Crisis at Home:
Mancession-induced Change in Intrahousehold Distribution

Olivier Bargain*

Laurine Martinoty+

August 3, 2015

Preliminary version, please do not quote

Abstract

The Great Recession has often been referred to as a 'mancession' in several countries including Spain and the US. Although women did experience substantial job losses during the recession, the crisis hit men harder than women for they were disproportionately represented in heavily affected sectors such as construction, manufacturing and financial services. To date, nothing is known about the way the mancession has translated within the household. More generally, we know little about how labor market opportunities affect intrahousehold distribution. To study this issue, we exploit the exogenous, gender-oriented evolution of the economic environment in Spain, using consumption data from 2006-2011. We adapt and estimate a collective model of consumption which allows testing original distribution factors. In particular, we allow the sharing rule to depend on regional-time variation in relative job opportunities during the mancession. Looking more specifically at the gender-differentiated shock from the construction sector, we also suggest a difference-in-difference estimation originally embedded in the structural model. We find that the mancession strongly impacts the way the resources are shared within the household. On average, following the improvement of their relative opportunities on the labor market, the resource share accruing to Spanish wives increased by around 5-6 percents in stable marriages. This effect is similar, in magnitude, to the distributional impact of actual husbands' unemployment. The difference-in-difference estimates confirm that most of the effect is driven by the construction sector.

Keywords: mancession, intrahousehold allocation, unemployment risk

JEL: C3, D12, D13

*Aix-Marseille University (Aix-Marseille School of Economics), CNRS & EHESS

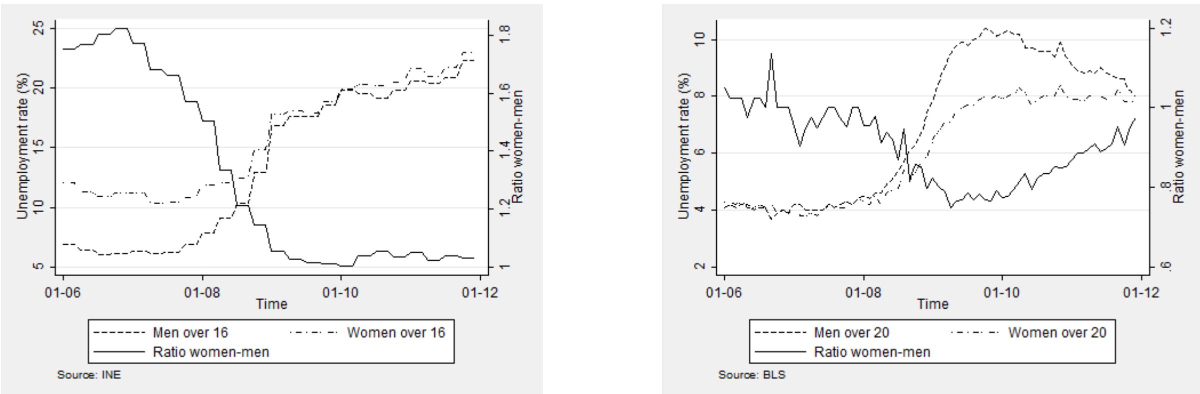
+ Université de Lyon, Lyon, F-69007, France ; CNRS, GATE Lyon Saint-Etienne, Ecully, F-69130, France ; Ecole Normale Supérieure de Lyon, Lyon, F-69007, France ; Aix-Marseille University (Aix-Marseille School of Economics), CNRS & EHESS ; Contact: martinoty@gate.cnrs.fr

Introduction

The Great recession witnessed a dramatic increase in unemployment. In the United States, by the end of 2009, unemployment hit just under 10%, which was more than double the 2007 rate of 4.6%. Some countries in the Euro zone experienced a comparable variation: in Spain, the unemployment rate rose from 8.8% to 19.7% between 2007 and 2010 and in Ireland it trippled from 4.8% to 13.2% (OECD, 2010).

A number of studies have pointed out that during recessions, the unemployment rate rises more for men than for women, resulting into the downturns being characterized as 'mancessions'. Empirically, the mancession hypothesis is clearly supported by the data. Figure 1 displays the unemployment gap between men and women aged 15-64 in Spain and in the US. In the US, men and women were as likely to be unemployed before 2008; between 2008 and 2010, this ratio clearly evolves in favor of women. In Spain, at the onset of the recession in 2008, women's unemployment was double the size of men's. By 2010, women's unemployment had risen, but men's unemployment had risen even more sharply, so that men and women were exactly as likely to be unemployed.[3]

FIGURE 1 – Unemployment rate by gender, 2006-2011



(A) Spain

(B) United States

The evidence of a mancession has been previously assessed by the literature. Studying the Great recession in the US, Sierminska and Takhtamanova (2010) note that men face higher job separation probabilities, lower job finding probabilities and, as a result, higher

[3]One may object that the labor force participation of women is more elastic to the economic opportunities, unemployed women becoming inactive. If this is the case, then the reduction in the employment gap may mechanically arises from a discouragement effect. However, the labor force participation rate of Spanish women (men) grew by 4.9% (0,3%) between 2006-2008 and by 3.4% (-0,7%) between 2009-2011. The observed men's unemployment rate is lower than the unemployment rate accounting of the discouraged workers, so that if anything, the unemployment gap may actually rather be understated.

unemployment rates than women in the US. Wall (2009)'s report points out that the large difference in the relative effects of the recession on the employment of men and women is not unusual. Between 1969 and 1991, male employment fell by an average of 3.1 percent during the five recessions experienced during the period. Female employment, on the other hand, actually tended to rise by an average of 0.3 percent during recessions (Goodman et al., 1993). The mancession does not only concern the developed economies. Concentrating on a subsample of 17 middle income countries, Cho and Newhouse (2013) find that because they were employed in the hard-hit industrial sector, men were significantly more exposed to adverse labor market events than women.

The gender composition of the sectors is the key explanatory factor to the phenomenon (Sahin et al., 2010): when the aggregate demand drops, the spendings on durable goods, capital goods, as well as investment in housing decline even more sharply. Since the employment in these sectors is more men intensive, while female tend to be more employed in the service sectors (ILO, 2010)[4], a recession implies a greater displacement rate among men than women.[5]

Spain makes no exception to this pattern. The construction sector plays a key role in explaining the variation in the gender employment gap (Bentolila et al., 2012). During the overconfident climate of the mid-2000s, the sector had been flourishing ; symmetrically, as the housing bubble burst, it was the first enter the recession, and it was particularly strongly hit. In 2006, 11,9% of the existing jobs were offered in the construction sector which represented 10,4% of the total GDP ; by 2011, the share of employment within the construction sector had shrunk down to 6,9% of the total job pool, and to 6,8% of the total GDP.[6] In parallel, construction is by far the most male specific sector: only one worker out of ten is a woman.

This exogenous change in unemployment risk may have considerably affected intrahousehold allocation of resource. As a matter of fact, there is hardly any evidence on the effect of unemployment and unemployment risk on the balance of power within the household. In this paper, we exploit the Spanish mancession as a natural experience to investigate

[4]According to the ILO, in 2008, 85 percent of women in the developed economies worked in services, 3 percent in agriculture, and 12 in manufacturing. The corresponding percentages for men are 61, 4, and 35.

[5]Of course, some studies also point out that women may be relatively more vulnerable than men in other dimensions than the employment or wage gap (Sabarwal et al., 2011). For example, newly active women may suffer from the burden of the home production in addition to the job search or the market working time. The ILO (2010) report insists on the fact that women are predominant in short term, export oriented industries.

[6]Source: Instituto nacional de Estadisticas de Espaa, `www.ine.es/daco/daco42/cne10/rem_empleo0013.xls`.

how a change in the gender unemployment gap translated into the household. While actual unemployment of a spouse may carry endogeneity issues, we can treat mancession as an exogenous shock on unemployment risk and more specifically on the gender differential in unemployment probability.

In our view, this paper contributes to the existing literature in shedding additional, well-needed light on the consequences of the economic crisis on the distribution of resources between individuals. So far, the existing literature on the mancession essentially focused on measuring the gender gap emerging with the economic crisis, and assessed the vulnerability *across* different demographic groups (Sierminska and Takhtamanova, 2010; Hoynes et al., 2012; Cho and Newhouse, 2013). While there exists widespread evidence over the redistributive impacts of economic crises between the households, little is known about the changes in the relative welfare of individuals living in these households. Indeed, a virtually substantial amount redistribution happens at the individual level. Ignoring the potential for intra-household inequality may lead to a severe underestimation of the individual-level consumption inequality (Lise and Seitz, 2011).

Precisely, we suggest a collective model identified on observation from both singles and individuals in couple without children. The model is estimated before and during crisis years (2006-2011). We retrieve the complete sharing rule and test original determinants. The first empirical approach consists of using time and spatial variation in women's relative unemployment risk. In this way, we can assess the impact of varying labor market opportunities on the intrahousehold distribution of resources in stable marriages. The second approach focuses more specifically on the role played by the construction sector. It consists in embedding a double difference approach within the sharing rule, i.e. identifying the effect of husbands being in the construction sector after the outburst of the crisis.

Controlling for price variations during the period, we find that the mancession strongly impacts the way the resources are shared within the household. Depending on the specification, we evaluate the average resource share accruing to Spanish wives for their own consumption at baseline to 0.46-0.60. Following the improvement of their relative opportunities on the labor market, their baseline share increases by 5-6 percents. Importantly, this result holds after controlling for the variation in the relative wage of both spouses, as well as their current labor market status. In terms of magnitude, the effect of this change of economic context alone has the same effect as the *actual* unemployment of the household head on the sharing of resources. In addition, using the alternative difference-in-difference setting based on the specificity of the construction sector, we conclude that this sector is a key driving force to explain the observed changes in the sharing of resources between life partners. Wives with husbands employed in the construction sector experience a 8-12 percent increase in their resource share.

The paper is structured as follows. In Section 1, we describe the model and its identification. In Section 2, we present the functional forms and the estimation method, and we motivate the choice of the distribution factors. Section 3 presents the Spanish consumption survey and provides a first look at the data. Section 4 reports the results, complemented with a discussion and various robustness checks. Section 5 concludes.

1 Model and Identification

1.1 Overview

Our approach is closely related to the most recent developments of the literature on collective models. In particular, Browning et al. (2013) and Lewbel and Pendakur (2008) consider a model where each individual in a household is characterized by a specific utility function. They suggest the complete identification of the resource sharing function and of economies of scale, exploiting data on couples and single-person households simultaneously. Browning et al. (2013) model economies of scale for each composite good using Barten scales, which reflect how much each good is jointly consumed by household members. This model is highly nonlinear in prices, expenditures and other characteristics, and is consequently difficult to estimate, both numerically and in terms of data requirements. Lewbel and Pendakur (2008) suggest a model which is slightly more restrictive but much more tractable. They posit a single function representing the economies from joint consumption and assume it is independent of total expenditure ('independence of base'). With this (testable) simplification, they can identify both resource sharing and economies of scales without observing price variation: the demand system reduces to a mildly nonlinear system of Engel curves, estimated on cross-sectional data.

The model we use is somewhat intermediary. Indeed, our six years of data surrounding the crisis are not enough variation to identify Barten scales. At the same time, we cannot ignore the little price variation that has taken place. Our aim is to retrieve the sharing rule over the period and estimate some original distribution factors like the gender differentials in unemployment risk. Hence, we rely on a model similar to Lewbel and Pendakur (2008), yet explicitly accounting for prices. We use the same basic behavioral identifying assumptions, namely the existence of some private, assignable goods, the fact that individual preferences do not change across household compositions, and the 'independence of base' assumption. This middle ground model is very convenient when using data in which spatial or time variation in prices cannot be ignored but is not big enough to be used for Barten scale identification.

