTIVE ANALYSIS OF WELFARE. ($\hat{\gamma} = 100.00\%$)

Income decile	R^1	ΔR	EV	CV	EE^I	EE^F	EB^{ev}	EB^{cv}
1: 13,632	76.41	16.93	−18.54	−19.71	542.28	497.28	−1.02	−1.41
2: 15,540	118.16	25.38	−29.14	−30.05	820.99	756.48	−1.14	−2.18
3: 20,424	143.66	30.00	−33.16	−34.68	998.58	915.74	−0.59	−0.98
4: 24,480	168.95	34.55	−38.62	−40.47	1140.89	1057.66	−2.14	−2.94
5: 26,988	162.81	38.41	−44.60	−46.54	1188.43	1078.90	−0.66	−1.66
6: 32,340	190.45	41.52	−50.29	−51.62	1316.02	1236.33	−1.55	−2.95
7: 36,000	174.83	39.47	−44.35	−45.83	1268.23	1180.62	−0.82	−2.14
8: 41,532	200.96	45.71	−53.92	−55.98	1411.28	1312.21	−0.97	−2.23
9: 52,116	234.39	51.15	−62.45	−65.04	1682.83	1561.27	−1.25	−2.39
10:199,608	254.90	63.33	−73.85	−77.07	1895.68	1728.07	−2.93	−4.29
Median	167.94	35.71	−41.47	−43.27	1200.08	1101.81	−1.17	−2.30

Key to Table A.1. Breakdown of household's tax and welfare change by income deciles. See key to Table 9. EB^{ev}: excess burden for the equivalent variation; EB^{cv}: excess burden for the compensating variation. All figures are measured at their median values in 2011 euros.

Table A.2. Distributive analysis of welfare relative to annual income ($\hat{\gamma} = 100.00\%$)

Income decile	R^1	ΔR	EV	CV	EE^I	EE^F	EB^{ev}	EB^{cv}
1: 13,632	8.52	1.84	−2.15	−2.26	59.19	54.74	−0.11	−0.19
2: 15,540	8.16	1.77	−2.00	−2.05	56.04	52.64	−0.08	−0.15
3: 20,424	7.68	1.63	−1.79	−1.87	53.18	49.77	−0.03	−0.06
4: 24,480	7.69	1.61	−1.83	−1.89	53.10	48.75	−0.10	−0.14
5: 26,988	6.16	1.46	−1.70	−1.77	45.11	41.24	−0.03	−0.06
6: 32,340	6.26	1.38	−1.65	−1.73	42.89	40.05	−0.06	−0.10
7: 36,000	5.29	1.16	−1.34	−1.41	37.67	34.89	−0.02	−0.06
8: 41,532	5.13	1.18	−1.38	−1.44	36.83	33.94	−0.02	−0.05
9: 52,116	5.22	1.12	−1.33	−1.38	36.53	33.45	−0.03	−0.05
10:199,608	3.79	0.88	−1.04	−1.08	27.55	25.58	−0.04	−0.06
Median	6.06	1.37	−1.57	−1.64	42.98	39.52	−0.04	−0.08

Key to Table A.2. Breakdown of household's tax revenues and welfare change by income deciles. See key to Table 9. EB^{ev}: excess burden for the equivalent variation. EB^{cv}: excess burden for the compensating variation. All figures (except those in column 1) are expressed in per thousand terms relative to annual income.

Table A.3 shows the counterpart to Table 11: social welfare effects of the VAT reform, but this time under the assumption that the change in tax rates is completely shifted to consumers. Comparing columns 5 in Tables 11 and A.3 the observed result is natural: everyone would expect that if the shift of the increments in tax rates are higher, for any inequality aversion parameter, the welfare loss will be higher (the values of King's λ in Table A.3 fall below those in Table 11). The ordering in the Atkinson's inequality indices remains the same, i.e. $A^I(\varepsilon) < A^0(\varepsilon) < A^F(\varepsilon)$ for all the ε's considered, so that one cannot determine whether the VAT reform would have reduced or increased the inequality of the distribution of cultural expenditure among Spanish households.

Table A.3. King's proportional increase in initial equivalent income ($\hat{\gamma} = 100\%$)

Inequality Aversion, ε	Atkinson's $A^0(\varepsilon)$ index	Atkinson's $A^I(\varepsilon)$ index	Atkinson's $A^F(\varepsilon)$ index	*King's* $\lambda(\varepsilon)$
0	0	0	0	0.92109
0.5	0.14614	0.14581	0.14655	0.92030
1.0	0.28397	0.28280	0.28528	0.91792
1.5	0.42118	0.41827	0.42432	0.91152
2.0	0.55595	0.55070	0.56148	0.89899
2.5	0.67398	0.66711	0.68105	0.88253
3.0	0.76186	0.75484	0.76893	0.86814
3.5	0.82055	0.81422	0.82685	0.85844
4.0	0.85884	0.85332	0.86428	0.85232

Key to Table A.3. See key to Table 11. The values for $A^0(\varepsilon)$ are, of course, the same as there.

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