1.2 Model and Assumptions

We model decisions about consumption only. Individuals are indexed by subscript $i = m, w$ for men and women respectively while superscript $k = 1, ..., K$ denotes goods. The log total expenditure in a household is denoted by x and the vector of log prices by $\boldsymbol{p}$. For a *single person*, individual log resources simply coincide with household log expenditure x. His/her welfare level is represented by:

$$u_i = v_i(x, \boldsymbol{p}, \boldsymbol{z}_i) \tag{1}$$

where $v_i(\cdot, \boldsymbol{p}, \boldsymbol{z}_i)$ is a well-behaved indirect utility function and $\boldsymbol{z}_i$ is a vector of individual characteristics. In a *couple*, individual log resources are equal to:

$$x + \log \eta_i(\boldsymbol{p}, \boldsymbol{z}) - \log s_i(\boldsymbol{p}, \boldsymbol{z}), \tag{2}$$

where $\eta_i > 0$ is a function representing the share of total expenditure accruing to individual i and $s_i > 0$ is a scale that may represent economies of scale or externalities in consumption, with $\boldsymbol{z}$ a vector of individual and household characteristics. We assume that each individual i in a couple has her/his own indirect utility function, so that the indifference curves of this individual satisfies the condition:

$$u_i = v_i(x + \log \eta_{i,n}(\boldsymbol{p}, \boldsymbol{z}) - \log s_{i,n}(\boldsymbol{p}, \boldsymbol{z}), \boldsymbol{p}, \boldsymbol{z}_i). \tag{3}$$

Function $v_i(\cdot, \boldsymbol{p}, \boldsymbol{z}_i)$ does not depend on the type n of the household. Hence, differences in expenditure patterns between a person living alone and a person living in a couple can be attributed to scaling and sharing functions only. The stability of individual preferences across household types is the key hypothesis behind identification results.[7]

We model consumption decisions in a couple as a repeated choice for which the assumption of efficiency is plausible. The most general representation of an efficient household

[7] The idea of combining data on people living alone and in couples to retrieve the complete resource sharing rule is applied in Couprie (2007); Lise and Seitz (2011); Browning et al. (2013); Lewbel and Pendakur (2008) and Bargain and Donni (2012). The assumption of stable preferences across marital status is necessary to make "situation comparisons" (i.e., compare the welfare of adults when living alone or with others) in the terminology of Pollak (1991).

decision-making process is the collective approach, which can be seen as a two-stage budgeting process (Browning et al., 1994). In a first stage, household resources are supposed to be allocated between spouses according to a sharing rule, i.e., the outcome of an unspecified decision process. Individual i living in a couple receives a share $\eta_i(\boldsymbol{p}, \boldsymbol{z})$ of total expenditure $\exp(x)$. In a second stage, expenditures on all goods are chosen *as if* each individual solved her/his own utility maximization problem subject to an individual budget constraint, i.e., spent her/his own resources $\eta_i \cdot \exp(x)$. The sharing functions $\eta_i(\boldsymbol{p}, \boldsymbol{z})$, $i = 1, 2$, are differentiable, comprised between zero and one, and sum up to unity. Following Lewbel and Pendakur (2008), we assume for the sake of identification that they do not depend on household total expenditure.[8] Yet they vary possibly with prices and a vector of household characteristics $\boldsymbol{z}$, which includes individual characteristics $\boldsymbol{z}_i$ and possibly some distribution factors $\boldsymbol{z}^d$ that capture spouses' relative bargaining position. Formally, distribution factors are variables that affect intra-household bargaining without influencing preferences or the budget constraint. They are sometimes used to identify collective models (Bourguignon et al., 2009), which is not the case in the present framework. Hence, we are free to incorporate any such factors, and in particular to test the potential role of original environmental factors like the spouses' relative risk of unemployment.

Beyond sharing, life in a couple can be characterized by other events affecting individual welfare. These are summarized here by Engel scales $s_i(\boldsymbol{p}, \boldsymbol{z})$, $i = 1, 2$. Following Lewbel and Pendakur (2008), we assume 'independent of the base' (IB), i.e. the fact that these scales are independent of the base expenditure – and, hence, of the utility level – at which they are evaluated.[9] The scale $s_i(\boldsymbol{p}, \boldsymbol{z})$ can be interpreted as a measure of the cost savings experienced by person i as a result of scale economies in the household. That is, the 'value' of total expenditure is inflated because of (partially) joint consumption by the spouses (e.g., when they ride the car together, they 'consumes' gasoline twice). With this interpretation, scales should be smaller than 1 (purely private consumption) and larger than $\eta_i(\boldsymbol{p}, \boldsymbol{z})$ (purely public consumption). Yet these scales may reflect other aspects of couple life, including consumption externalities (for instance, a man may decide to stop

[8] This assumption implies that the indifference scales derived from the model are independent of the level of utility, a desirable property most often imposed in the traditional literature on equivalence scales. Bargain and Donni (2012) show that identification results still hold, theoretically at least, when sharing functions depend on total expenditure. Also, this assumption can be mitigated in empirical applications by including measures of household wealth other than total expenditure in resource shares.

[9] This assumption is similar to the IB restriction in the equivalence scale literature (Lewbel, 1991), but it concerns individual utility functions rather than aggregated household utility functions. The IB scales can be seen as an approximation of Barten scales (used by Browning et al. (2013)) in the sense that indirect utility functions can be both IB and Barten scaled if at least one linear restriction exists on the log of Barten scales (Lewbel, 1991). SeeLewbel and Pendakur (2008).

smoking after marriage), or the departure from our assumption of preference stability (Browning et al., 2013). Hence, scaling function $s_{i,n}(\boldsymbol{p}, \boldsymbol{z})$ generally depends on all the individual characteristics of the persons living in the household, $\boldsymbol{z}$. Moreover, the fact that scaling depends on prices makes the IB scale far more general than traditional Engel scales: the idea that some goods are consumed in common (and thereby largely affected by economies of scale) can be represented here by the derivative of $s_{i,n}(\boldsymbol{p}, \boldsymbol{z})$ with respect to prices.

1.3 Model Identification

We now discuss the identification of structural components. For singles, the budget share of individual i for good k is defined by

$$w_i^k(x, \boldsymbol{p}, \boldsymbol{z}_i) = -\frac{\partial v_i(x, \boldsymbol{p}, \boldsymbol{z}_i)/\partial p^k}{\partial v_i(x, \boldsymbol{p}, \boldsymbol{z}_i)/\partial x}, \tag{4}$$

using Roy's identity. Turning to couples, we can denote log individual resource shares $x_i = x + \log \eta_i$ and apply Roy's identity to equation (3) to define the 'reduced-form' budget share on good k of spouse i as:

$$\omega_{i,n}^k(x, \boldsymbol{p}, \boldsymbol{z}) = - \left. \frac{\partial v_i(x_i - \log s_i(\boldsymbol{p}, \boldsymbol{z}), \boldsymbol{p}, \boldsymbol{z}_i)/\partial p^k}{\partial v_i(x_i - \log s_i(\boldsymbol{p}, \boldsymbol{z}), \boldsymbol{p}, \boldsymbol{z}_i)/\partial x_i} \right|_{x_i = x + \log \eta_i(\boldsymbol{p}, \boldsymbol{z})}.$$

This is the fraction of spouse i's resource share, $\exp(x) \cdot \eta_{i,n}(\boldsymbol{p}, \boldsymbol{z})$, spent on good k, expressed as a a function of household (log) expenditure x, prices $\boldsymbol{p}$ and household characteristics $\boldsymbol{z}$. Developing the derivatives on the right-hand side easily leads to:

$$\omega_{i,n}^k(x, \boldsymbol{p}, \boldsymbol{z}) = \lambda_{i,n}^k(\boldsymbol{p}, \boldsymbol{z}) + w_i^k(x + \log \eta_{i,n}(\boldsymbol{p}, \boldsymbol{z}) - \log s_{i,n}(\boldsymbol{p}, \boldsymbol{z}), \boldsymbol{p}, \boldsymbol{z}_i) \tag{5}$$

where $\lambda_{i,n}^k(\boldsymbol{p}, \boldsymbol{z}) = \partial \log s_{i,n}(\boldsymbol{p}, \boldsymbol{z})/\partial p^k$ is the elasticity of $s_{i,n}(\boldsymbol{p}, \boldsymbol{z})$ with respect to the k-th price. The right-hand side puts some structure on individual budget shares as a result of the IB restriction, using the *basic* budget share function $w_i^k(\cdot, \boldsymbol{p}, \boldsymbol{z}_i)$ used for single individuals. The consequence of this assumption is that the budget share equations of person i living in a couple differ from when living alone only in that they are translated over by the elasticity $\lambda_{i,n}^k(\boldsymbol{p}, \boldsymbol{z})$ and depend on her/his individual resources adjusted by the scaling $s_{i,n}(\boldsymbol{z})$. This property of "shape invariance", as defined by Pendakur (1999), implicitly means that single individuals are used as the demographic structure of reference.[10] We can also define an *indifference scale*, $I_i(\boldsymbol{p}, \boldsymbol{z}) = \eta_i(\boldsymbol{p}, \boldsymbol{z})/s_i(\boldsymbol{p}, \boldsymbol{z})$, as the

[10] The translation function $\lambda_i^k(\boldsymbol{p}, \boldsymbol{z})$ is specific to good k and related to the differences that may exist

adjustment applied to total expenditure that allows a person living in a couple to reach the same indifference curve as if living alone (Lewbel, 2003). [11]

We denote W_n^k the household budget share for good k and household type $n = 1, 2$ for single individuals and couples respectively. For singles, the total budget share for good k is simply defined by

$$W_1^k(x, \boldsymbol{p}, \boldsymbol{z}_i) = w_i^k(x, \boldsymbol{p}, \boldsymbol{z}_i) \tag{6}$$

For couples, total expenditure on each good k can be written as the sum of individual expenditure $\omega_{i,n}^k(x, \boldsymbol{p}, \boldsymbol{z}) \cdot x_i$, $i = m, w$, on that good. Dividing this identity by the total outlay $\exp(x)$, we directly obtain the couple's budget share function for good k as:

$$W_2^k(x, \boldsymbol{p}, \boldsymbol{z}) = \sum_{i=m,w} \eta_i(\boldsymbol{p}, \boldsymbol{z}) \cdot \left(\lambda_i^k(\boldsymbol{p}, \boldsymbol{z}) + w_i^k\left(x + \log \eta_i(\boldsymbol{p}, \boldsymbol{z}) - \log s_i(\boldsymbol{p}, \boldsymbol{z}), \boldsymbol{p}, \boldsymbol{z}_i\right)\right). \tag{7}$$

This is simply the sum of individual budget share equations weighted by individual resource shares.

Identification concentrates on how to retrieve the structural components s_i and η_i, for $i = m, w$, from the knowledge of the deterministic components $W_n^k(\cdot)$. The approach derives from Lewbel and Pendakur (2008) and Bargain and Donni (2012), with several differences compared to these studies: Lewbel and Pendakur (2008) do not use individual-specific goods for their demonstration, Bargain and Donni (2012) do so but consider couples with children and ignore price variation. In our version of the model, we do include price variation, as motivated above, and use exclusive goods (male and female clothing). The main result is summarized in the following proposition:

Proposition 1. *Assume that there exists at least one exclusive good for each adult in the household. More precisely, one good k_m is consumed by men but not by women and one other good k_w is consumed by women but not by men. Assume that $\nabla_x w_i^{k_i} \neq 0$ and $\nabla_{xx} w_i^{k_i} \neq 0$ almost everywhere for $i = m, w$ and the functions $\Delta_i^{k_i}(\cdot, \boldsymbol{p}, \boldsymbol{z}_i) \equiv$*

between goods with respect to the possibility of joint consumption or externalities. Intuitively, economies of scale may have a wealth effect and a substitution effect. The former is represented by $\log s_{i,n}(\boldsymbol{p}, \boldsymbol{z})$ and the latter by $\lambda_{i,n}^k(\boldsymbol{p}, \boldsymbol{z})$. The substitution effect is positive (negative) if good k is essentially public (private).

[11] This concept differs from an ordinary equivalence scale, which attempts to compare the welfare of an individual to that of a household, and hence suffers from the fundamental identification problem associated with interpersonal comparisons. In contrast, indifference scales can be seen as comparing the same individual in two different marital situations, respecting individualism.

$\nabla_x w_i^{k_i}(\cdot, \boldsymbol{p}, \boldsymbol{z}_i) \cdot \left[\nabla_{xx} w_i^{k_i}(\cdot, \boldsymbol{p}, \boldsymbol{z}_i)\right]^{-1}$ *are not periodic in their first argument. The sharing functions* $\eta_i(\boldsymbol{p}, \boldsymbol{z})$ *and the scaling functions* $s_i(\boldsymbol{p}, \boldsymbol{z})$*, for* $i = m, w$ *in couples can be identified from the estimation of the budget share equations* $W_n^{k_i}$ *on the exclusive goods.*

The proof follows in two steps. First, the basic budget share equations can be identified from single individuals since preferences are stable across household types n. That is, we simply have:

$$W_1^k(x, \boldsymbol{p}, \boldsymbol{z}) = w_i^k(x, \boldsymbol{p}, \boldsymbol{z}_i),$$

for any k, with $i = m, w$, and identification of the functions $w_i^k(\cdot, \boldsymbol{p}, \boldsymbol{z}_i)$ can be obtained from a sample of single male and female individuals. Second, the resource sharing functions and scaling functions for $n = 2$ can be identified from a sample of couples. Precisely, household budget share equations for adult good k_i in a couple can be written as:

$$W_2^{k_i}(x, \boldsymbol{p}, \boldsymbol{z}) = \eta_i(\boldsymbol{p}, \boldsymbol{z}) \cdot \left[\lambda_i^{k_i}(\boldsymbol{p}, \boldsymbol{z}) + w_i^{k_i}(x - \log I_i(\boldsymbol{z}), \boldsymbol{p}, \boldsymbol{z}_i)\right], \tag{8}$$

for $i = m, w$. Then, we eliminate function $\lambda_{i,2}^{k_i}(\boldsymbol{z})$ by computing the first-order derivative of this expression with respect to x:

$$\nabla_x W_2^{k_i}(x, \boldsymbol{p}, \boldsymbol{z}) = \eta_{i,2}(\boldsymbol{p}, \boldsymbol{z}) \nabla_x w_i^{k_i}(x - \log I_i(\boldsymbol{p}, \boldsymbol{z}), \boldsymbol{p}, \boldsymbol{z}_i), \tag{9}$$

where the left-hand side is identified. Differentiating this expression again with respect to x gives the second-order derivative:

$$\nabla_{xx} W_2^{k_i}(x, \boldsymbol{p}, \boldsymbol{z}) = \eta_{i,2}(\boldsymbol{p}, \boldsymbol{z}) \nabla_{xx} w_i^k(x - \log I_i(\boldsymbol{p}, \boldsymbol{z}), \boldsymbol{p}, \boldsymbol{z}_i), \tag{10}$$

and taking the ratio of (9) and (10), we have:

$$\frac{\nabla_x W_2^{k_i}(x, \boldsymbol{p}, \boldsymbol{z})}{\nabla_{xx} W_2^{k_i}(x, \boldsymbol{p}, \boldsymbol{z})} = \frac{\nabla_x w_i^{k_i}(x - \log I_i(\boldsymbol{p}, \boldsymbol{z}), \boldsymbol{p}, \boldsymbol{z}_i)}{\nabla_{xx} w_i^{k_i}(x - \log I_i(\boldsymbol{p}, \boldsymbol{z}), \boldsymbol{p}, \boldsymbol{z}_i)} = \Delta_i^{k_i}(x + \log I_i(\boldsymbol{p}, \boldsymbol{z}), \boldsymbol{p}, \boldsymbol{z}_i)$$

where function $\Delta_i^{k_i}(\cdot, \boldsymbol{z})$ is known from the first step. This condition uniquely identifies the indifference scales $I_i(\boldsymbol{p}, \boldsymbol{z})$ for $i = m, w$, provided the function $\Delta_i^{k_i}(\cdot)$ is not periodic in its first argument – a rather natural requirement. Then, for $i = m, w$, identification of sharing functions $\eta_i(\boldsymbol{p}, \boldsymbol{z})$ follows from (9) and identification of translation functions $\lambda_i^{k_i}(\boldsymbol{p}, \boldsymbol{z})$ from (8). Finally, the scaling functions $s_i(\boldsymbol{p}, \boldsymbol{z})$ can be computed for $i = 1, 2$ from the definition of $I_i(\boldsymbol{z})$.

Generic identification requires that budget share equations for exclusive goods are non-linear in log total expenditure (i.e., the second order derivative of the budget share equation must be different from zero), at least for some values of x. This is not necessarily a serious issue: as recognized by Banks et al. (1997), budget share equations are generally non-linear. Yet the functional form must be sufficiently flexible to account for it and this regularity condition must be checked in a preliminary step of the empirical analysis.

2 Empirical Implementation

2.1 Functional Forms

We turn to the empirical specification of the complete model with 8 goods which includes 7 equations. The model with only adult-specific goods, which will also be estimated, is simply a particular case. For the functional form, we suggest a parameterization that balances flexibility and empirical tractability. The first component, which appears in the specification of the different demographic groups, is the "basic" budget share equation. We adopt the following QUAIDS-type quadratic specification:

$$\begin{aligned} w_i^k(x_{i,n}, \boldsymbol{p}, \boldsymbol{z}_i) &= \bar{a}_i^k + \sum_j a_{i,j}^k z_j + \sum_q b_{i,q}^k p_q + c_i^k \left(x_{i,n} - \sum_j e_{i,j} z_j \right) \\ &\quad + d_i^k \left(x_{i,n} - \sum_j e_{i,j} z_j \right)^2, \qquad \text{for } i = 1,2,3 \text{ and } k = 1,...,K, \end{aligned}$$

where $x_{i,1} = x$ and $x_{i,2} = x + \log \eta_i$. Parameters $\bar{a}_i^k, a_{i,j}^k$, $b_{i,q}^k$, c_i^k, d_i^k and $e_{i,j}$ are specific to individual $i = m, w$ but do not depend on the demographic type n since the "basic" budget share equations are the same for single women (resp. men) and for women (resp. men) living in a couple. The demographic variables enter the specification both as a translation of budget share equations and as a translation of log scaled expenditure.

In line with the existing literature (Lewbel and Pendakur, 2008; Bargain and Donni, 2012), the basic characteristics entering $\sum_j e_{i,j} z_j$ are mainly dummies, to ease the estimation process. Age is a dummy equal to 0/1 if the individual is aged over 35, education is a dummy signaling whether the individual has a university degree. The effect of the living place is captured by two variables: a dummy for living in Madrid/Barcelona, and a dummy for living in a rural area. Wealth effects are captured by a dummy for home ownership without a pending loan, and by the ownership of a costly durable good in terms of fixed and variable costs, i.e. a car.[12] Those entering $\sum_j a_{i,j}^k z_j$ include the same variables, plus $K-1$ own and cross relative prices, the rescaled log of total expenditures and its square. In addition, the existing collective models of consumption assume that prices are constant over time (Lewbel and Pendakur, 2008; Bargain and Donni, 2012). In the context of the Great Recession, the hypothesis that prices are time-invariant is unrealistic. We account for the variation on prices by introducing a full vector of relative prices within each of the budget share equations.

Next, we specify the household budget share equations. To account for unobserved factors, we add error terms to the household budget shares previously defined:

$$\begin{aligned} \widetilde{W}_n^k(x, \boldsymbol{p}, \boldsymbol{z}) &= W_n^k(x, \boldsymbol{p}, \boldsymbol{z}) + \varepsilon_n^k, \\ &\text{for} \quad n = 1, 2 \quad \text{and} \quad k = 1, \ldots, K, \end{aligned} \tag{11}$$

where $\widetilde{W}_n^k(\cdot)$ is the stochastic extension of $W_n^k(\cdot)$. Error terms ε_n^k are traditionally interpreted as optimization/measurement errors or unobservable heterogeneity in the individual budget share equations (hence assuming random utilities), in the scales or in the resource shares. For single adults, budget shares coincide with the "basic" budget share equations plus the additive error term:

$$\widetilde{W}_1^k(x, \boldsymbol{p}, \boldsymbol{z}) = w_i^k(x, \boldsymbol{z}_i) + \varepsilon_1^k. \tag{12}$$

For couples, and for non-exclusive goods, the household budget share equations,

$$\widetilde{W}_n^k(x, \boldsymbol{p}, \boldsymbol{z}) = \sum_{i=m,w} \eta_i(\boldsymbol{p}, \boldsymbol{z}) \left[\lambda_i^k(\boldsymbol{p}, \boldsymbol{z}) + w_i^k(x - \log I_i(\boldsymbol{p}, \boldsymbol{z}), \boldsymbol{p}, \boldsymbol{z}_i)\right] + \varepsilon_2^k, \tag{13}$$

comprise the individual functions $w_i^k(\cdot, \boldsymbol{p}, \boldsymbol{z}_i)$ as already specified and structural components, defined as follows. First, the *sharing functions* are specified using the logistic form:

$$\begin{aligned} \eta_w(\boldsymbol{z}, \boldsymbol{p}) &= \frac{\exp(\overline{\beta} + \sum_j \beta_j z_j + \sum_k \beta_k p_k)}{1 + \exp(\overline{\beta} + \sum_j \beta_j z_j + \sum_k \beta_k p_k)}, \\ \eta_m(\boldsymbol{z}, \boldsymbol{p}) &= 1 - \eta_w(\boldsymbol{z}, \boldsymbol{p}), \end{aligned}$$

[12]Note that the terms e_i introduce some flexibility in the budget share equations. The elasticity of the Engel curve with respect to the total expenditures is allowed to vary with a few basic characteristics z_j: at a given level of $x_{i,n}$, the demand for good k is allowed to respond differently to a 1 euro increase in $x_{i,n}$. Note that it allows to relax the constraint imposed by the 'independence of the base' hypothesis. The rescaling of the total expenditures notably captures some wealth effects through the home and car ownership, which are then translated to the estimation of the sharing and scaling functions.

where $\overline{\beta}$, β_j and β_k are parameters. Variables in $\sum_j \beta_j z_j$ are commented more in detail below. They include distribution factors, basic controls, enriched with additional controls according to the specifications. $\beta_k z_k$ are time-region relative prices of good k with respect to the other $k-1$ other categories of goods.

Second, the log *scaling functions* that translate expenditure within the basic budget shares can be written as:

$$\log s_i(\boldsymbol{z}, \boldsymbol{p}) = \overline{\alpha}_i + \sum_j \alpha_{i,j} z_j + \sum_k \alpha_{i,k} p_k, \qquad \text{for } i = m, w,$$

where $\overline{\alpha}_i$, $\alpha_{i,j}$ and $\alpha_{i,k}$ are parameters. The scaling functions can in principle vary with all the variables entering preferences (i.e., $\boldsymbol{z}_i$ for $i = m, w$). In our specification, however, it is restricted to depend only on variables relating to individual i. Concretely, variables in $\sum_j \alpha_{i,j} z_j$ include the same variables as in $\sum_j e_{i,j} z_j$. $\sum_k \alpha_{i,k} z_k$ are the relative prices.

Third, the function that translates the basic budget shares $\lambda_i^k(\boldsymbol{z})$ is a *price elasticity*. Measuring price effects is generally challenging – and it is all the more difficult to capture their interaction with demographics in any plausible way. Therefore we restrict these terms to be constant:

$$\lambda_i^k(\boldsymbol{z}) = \overline{\lambda}_i^k, \qquad \text{for } i = m, w \text{ and } k = 1, \ldots, K.$$

2.2 Estimation Method

The complete model is estimated by the iterated SURE method. To account for the likely correlation between the error terms ε_n^k in each budget share function and the log total expenditure, each budget share equation is augmented with the 'Wu-Hausman' residuals (Banks et al., 1997; Blundell and Robin, 1999). These are obtained from reduced-form estimations of x on all exogenous variables used in the model plus some excluded instruments (a third order polynomial in household disposable income). Since budget shares sum up to one, equation for good K is unnecessary. The household budget share equations for the $K-1$ goods and for the three demographic groups are estimated simultaneously. The error terms are supposed to be uncorrelated across households but correlated across goods within households. They are also assumed to be homoskedastic for each family type. The method is described in Bargain and Donni (2012). Let h denote the observations, H_n the number of households of type n. Let $W_{h,n}$ ($\hat{W}_{h,n}$) be the vector of observed (predicted) budget shares for the $K-1$ goods consumed by household h (for some parameter θ). The vector of residuals is then: $\epsilon_{n,h}(\theta) = W_{h,n} - \hat{W}_{h,n}(\theta)$. Let $\hat{\theta}_0$ an initial consistent estimation of the vector of parameters, and $\hat{\epsilon}_{n,h}$ the corresponding vector of residuals. Then the estimated variance-covariance matrix is: $\hat{V}_n = H_n^{-1} \times \hat{\epsilon}_{n,h}\hat{\epsilon}'_{n,h}$. The SURE criterion is then:

$$\min_{\theta} \sum_{n=1}^{2} \sum_{h=1}^{H_n} (\epsilon_{n,h}(\theta))'(\hat{V}_n)^{-1}(\epsilon_{n,h}(\theta)) \tag{14}$$

This gives a new value for $\hat{\theta}$. This procedure is iterated until the variance-covariance matrix converges.

2.3 Sharing Rule Specification

To capture the effects of the mancession on the intrahousehold distribution of resources, we proceed in two steps. First, choosing a general framework, we proxy the change in the relative gender economic opportunities with the regional gender unemployment ratio, and examine its effect on the sharing of resources. Then, we have a deeper look at the mancession by concentrating on its epicenter: the construction sector, which represents 12% of the Spanish GDP in 2006, and is responsible for 50% of the male unemployment by 2011.

The regional gender unemployment ratio – As a proxy for the economic context, we use the regional men-to-women gender unemployment ratio, computed based on the regional database provided by the Spanish national statistical agency INE.[13] Then, the argument of the logistic function for the sharing rule $\boldsymbol{z}$ takes the form:

$$\boldsymbol{z} \;=\; \overline{a}_1 + a_2 u_{r,t}^{ratio} + \sum_{j=3}^{n} a_j z_i^j + \phi_r + \sum_{k=1}^{K-2} c_k p_{r,t}^k,$$

where u^{ratio} is the regional gender unemployment ratio, ϕ_r are regional fixed effects, and z^j stands for basic controls (age and education of the wife) and additional controls according to the specifications: wage ratio, unemployment of the husband, participation of the wife to the labor market. p^k are the relative prices.[14]

We are not the first paper to use the gender specific unemployment rates to study aspects of the bargaining power within the household, and recent findings suggests that this

[13]The full dataset on time-regional unemployment ratio is available in Table 7 (see Appendix).

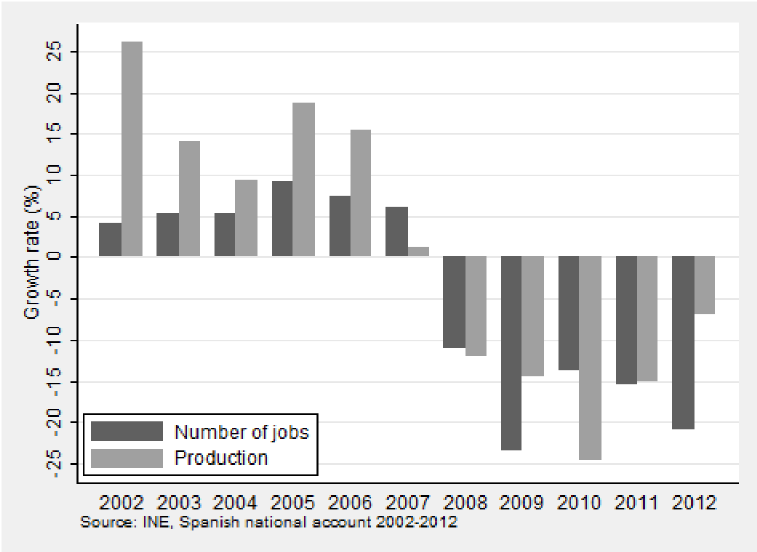
[14]The vector p contains exactly $(K-2) \times t \times r$ prices, that is, for each period and region, the number of goods present in the model, minus the price of the composite good (excluded from the system of budget share equations) and the price of male clothing (supposed identical to the price for female clothing). The model with three goods thus includes the relative price of clothing with respect to other goods, while the model with eight goods considers a vector of regional-time relative prices for six goods.

measure is relevant and promising. In Britain, Anderberg et al. (2015) observe that the unemployment across regions and genders varies greatly with the onset of the late-2000s recession, and combine this data with individual-level information on intimate partner violence. Controlling for time and area fixed effects, they find that while the general unemployment rate has no effect on the incidence of domestic violence ; on the opposite, while female unemployment increases the risk of domestic abuse, unemployment among males reduces it.

As repeatedly assessed throughout the literature (Cho and Newhouse, 2013; Hoynes et al., 2012), age categories are another important dimension of the Great Recession. Within our framework, we may want to account for a plausible age heterogeneity of the regional gender unemployment ratio. In 2006-2008, as well as in 2009-2011, the men-women gender unemployment gap is more favorable to younger women than older ones (INE). On average, the regional men-women unemployment ratio amounts to 0.73 for women aged 20-24, against 0.61 for women aged 25-54. After 2009, the opportunities increase for both age groups: women aged 20-24 are less likely to be unemployed than their male counterparts (ratio of 1.1); for women over 25, the unemployment ratio between 2009-2011 reaches 0.9. What matters for our analysis is that during the mancession episode, a woman is exposed to the same improvement in her perceived relative opportunities, whatever her age group. Using regional data on the gender unemployment gap by gender and age groups between 2006-2011, we show that there is no statistically significant difference between the growth rate of the gender relative opportunities of the two age groups. Over the period 2006-2011, younger women are exposed to a stable 8% higher relative men-women unemployment gap than older women. In addition, between 2006-2008 and 2009-2011, the relative men-women unemployment gap increases by 52% for each age group.

The construction sector before and after 2008 – Construction is an important sector within the economy: in 2008, one man out of six is employed in the construction sector. As such, it is a fact that the regional variation in unemployment is closely related to the regional shares of employment in the construction industry, which has plummeted during the recession years. Indeed, in the context of a bursting real estate bubble, the construction sector, which had been flourishing in the previous overconfident period, was the first exposed sector, and was particularly strongly hit from 2008 onward, as displayed on Figure 2. By 2011, ten percentage points of the post crisis unemployment rate were imputable to the construction sector alone (Pissarides, 2013). The reduction of employment in this sector was 36%, with regional rates varying between 18-55%. Bentolila et al. (2012) note that the raw correlation between the changes in total and construction employment shares across Spanish regions is 0.7.

Figure 2 – The construction sector in Spain, 2002-2012



Finally – and unlike the manufacturing sector, construction is the most gender specific sector of all: 91.4 of the workers within the sector are men (ILO, 2010).

We thus deepen our analysis and concentrate on the construction sector constituting the epicenter of the mancession. We run a standard difference-in-difference analysis, where the argument of the logistic function for the sharing rule $\boldsymbol{z}$ now takes the form:

$$\boldsymbol{z} \;=\; \bar{b}_1 + b_2 constru_i + b_3 post_t^{2008} + b_4 constru_i \times post_t^{2008} + \sum_{j=5}^{n} b_j z_i^j + \sum_{k=1}^{K-2} d_k p_{r,t}^k$$

We control for the specific effect of being married to a man employed in the construction sector, whatever the time period ($constru_i$); we single out the impact of the post 2008 period in the share accruing to each spouse, whatever the economic sector of the life partner ($post_t^{2008}$). In addition, we control for the same set of controls z^j presented above. We argue that b_4, the parameter of the interaction term between the post 2008 dummy and the fact for a husband to be employed in the construction sector, captures the effect of the unemployment risk on the sharing of resources at the intra-household level.

3 Data

3.1 Sample Selection

We use data on consumption from the *Encuesta de Presupuestos Familiares* collected by the Spanish national statistical agency INE. The EPF is a nationally representative panel survey of households living in Spain, and provides information on consumption, unemployment, as well as socioeconomic characteristics at the household level and the individual level for 2006-2011.[15] It surveys around 24 000 households on a yearly basis.

The original EPF sample contains detailed information on the consumption and demographic characteristics of 129 722 households during the period 2006-2011. The sample is drawn from the EBF according to the following lines. To begin with, we select single individuals and couples without children living in the home. Couples with children under 16 account for around 20% of the sample. Other type of families (alone mother or father, or extended families) account for another 40% of the original sample. This leaves us with 37.14% of the original sample (around 48,000 observations). We then discard individuals whose age is below 20 or above 45. This age category represents three quarter of the remaining sample, which now consists in 10,129 observations. To simplify our analysis, we restrain the sample to individuals who are not retirees nor students (1.5% of the remaining observations). Furthermore, we do not allow for men to be inactive so that we do not have a corner solution issue for men (0.7% of the observations – housewives are included into the sample). Around 3.5% of the households left in the sample have missing information on education or income. Income is redefined as 0 whenever the survey states that the question is "not applicable" (around 10.4% of women, and 2.9% of men). We are left with a sample of 8,875 observations, composed for 22.29% with single men, 15.26% with single women, and the remaining 62.46% with childless couples, all between 20-44.

Last, the estimation of the structural model in the difference-in-difference setting requires information on the sector of employment of the husband. In the EPF, this information is retrieved only when the husband is declared as the head of the household, a distinction left to the appreciation of the spouses. In addition to the 'large' sample, we use a 'restricted' sample, where couples with female household head are excluded. Note that within this sample, the population of singles is unchanged.

Table 1 displays standard statistical information on the individuals and their household composing this sample drawn form the EBF 2006-2011. Real data on expenditures and

[15] Although waves for 2012 and 2013 are also available, we restrict our analysis around the 2009 breaking point.

TABLE 1 – Summary statistics on individuals, by household types.

	Large sample			Restricted sample
	Single women	Single men	Couples	Couples
Women				
Aged over 35	0.46		0.21	0.20
	(0.50)		(0.41)	(0.40)
Tertiary education	0.62		0.55	0.51
	(0.49)		(0.50)	(0.50)
Income[a]	1308.90		993.69	869.74
	(639.74)		(639.63)	(591.65)
Men				
Aged over 35		0.50	0.32	0.33
		(0.50)	(0.47)	(0.47)
Tertiary education		0.45	0.43	0.43
		(0.50)	(0.50)	(0.49)
Income[a]		1366.28	1367.27	1461.94
		(698.79)	(668.56)	(644.13)
Construction sector		0.12		0.17
		(0.33)		(0.37)
Household				
Share of wife in total income	.	.	0.40	0.34
	(.)	(.)	(0.21)	(0.18)
House ownership, w/o loan	0.12	0.14	0.09	0.09
	(0.32)	(0.35)	(0.28)	(0.28)
Car ownership	0.51	0.63	0.81	0.82
	(0.50)	(0.48)	(0.39)	(0.39)
Rural area	0.12	0.21	0.20	0.21
	(0.33)	(0.41)	(0.40)	(0.41)
Madrid or Barcelona	0.13	0.11	0.10	0.10
	(0.33)	(0.31)	(0.31)	(0.31)
Observations	1354	1978	5543	4230

Note: Statistics on 2006-2011 EPF sample of working age couples (20-44). [a]Monthly nominal values. Standard error in parentheses.

income is obtained by adjusting for inflation using the inflation index (base 2008)[16]. Single individuals without children are older (35 on average for men and women, while the mean for couples is 33 and 32 respectively). While single and married men tend to have a similar education level, single women are more educated than married women. Single and married men tend to have the same level of income. In line with expectation, single women earn less than single men, and married women earn less on average than single women. The mean share of resources brought by the woman to the household is 0.4. 12% of single men work in the construction sector, which corresponds to the average share of this sector within the Spanish economy. This share amounts to 17% of household head men living in a couple. Unsurprisingly, with respect to the larger sample, couples from the

[16]Source: IMF, World Economic Outlook

restricted sample differ essentially in the fact that wives are less educated, and earn a less important share of the household income. On the other hand, interestingly, husbands and households have very similar characteristics in both samples. Looking at the bottom part of Table 1, single women tend to live relatively more in Madrid or Barcelona, while single men and couples are more likely to live in a rural area. Home ownership appears to be low for all household types. This is explained by the fact that we look at young individuals without children, age and fertility being important predictors for home ownership. In addition, to capture the wealth effect associated with home ownership, we consider only home ownership without a pending loan. The home ownership rate for young couples without children jumps to 74% once accounting for properties under mortgage.

Last, how does our selected sample of couples without children compare with the other family types, namely singles and couples with children, as well as non-nuclear family structures? Table 10 (in Appendix) reports basic demographic characteristics of household heads and their life partner according to their family type. In particular, the last three columns show that women in couple without children are younger and more educated than other women in a relationship. They are much more likely to participate into the labor market (94% against 77%) and less likely to be unemployed, so that they earn more than other married women and have a higher wage ratio within the household (0.4 against 0.3). Married men without children are also younger and more educated, but as likely to be employed or unemployed as married men in other family types. Finally, unsurprisingly, couples with children are more likely to own their home and live in a rural area than couples without children from our selected sample.

3.2 A First Look at the Data

Formally, only a pair of adult specific goods is required to identify the share of resources accruing to each spouse. We use clothing expenditures, which are gender specific.[17] In this case, the estimation relies on two gender-specific goods and one composite good: $K = 3$. Then, in the estimation, we also consider other non durable goods, to improve the efficiency of the estimations, namely: food and accommodation, transport and communication, alcohol, tobacco and gambling (commonly referred to as 'vice'), housing

[17] To ensure assignability, we define clothing expenditure of the opposite gender as 0 for singles. In the data, the share of non zero values for singles is in fact quite high : about 10% of single men report a positive value for female clothing expenses, and 20% of single women declare a non zero value for male clothing. However, the declared amounts are generally very low. The annual amount exceeds the median amount of clothing expenditures of couples only in 10% of the observations for single women, and 4% for single men.

services, leisure, personal care ($K = 8$).[18]

Table 2 presents descriptive statistics on the different consumption patterns of individuals according to their family type. In addition, it allows for a closer look at the effects of the Great recession by comparing the mean value of the consumption variables before and after the outburst of the Great Recession.

The first panel of Table 2 displays mean total expenditures according to the marital structure of the household. On average, single individuals spend about 20000 euros on consumption goods, against 31000 when living in a couple. Once excluding durable goods, investment goods and housing, the total expenditure amounts to 13000 euros for singles, and 23000 euros for individuals living in a couple. After the outburst of the Great recession, the yearly expenditures significantly dropped for all family structures.

The second panel of Table 2 presents the shares for non assignable goods. In line with expectations, the expenses are primarily dedicated to food (around 30%), transportation (20 to 25%), and services (15% to 20%). Food budget share for single men are higher because of the food outside the home component. Unsurprisingly, expenses for personal care concerns a higher budget share for single women than for single men. In general, the budget share of typically public goods decreases with the size of the household. Indeed, the expenditure share for housing services declines from 20% to 15%.[19] The economies of scale appear to be substantial, and we expect the share of privately consumed goods to increase with the scale economies allowed by the increase in the household size. The budget share devoted to private goods does increase with the size of the household when looking at men (1 additional percentage points in the budget share for leisure, 2 for personnal care), but not so much when comparing the budget share of single women to the expenditure pattern of couples. The expenditure patterns evolved little in time for single men and single women : the budget share devoted to vices and housing expenditures increased. Regarding couples, the relative share of leisure and personal care decreased, while the relative share of housing services, vices and food increased.

Table 2 give statistics for the assignable good. The expenditure for clothing increases

[18]The model being essentially static, we refrain from including expenditure on education or health. This would require adding dynamics and uncertainty in the structural model, which we leave for future research. Traditionally, expenditures on housing are not modeled because these expenditures may be difficult to evaluate for owners. We stick to the approach despite the presence of imputed rents in the EPF survey and because of measurement errors affecting this data.

[19]Perhaps surprisingly, this is not the case for transportation and communication expenditures. In fact, this reflects the differences in car ownership. Looking at Table 1, we see that the car ownership amounts to 80% when in couple, against 50 to 60% for singles. even accounting for scale economies, the car ownership typically comes with higher costs.

TABLE 2 – Summary statistics on budget shares before and after the outburst of the mancession, by household type

	Large sample						Restricted sample	
	Single women		Single men		Couples		Couples	
	2006-2008	2009-2011	2006-2008	2009-2011	2006-2008	2009-2011	2006-2008	2009-2011
Yearly expenditure								
Total yearly expenses	20754	19278***	21563	19546***	33303	30026***	33210	29924***
Selected goods+	19117	17661***	20110	18105***	30765	27731***	30711	27645***
W/o housing	13905	12339***	15150	12778***	25138	21753***	25136	21723***
Budget shares								
Food	0.3083	0.3149	0.3479	0.3453	0.3273	0.3334*	0.3286	0.3349
Transp. and comm.	0.2333	0.2261	0.2791	0.2676	0.2702	0.2630	0.2694	0.2639
Housing services	0.2019	0.2057	0.1703	0.1832**	0.1492	0.1586***	0.1494	0.1580***
Leisure	0.1068	0.1007	0.0946	0.0917	0.1052	0.0979***	0.1045	0.0969***
Vice	0.0287	0.0339*	0.0399	0.0501***	0.0390	0.0435***	0.0389	0.0437***
Personal care	0.0567	0.0594	0.0257	0.0236	0.0434	0.0417*	0.0437	0.0414*
Assignable good: yearly expenditures								
Female clothing	900	741***	0	0	957	817***	956	797***
Male clothing	0	0	669	521***	749	586***	745	588***
Budget shares								
Female clothing	0.0642	0.0593	0.0000	0.0000	0.0373	0.0361	0.0372	0.0351*
Male clothing	0.0000	0.0000	0.0424	0.0385	0.0284	0.0259**	0.0282	0.0261*
Proportion of positive values								
Female clothing	0.7859	0.7588	0.0000	0.0000	0.7920	0.7623***	0.7916	0.7559***
Male clothing	0.0000	0.0000	0.5798	0.5429	0.6723	0.6318***	0.6738	0.6344***
Observations	612	742	871	1107	2716	2827	2198	2032

Note: P-values of differences, * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. +Total expenditure for the non durable goods.

together with the change in marital status, but not as fast as the total expenditures, so that the share of clothing in the total expenses for both spouses decreases. Clothing expenditures are non-zero in 77% of the cases for women independently of their marital status, and they are strictly positive for 56% of the single men and 65% of the men living in a couple.

How did the expenses and shares of the gender specific good vary in time? In absolute terms, clothing expenditures dropped significantly during the period. In relative terms, looking at the large sample, only the clothing share for married men was downsized. The proportion of positive clothing expenditures decreases between 2006-2008 and 2009-2011.

Furthermore, as noted by Lewbel and Pendakur (2008), clothing expenditures allow us to get a first approximation of the resource shares, under the strong assumption that budget shares are independent of expenditures and demographic characteristics. In this very specific case, the resource share of married individuals would be reflected by the ratio of the budget share of clothing for married vs. singles. Computing these ratios, we find that the resource share for women increases from 0.57 to 0.61, while the share for men remains relatively stable around 0.67.[20]

The data on prices are collected by the Spanish national statistical agency INE as well. The INE reports price indexes disaggregated at the region (*Communidad Autonoma*) and good (*Classification of Individual Consumption According to Purpose*) levels. We account for the regional structural difference in prices by expressing the price indexes in base 2002 (i.e. four years prior to our first survey year). Since our analysis focuses on consumption goods only, and excludes the expenses for health care or education, we do not rely on the general price index to compute the relative prices of each goods. Instead, we compute a 'non-durable good' price index. Each non-durable good k contributes to the $NDGPI_{r,t}$ index according to its weight $w_{k,r,t}$ within the basket of non-durable goods, also available from the INE database, so that $NDGPI_{r,t} = \sum_{j=1}^{K} w_{j,r,t}.P_{j,r,t}$. The relative price of good k is then:

$$RP_{k,r,t} = \frac{P_{k,r,t}}{\sum_{j=1}^{K,j\neq k,t} w_{j,r,t}.P_{j,r,t}} \tag{15}$$

Figure 3a in Appendix plots the national price index for the selected categories of goods. Figure 3b displays the relative prices of the different goods k, with respect to the weighted evolution of the $K-1$ goods, as described in equation (15). Note that for the sake of readability, the graph displays the relative prices at the national level. The relative price that we use in the empirical estimation is the relative price of good K with respect to the index price of the $K-1$ other goods at the *regional* level.[21]

3.3 Nonlinearity of the Engel curves

As previously discussed, the generic identification of the model requires the non-linearity of the budget share equation for identifying (assignable) goods with respect to the total log expenditures. First, to check that budget share equations are indeed non linear, we estimate a reduced forms of the model on the subsamples of singles and couples. In addition, among the preliminary checks, we check for the endogeneity of the total expenditure. To do so, we first compute the Wu-Hausman residuals for each family

[20] Of course, statistically, these shares should sum to one according to the restrictions of the model.

[21] In addition, aslo in the Appendix, Figure 4 displays the regional relative price for clothing. Tables 8 and 9 report the full dataset on time-regional relative prices.

structure by regressing each budget share on all the exogenous variables of the model, plus some excluded instruments, i.e. a fourth order polynomial in the logarithm of the household income(Banks et al., 1997; Blundell and Robin, 1999, 2000). We then plug them into the Engel curves regressions.

Table 11 in Appendix reports the results of these reduced forms. The budget shares for assignable goods are indeed non linear with respect to the log of total expenditures, which is a standard result in the literature (Bargain and Donni, 2012).

In Appendix, Figure 5 graphically illustrates the non-linearity of the eight budget shares with respect to the total expenditure. Results clearly indicate a non linear behavior for most goods. Food, vice and housing services are necessity goods: holding prices constant, their demand increases with the total expenditures, but less rapidly. On the opposite, leisure, personal care and transport are luxury goods, in so far their shares in the total budget increase together with the total expenditure. In particular, women and men clothing seem to be a luxury good. At the top of the income distribution, clothing then becomes a necessity good, especially in the case of women clothing.

4 Results

4.1 Unemployment Risk

We first consider a three-equation model, containing the budget share equations for the two gender specific goods and the residual, composite good – the latter being omitted from the estimation.

The estimated coefficients for the budget share equations are presented in Table 12 in the Appendix.[22] Men and women's demand for clothing have the same determinants in terms of sign and magnitude. In line with the non-linearity identification condition presented above, it depends positively on the log of total expenditure, and negatively on its square. Besides, women tend to spend more on clothing items if they live in Madrid or Barcelona, and relatively less if they live in a rural area. Finally, note that the coefficient for the relative price is positive for men and women: the higher the relative price, the higher the expenditure on the item. This indicates that the demand for clothing is not very elastic with respect to variation to own and relative prices.

[22]For the sake of parsimony, we report the parameters corresponding to the specifications [1] and [4] of Table 3. For the sake of completeness, we also report the estimated parameters of the budget shares for the complete model with 8 goods. The parameters reported in Appendix, Table 13 correspond to the most simple specification displayed in Table 3.

Table 3 – Parameters of the sharing rule

Parameters	Model with 3 goods						Model with 8 goods		
	[1]	[2]	[3]	[4]	[5]	[6]	[7]	[8]	[9]
Constant	-0.0037	0.0478	0.0763	-0.025	0.0299	0.0485	-6.4232	-7.1899	-5.8413
	(1.4502)	(1.4571)	(1.454)	(1.451)	(1.456)	(1.4552)	(22.2237)	(22.0785)	(23.0038)
Reg. unempl. ratio	0.3831	0.3731	0.3687	0.3813	0.3735	0.3682	0.4364	0.4322	0.4419
	(0.1716)	(0.1718)	(0.1712)	(0.1718)	(0.1719)	(0.1713)	(0.2604)	(0.2542)	(0.2807)
Aged $\geq$ 35	0.1419	0.1411	0.14486	0.1432	0.1416	0.14473	0.1273	0.1263	0.1407
	(0.0853)	(0.0854)	(0.0855)	(0.0851)	(0.0853)	(0.0853)	(0.1285)	(0.1291)	(0.1308)
High ed.	0.02	0.0147	0.0161	0.018	0.0142	0.0137	0.0566	0.055	0.0501
	(0.0713)	(0.0714)	(0.0714)	(0.0711)	(0.0713)	(0.0713)	(0.1086)	(0.1096)	(0.1097)
Wage ratio		0.118			0.1206			0.0432	
		(0.0694)			(0.0748)			(0.0763)	
Husb. unempl.			0.093			0.0907			0.1457
			(0.0532)			(0.0537)			(0.0615)
Active				0.0390	-0.00559	0.0252	-0.0370	-0.0534	-0.0529
				(0.0680)	(0.0736)	(0.06866)	(0.0610)	(0.0666)	(0.0657)
Prices	YES	YES	YES	YES	YES	YES	YES	YES	YES
Region FE	YES	YES	YES	YES	YES	YES	YES	YES	YES

Note: Standard errors in parenthesis. Unless specified, the demographic characteristics are women's. Men to women unemployment ratio. Women to men wage ratio.

The estimated parameters for the sharing functions are shown in Table 3. In the baseline specification (column [1]), the sharing of resources depends on the age and education of the wife, as well as on the men-women unemployment ratio at the region level, controlling for regional fixed effects as well as for the regional price of clothing relatively to the price of the composite good. The most salient result is that the parameter for the gender unemployment ratio at the region level is positive. The higher the unemployment rate faced by the husband on the labor market relatively to the one faced by his wife, the higher the share accruing to the wife.

To have a better understanding of the results, we then estimate the shares accruing to a representative wife. Table 4 reports an increase in the average share accruing to Spanish wives between 2006 and 2011. Between 2006 and 2008, the average woman receive 46.4 % of the resources of the household. Between 2009 and 2011, the share increases by nearly 4 percentage points (i.e. 8 percents) and reaches 50.3%.[23] We decompose this variation by freezing the effects other than the change in the regional male-female unemployment gap. This exercise shows that the increase is almost entirely explained by the gender unemployment ratio: plugging the post 2009 mean of the ratio into the share computed

[23]Note that the standard errors associated with the estimated shares are not sufficiently low to assure that the positive increase is significant ; however, the fact that the parameter on the gender unemployment ratio variation is significant is reassuring: the noise comes from other parameters that are imprecisely estimated.

Table 4 – Estimated share of the average Spanish wife

	Model with 3 goods						Model with 8 goods		
Estimated	[1]	[2]	[3]	[4]	[5]	[6]	[7]	[8]	[9]
Pre 2009	0.46423	0.46527	0.47577	0.45901	0.46277	0.47212	0.59238	0.5928	0.60481
	(0.0793)	(0.0798)	(0.08006)	(0.07908)	(0.07955)	(0.07971)	(0.04858)	(0.04879)	(0.04782)
Pre 2009 (1)	0.49395	0.49422	0.50441	0.48855	0.49174	0.50072	0.62469	0.62479	0.63713
	(0.08017)	(0.0807)	(0.08086)	(0.07999)	(0.08046)	(0.08052)	(0.05165)	(0.05168)	(0.05137)
Post 2009 (2)	0.49558	0.49677	0.50782	0.49031	0.4943	0.50414	0.62609	0.62648	0.64107
	(0.08004)	(0.08056)	(0.08078)	(0.07985)	(0.08032)	(0.08044)	(0.05172)	(0.05175)	(0.05145)
Post 2009	0.50325	0.50568	0.51595	0.49849	0.50303	0.5124	0.60878	0.60926	0.62084
	(0.08345)	(0.08404)	(0.08412)	(0.08323)	(0.08375)	(0.08378)	(0.04994)	(0.05007)	(0.04921)
Heterogeneity									
Low ed.	0.46161	0.46334	0.47364	0.45665	0.4609	0.47032	0.58515	0.58567	0.59837
	(0.08053)	(0.081)	(0.08134)	(0.08029)	(0.08076)	(0.08099)	(0.05189)	(0.05211)	(0.05109)
High ed.	0.46659	0.467	0.47767	0.46112	0.46445	0.47375	0.59883	0.59916	0.61054
	(0.07913)	(0.07965)	(0.07985)	(0.07891)	(0.0794)	(0.07949)	(0.04886)	(0.04911)	(0.04821)
Aged <35	0.45764	0.45871	0.46901	0.45236	0.45619	0.46538	0.58662	0.58708	0.5985
	(0.07967)	(0.08015)	(0.0804)	(0.07944)	(0.0799)	(0.08005)	(0.04849)	(0.04868)	(0.04777)
Aged ≥35	0.49301	0.49391	0.50519	0.48803	0.49148	0.50151	0.61712	0.61734	0.6318
	(0.07966)	(0.08023)	(0.08055)	(0.07949)	(0.08)	(0.08021)	(0.05594)	(0.05623)	(0.05515)
Wage ratio 0%		0.45392			0.45117			0.58874	
		(0.08003)			(0.07976)			(0.04916)	
25%		0.46124			0.45865			0.59136	
		(0.07983)			(0.07956)			(0.0488)	
50%		0.46858			0.46615			0.59397	
		(0.07982)			(0.0796)			(0.04887)	
75%		0.47593			0.47366			0.59658	
		(0.08001)			(0.07987)			(0.04937)	
100%		0.48329			0.48119			0.59918	
		(0.08041)			(0.08038)			(0.05027)	
Husb. empl.			0.47458			0.47097			0.60302
			(0.08001)			(0.07966)			(0.04781)
Husb. unemp.			0.49784			0.49363			0.63734
			(0.08186)			(0.08161)			(0.04972)
Inactive				0.44997	0.46407	0.46626	0.6007	0.60478	0.61655
				(0.08105)	(0.08181)	(0.08186)	(0.04998)	(0.05065)	(0.04944)
Active				0.45966	0.46268	0.47255	0.59177	0.59192	0.60395
				(0.07905)	(0.07953)	(0.07968)	(0.04862)	(0.04882)	(0.04787)

Note: Standard errors in parenthesis. (1) Share computed with pre-2009 averages but post-2009 regional unemployment ratio. (2) Share computed with post-2009 averages but pre-2009 regional relative prices. Unless specified, the demographic characteristics are women's. Men to women unemployment ratio. Women to men wage ratio.

using the pre 2009 values still yields a 6 percent increase in the share accruing to the wife. The remainder of the increase is explained by the variation in the relative price of clothing. Changes in the relative prices do not seem to influence greatly the sharing of resources.

Specification [2] then allow the sharing of resources to depend on the wage ratio within the couple. The wage ratio is the standard distribution factor in the literature, and equals the ratio of the wife labor market earnings on the total household labor income. As expected, the wage ratio has a positive influence on the sharing of resources: the higher the relative contribution of the wife to the total household income, the higher her share for private consumption. The gender unemployment ratio parameter is robust to the inclusion of the wage ratio, indicating that labor market opportunities play an specific role, independently from the effective financial power represented by the wage ratio. The average shares calculated for different values of the wage ratio suggest an elasticity of the resources to the wage ratio of around 0.03. For a Spanish wife, quantitatively, the effect of the average increase in the regional unemployment ratio on the share of resources she receives is equivalent to a variation of her contribution to the household labor income from 0 to 100%.

In specification [3], we allow the sharing rule to depend on the actual unemployment of the husband. Before the economic crisis, the actual unemployment of the husband is not exogenous: husbands who lost their job have particular characteristics simultaneously affecting the parameters of the sharing rule. For this reason, no structural model relies on the impact of unemployment on the intrahousehold distribution. This arguments is not as relevant in times of crisis, where the exogenous negative demand shock generates massive lay-offs. The endogeneity issue remains; however, to compare the effect of an effective job loss with the effect of the gender specific unemployment risk is certainly insightful to understand the magnitude of the effect we observe. The unemployment status of the husband (column [3]) displays the expected sign. The fact of being married to an unemployed husband is associated with a 5 percents higher share of resources. As for specification [2], the gender unemployment ratio parameter is robust to the inclusion of the additional variable capturing the actual experience on the labor market. In terms of magnitude, the increase in the resource share of the average Spanish wife following her husband's job loss represents two thirds of the shift in resources happening during the Great recession.

The estimation of the scale economies within the household is not the purpose of our paper. Still, the parameters of the scaling function and the estimated scale economies of the life partners are reported in Table 5. In line with the existing literature, the

Table 5 – Estimated scale economies and scale parameters

	Model with 3 goods						Model with 8 goods		
Estimated scales	[1]	[2]	[3]	[4]	[5]	[6]	[7]	[8]	[9]
Wife	0.64501	0.62796	0.70161	0.63422	0.62958	0.69179	0.52976	0.52415	0.5787
	(0.20058)	(0.19506)	(0.2188)	(0.19754)	(0.19581)	(0.21706)	(0.10926)	(0.10846)	(0.1139)
Husband	1.80107	1.91727	1.72899	1.84417	1.9117	1.72247	1.30482	1.33416	1.19572
	(0.67228)	(0.73059)	(0.65576)	(0.69126)	(0.72841)	(0.65119)	(0.20413)	(0.21109)	(0.19212)
Parameters, wife									
Constant	1.22741	1.36964	1.26645	1.32206	1.37556	1.29097	-18.81754	-19.02719	-17.78617
	(1.39747)	(1.405)	(1.34775)	(1.41304)	(1.40786)	(1.36053)	(15.74509)	(15.82613)	(15.07674)
Age $\geq$ 35	0.08265	0.08187	0.0681	0.08394	0.08238	0.06856	0.13086	0.13092	0.1282
	(0.17653)	(0.17702)	(0.17143)	(0.17743)	(0.17712)	(0.1723)	(0.11759)	(0.11905)	(0.1093)
High ed.	-0.02014	-0.01579	-0.01578	-0.02069	-0.01744	-0.01775	0.03935	0.04235	0.02679
	(0.16615)	(0.16655)	(0.16152)	(0.1671)	(0.16673)	(0.16228)	(0.10472)	(0.10654)	(0.09627)
Rural	0.14113	0.13616	0.1408	0.14404	0.13824	0.14157	-0.05039	-0.05024	-0.06502
	(0.19205)	(0.19306)	(0.19013)	(0.19351)	(0.19315)	(0.19059)	(0.0913)	(0.09042)	(0.08824)
Madrid	-0.1603	-0.15967	-0.16886	-0.1627	-0.16082	-0.16789	-0.0786	-0.07626	-0.08687
	(0.18848)	(0.19041)	(0.18245)	(0.19002)	(0.19042)	(0.18348)	(0.11387)	(0.11245)	(0.11133)
Car owner	-0.31128	-0.31492	-0.29628	-0.31104	-0.31284	-0.29779	-0.255	-0.25181	-0.25061
	(0.13765)	(0.13839)	(0.13455)	(0.1384)	(0.13845)	(0.13521)	(0.06964)	(0.06839)	(0.06866)
Home owner	0.11013	0.11473	0.09414	0.1125	0.11476	0.09512	0.10891	0.10728	0.10956
	(0.18304)	(0.18434)	(0.17833)	(0.18376)	(0.18412)	(0.17927)	(0.0924)	(0.09129)	(0.08959)
Prices	YES	YES	YES	YES	YES	YES	YES	YES	YES
Region FE	YES	YES	YES	YES	YES	YES	YES	YES	YES
Parameters, husband									
Constant	-2.67216	-2.64394	-2.81874	-2.69623	-2.64119	-2.90722	7.07747	7.18896	7.26803
	(1.72048)	(1.71449)	(1.7678)	(1.71143)	(1.71535)	(1.7638)	(8.10426)	(7.93363)	(8.7012)
Age $\geq$ 35	-0.07485	-0.08169	-0.06689	-0.08464	-0.08814	-0.06907	-0.01826	-0.01967	-0.0174
	(0.21481)	(0.21315)	(0.21771)	(0.21449)	(0.21363)	(0.21787)	(0.04841)	(0.04809)	(0.05002)
High ed.	-0.19393	-0.19837	-0.17956	-0.19494	-0.19976	-0.17837	-0.16114	-0.16384	-0.16438
	(0.21183)	(0.2107)	(0.2146)	(0.21164)	(0.21121)	(0.21456)	(0.05068)	(0.05054)	(0.05214)
Rural	-0.05082	-0.04806	-0.04829	-0.05045	-0.05056	-0.04753	0.11025	0.11005	0.13345
	(0.25117)	(0.24831)	(0.25471)	(0.25)	(0.24853)	(0.25464)	(0.07449)	(0.07382)	(0.07752)
Madrid	0.2978	0.28426	0.30425	0.2911	0.28199	0.30329	0.21782	0.2143	0.23158
	(0.34603)	(0.34601)	(0.34946)	(0.34536)	(0.34517)	(0.3495)	(0.1086)	(0.10695)	(0.11488)
Car owner	0.36694	0.37922	0.36699	0.37244	0.38068	0.36597	0.11268	0.10813	0.13086
	(0.23283)	(0.23198)	(0.23564)	(0.2325)	(0.23238)	(0.2353)	(0.08033)	(0.07848)	(0.08792)
Home owner	-0.31511	-0.29628	-0.33548	-0.29135	-0.2895	-0.32766	-0.16697	-0.16487	-0.18487
	(0.29833)	(0.29052)	(0.30764)	(0.29352)	(0.29047)	(0.30712)	(0.09554)	(0.09408)	(0.10154)
Prices	YES	YES	YES	YES	YES	YES	YES	YES	YES
Region FE	YES	YES	YES	YES	YES	YES	YES	YES	YES

Note: Standard errors in parenthesis.

results indicates that the parameters are imprecisely estimated. Recall that the closest the estimated scale is to 1, the lowest the scale economies. For women, car ownership has a negative, significant impact on the scaling function, meaning that the scale economies increase with the ownership of a car. The estimated scales suggest that women benefit from more economies of scale than men within the household. For the average husband, the scale estimate is well above 1; however, it is very imprecisely estimated, so that we cannot statistically reject the fact that it may be inferior than 1, as predicted by the theory. In any case, the fact of changing family structure comes with less scale economies for men than for women.[24]

We now turn to the analysis of heterogeneous effects. Our aim is to compare the magnitude of the 'mancession' effect explored above with the impact of other demographic or social characteristics of the household. The representative wife is the average pre 2009 wife. We then manipulate the value of each demographic or social characteristic separately. Given the value of the parameters, it is not surprising that education has no impact on the share of resources accruing to the spouses. This is easily rationalized considering that the education degree is largely correlated between spouses. On the other hand, the fact that a woman is aged 35 and more increases her share of resources by 3.5 percentage points.

Even more interesting for our purpose are the estimated shares for different values of the wage ratio and different employment statuses of the husband. Here, the main takeaway form specification [2] and [3] is that the impact of the real changes in the wage ratio or the employment status of the husband on the share of resources accruing to the wife never exceeds the effect captured by the variations in the regional gender unemployment ratio. The increased perceived probability of an adverse shock has at least the same impact than the actual adverse shocks hitting the household. To go deeper into the details, specification [2] tells us that a housewife (earning 0) who would then become the unique provider of the couple (earning 100% of the household income) would receive an average 3 percentage point higher share of common resources for her private consumption.

[24] Overall, these estimates are in line with the existing findings of collective models of consumption in developed economies. In their empirical exercise on the 1990 and 1992 Canadian Family expenditure Surveys, Lewbel and Pendakur (2008) find that the average woman benefits from 46% of the total household resources. The average scale economies are 0.7 for women, and 0.8 for men, but the hypothesis that scales are indeed under 1 cannot be rejected by the data, due to the large standard error associated with the estimate. Similarly, using the 2000 French Household Budget Survey, Bargain and Donni (2012) show that in households without children, wives get 55% to 62% of the resources. The average scale economies for women (men) without children vary between 0.64 and 0.84 (0.70 and 0.97); in the simple model, the scale economy for men without children is not significantly different from 1 at the 95% confidence interval.

Specification [3] considers the impact of the husband's employment status on the sharing process: his job loss corresponds to an increase of the resource share of his spouse by 2.3 percentage points, which is again below the 4 percentage point increase related to the changes in the economic environment itself.

We now run a series of estimations to checks to test the robustness of our results. A first potential concern relates to the added worker hypothesis, namely, the fact that women enter the labor market in times of hardship as a market oriented shock coping strategy to compensate for the income loss of the primary earner. If the added worker effect is high enough, it could partly explain the apparent positive correlation between the economic context and the resource share accruing to the wife. A first answer to this concern stems form the existing literature on the added worker effect in Spain. Using a cointegration approach on quarterly Spanish data between 1976-2008, Congregado et al. (2011) identify an unemployment threshold of 11.7% up to which the added worker effect disappears due to the overwhelming discouragement effect. The threshold is clearly reached by 2009, so that the eventuality of a large scale added worker effect is limited. Still, we take the concern seriously. So far, no collective model has jointly modeled the consumption and labor allocation decisions. In this paper, we thus suppose that the decision to enter the labor market is separable. In the spirit of the two step budgeting approach, the quantity of labor supplied by the wife does not directly enter into the private utility functions: the supply of labor is decided upfront. This decision as given, the individuals maximize their utilities taking the participation of the wife as given. While this hypothesis may be strong, since participation may be endogenous to the sharing process, it has already been repeatedly made in the literature (see e.g.Zamora (2011) or Zhang (2014)). The separability assumption allows us to have the participation enter the sharing rule. We thus run the exact same regressions as in columns [1]-[3] of Table 4, but include the wife's decision to participate as an additional explanatory variable for the sharing. Across specifications [4]-[6], the participation decision never crowd out the effect of the 'mancession', which still accounts for a full 4 percentage point increase in the share accruing to wives. Note that in specification [4] and [6], the participation decision positively impacts the share – though not significantly. This positive correlation disappears in specification [5] once we take into account the wage ratio, suggesting that the positive correlation arising from the decision to enter the labor market plays through the violation of the income pooling hypothesis.

Another related labor market concern comes from the fact that clothing expenditure may not be a purely private, non durable consumption good, casting doubt on the interpretation of our result. In the context of an economic crisis, expenses for clothing items could alternatively stand for a labor market investment made by employed individuals, as well as for new job seekers. Then, the immediate question that comes to mind is whether the investment motive could explain why women's budget share for clothing increases together with the decline in the unemployment gap. Such an interpretation would favor the insurance role of marriage over the bargaining hypothesis put forward by the paper. However, a key element to bear in mind is that the identification of our model relies on the budget share of single individuals, for given characteristics and at a given point in time. The sharing rule for couples is identified using information on the single individuals with similar characteristics. In our view, there is no obvious reason why married women should invest more in clothing than single women to enter on the labor market or keep their job. The investment motive is not a convincing alternative explanatory factor to our result.

The household formation and dissolution is another important dimension that we do not jointly model into our collective model of consumption decisions.[25] Nonetheless, the regional gender unemployment gap could easily interpreted as a shift in the 'outside options' or 'extra-environmental parameters' put forward by the early bargaining models of the household à la McElroy and Horney (1981), where the position along the efficiency frontier is explained by spouses' utility in case of a divorce. In a standard vision of the bargaining power, the changes in relative opportunities raises the value of the outside option of the wife relatively to the husband. The marriage ends whenever the individual cost that each spouse bears from living in a couple outreaches the benefits associated to life sharing. A legitimate concern arises if the exogenous shift in opportunities is likely to affect the household formation and dissolution probabilities, so that the sample composition of the single and the couples varies in time.

A first answer is to recall that by adopting the widely used collective consumption model, we have maintained the assumption that households make Pareto efficient consumption decisions. This implies that household members act cooperatively: by assumption, the model does not allow for free-riding on the consumption of the public goods, nor for Nash equilibrium allocations where the divorce is a relevant threat point.[26] Even so, we

[25]As a matter of fact, very few studies have attempted to endogenize the household formation in a collective model of consumption. Recently, Mazzocco et al. (2014) study the relationship between household consumption decisions (on labor supply and savings behavior) and marital choices. (Cherchye et al., 2014) add the assumption that marriages are stable to the standard Pareto efficiency of the household consumption decisions hypothesis. They endogenize the marriage matching decisions and show that combining these two assumptions generates strong testable implications for household consumption patterns.

[26]For non cooperative Nash bargaining models, see Browning et al. (2010) or Lechene and Preston (2011).

empirically investigate the likelihood that our sample may suffer from a selection bias linked with the dynamics on the marriage market by examining the number of divorces for childless couples by region between 2006 and 2011. Independently of the number of children, the number of divorces actually tend to decrease with the economic crisis[27] A simple OLS regression analysis between the regional log number of divorces for childless couples and the regional unemployment ratio suggests that divorce and gender relative economic opportunities are unrelated.[28] This result indicates that the sample composition alone cannot entirely account for the magnitude of the effect displayed in Table 4.

The model with three goods has the advantage of simplicity and should yield robust results ; on the other hand, the limited information may decrease the efficiency of the estimates (Bargain and Donni, 2012). We then estimate a complete model including eight budget share equations. Econometrically, two gender specific goods model are just what we need to identify the sharing rule. One of the main advantages of the complete model is to improve the efficiency of the scaling point estimates: indeed, the scaling estimates for men are still non significantly different than 1, but closer to 1 on average than in the simple model. The complete model comes in support of the previous findings. The resource share accruing to women at baseline is higher than in the simple model, and reaches 59.2 to 60.4 percentage points according to the specification. During the crisis period, the share increases by 1.6 to 1.8 percentage points, which is smaller than the increase observed in the simple model. When we single out the impact of the regional gender unemployment ratio from the effect of price variation, we find that had the relative prices not change, the share would have increased by 4 percentage points as well. In the complete model, the full vector of relative prices plays against the private consumption of married women, and moderates the increase in the share accruing to them. This is an important remark, because none of the price parameters are actually significantly different from zero, while the effect of the gender employment gap is significant and positive. When we compute the estimated share, the significant increase due to one single interest variable is thus mechanically partly offset by the non-significant effect of six relative prices.

4.2 Difference-in-Difference

The literature has stressed the leading role of the construction sector in the Great recession in countries such as Spain and Ireland (Bentolila et al., 2012; Pissarides, 2013). Consequently, we narrow our analysis and concentrate on the epicenter of the mancession.

[27]Source: INE.

[28]This result also holds accounting for regional and time fixed effects.

Recall that due to data limitation on the sector of activity of the household head, we use a restricted sample for the difference-in-difference analysis. Table 14 in Appendix allows for a comparison between the original sample and the restricted sample used for the difference-in-difference estimation. Reassuringly, it appears that the results commented above are robust to the sample restriction.

Table 6 displays the results obtained exploiting the specificity of the construction sector during the Great recession. For the purpose of completeness, we propose the exact same specifications as in previous result tables 3 to 5.[29] In all specifications, the estimates indicate that women whose husband is employed in the construction sector get on average substantially less from the household resources than other women with the same age and education degree. Looking at specifications [1] to [6], before 2008, the share accruing to them varies between 44.9-45.6%, while wives whose husband is employed elsewhere or unemployed benefit from 48.8 to 49.3% of the total household resources.

The dramatic adverse shock in the construction sector has important consequences on the sharing of resources. After 2008, specifications [1]-[6] report that the share accruing to wives whose husband works in the construction sector increases by 6 to 7 percentage points, while the share of wives from other households remains statistically stable (the parameter of the time dummy is positive but non significantly different from 0). Note that the change in relative prices contributes to the increasing share by 2 percentage points, so that the pure effect of the Great recession on women with husbands working in the construction sector is around 5 percentage points.

The results are robust to the estimation of the complete model presented in columns [7]-[9]. Wives with husbands employed in the construction sector after 2008 receive a 3 to 4 percentage point higher share of the total resources for private consumption after the outburst of the economic crisis. Finally, the sign and magnitude of the parameters for the wage ratio and employment status of the spouses deserve some attention. As it was the case for the previous results, the impact of the participation of the wife to the labor market on the sharing process is virtually 0. In the simple model, the estimate for the husband job loss is negative, but small and insignificant; in the complete model, it is higher and positive, which is more in line with expectation, although not significant. The only counter-intuitive result comes from the parameter associated with the wage ratio in specification [9]. The parameter is not significant, but high in magnitude and negative,

[29] The estimates for the scale economies are reported in Appendix, Table 6. Unlike in the previous estimations, we do not exploit a regional source of variation. In the difference-in-difference estimations, to account for the regional fixed effects thus makes little economic sense. Consequently, we get rid of the regional fixed effects. Note that keeping them actually leaves the results unchanged.

Table 6 – Estimated share of Spanish wives and sharing rule parameters

	Model with 3 goods						Model with 8 goods		
Estimated	[1]	[2]	[3]	[4]	[5]	[6]	[7]	[8]	[9]
Other sector before 2008	0.4885	0.4927	0.4888	0.4868	0.4904	0.493	0.6268	0.6353	0.6307
	(0.0888)	(0.0895)	(0.0892)	(0.0888)	(0.0895)	(0.0897)	(0.0563)	(0.0559)	(0.0557)
Other sector after 2008	0.5051	0.5095	0.5052	0.5034	0.5073	0.5097	0.6158	0.6255	0.6191
	(0.0927)	(0.0936)	(0.093)	(0.0928)	(0.0936)	(0.0936)	(0.0555)	(0.0551)	(0.0549)
Construction before 2008	0.451	0.455	0.4511	0.4493	0.4528	0.4554	0.5787	0.5884	0.5824
	(0.0896)	(0.0903)	(0.0899)	(0.0896)	(0.0903)	(0.0904)	(0.0613)	(0.0607)	(0.0611)
Construction after 2008	0.523	0.528	0.523	0.5214	0.526	0.5274	0.6167	0.6246	0.6192
	(0.0932)	(0.0941)	(0.0934)	(0.0931)	(0.094)	(0.0939)	(0.0567)	(0.0563)	(0.0563)
Construction after 2008 (1)	0.505	0.5095	0.5052	0.5033	0.5073	0.5093	0.6186	0.6258	0.6212
	(0.0894)	(0.0901)	(0.0897)	(0.0894)	(0.0901)	(0.0902)	(0.0597)	(0.0593)	(0.0595)
Parameters									
Constant	1.0688	1.1212	1.0593	1.0532	1.0999	1.083	-13.216	-12.9339	-13.1151
	(1.2959)	(1.3065)	(1.2942)	(1.296)	(1.306)	(1.3026)	(20.6767)	(20.8455)	(21.0048)
Husband in construction	-0.151	-0.1513	-0.1514	-0.1505	-0.1508	-0.1507	-0.2009	-0.1976	-0.2024
	(0.0688)	(0.0687)	(0.0689)	(0.069)	(0.069)	(0.0688)	(0.0825)	(0.0807)	(0.0882)
Post 2008	-0.0117	-0.0116	-0.0115	-0.0117	-0.0115	-0.0116	-0.0429	-0.0434	-0.0461
	(0.0522)	(0.0523)	(0.0522)	(0.0523)	(0.0523)	(0.0522)	(0.074)	(0.0733)	(0.0781)
Construction $\times$ Post 2008	0.2227	0.2256	0.2227	0.2225	0.2256	0.2216	0.2046	0.1938	0.2028
	(0.0863)	(0.0865)	(0.0862)	(0.0864)	(0.0866)	(0.0861)	(0.0959)	(0.0939)	(0.1015)
Age $\geq$ 35	0.1415	0.1378	0.1416	0.1418	0.1382	0.1396	0.113	0.116	0.1139
	(0.0977)	(0.0985)	(0.0977)	(0.0977)	(0.0984)	(0.0982)	(0.1487)	(0.1512)	(0.1496)
High ed.	0.0294	0.031	0.0296	0.0287	0.0302	0.029	0.0161	0.0159	0.0187
	(0.0801)	(0.0805)	(0.0801)	(0.08)	(0.0804)	(0.0805)	(0.123)	(0.1248)	(0.123)
Husb. Unempl.		-0.0437			-0.0462			0.0834	
		(0.0781)			(0.0785)			(0.0857)	
Wage ratio			-0.0161			-0.0348			-0.1753
			(0.1017)			(0.1101)			(0.1177)
Active				0.0178	0.0206	0.0271	-0.0564	-0.0596	-0.0022
				(0.0697)	(0.07)	(0.0756)	(0.0594)	(0.0594)	(0.0696)
Prices	YES	YES	YES	YES	YES	YES	YES	YES	YES

Note: Standard errors in parenthesis. Unless specified, the demographic characteristics are women's. Men to women unemployment ratio. Women to men wage ratio.

which suggests that a higher contribution of the wife to the total income of the household is associated with a *lower* share accruing to her once sitting at the negotiating table. We explain this result by the fact that household-declared female household heads, which are more likely to bring home a higher share of the total household income, are excluded from the sample. The effect of wage ratio on the sharing may well be non linear, a women contributing with a very low share of earnings being transferred less utility than than a housewife.

Overall, all the specifications converge to one main result, namely, that the fact of having a husband in the construction sector strongly and significantly increases one's resource share devoted to private consumption within the household. Now, is it possible that our results are driven by a sample composition effect? As already mentioned, using the restricted sample, we are able to successfully replicate the findings presented above obtained on the gender employment ratio. Another way to test the coherence of the results is to exclude unemployed men from the sample. Indeed, it is likely that a consequent share of the pool of unemployed after 2008 stem from the construction sector. The results (not reported here) are unchanged. Finally, more importantly, recall that the structural model builds on the information on singles to retrieve the parameters of the sharing and scaling functions. Even in the case where men still employed in the construction sector would be very different from men employed in the sector before the economic crisis, there is no reason why the selection of the remaining employees of the construction sector based on unobserved characteristics would operate differently for singles and for men in a couple. The population of construction employees may have changed; but the comparability of preferences across household types is unlikely affected.

5 Conclusion

This paper studies the dynamics of intra-household resource allocation among Spanish households during a significant economic downturn. We first show that the Great recession starting in 2009 caused a dramatic, exogenous change in the relative employment opportunities in favor of Spanish women, commonly described as a 'mancession'. Then, we measure the extent to which the intra-household sharing of resources responds to this new order of relative gender opportunities. Beyond the closing gender gap observed at the aggregate level, we show that the mancession also invites itself over to the family negotiating table. The changing economic context has implications all the way to the core of the household consumption decisions, and as such impacts the intra-household distribution of welfare.

More broadly, this papers revisits the intra-household bargaining for the sharing of resources from an innovative angle. First, by exploiting the features of the mancession, we complement the existing literature with an alternative distribution factor, which we believe is more credible than the standard distribution factors, such as the wage or sex ratio. In addition, while previous studies generally identify the sharing rule up to a constant, we estimate an intermediary model between the structures proposed by Browning et al. (2013), Lewbel and Pendakur (2008) and Bargain and Donni (2012). Thanks to the presence of private, gender specific goods, we fully identifies the parameters of the sharing function using data on revealed preferences from individuals living in two different family structures (singles, and life partners), taking into account prices.

In line with expectations, we show that the exogenous variation in gender relative opportunities does create room for marital bargaining. As the regional gender unemployment gap closes in favor of women, their share devoted to private consumption increases by 3 to 5 percentage points (5-6 percents). Interestingly, the magnitude of this effect is higher than the effect of changes in the actual situation of individuals, like a job loss for the husband, or an increase in the relative earnings for the wife. Narrowing the analysis to the epicenter of the mancession, we find that in line with expectations, the increased share accruing to wives is not uniform across economic sectors: the essential of the measured shift in resources actually concerns women married to construction workers, who are exposed to the earlier, and higher risks of unemployment during the Great recession.

Finally, this paper leaves open several interesting questions for future research. First, while the distributive effects of the Great recession in Spain are clearly stated in the paper, the issue of welfare is far more tricky. Clearly, in absolute terms, the Great recession causes the household consumption to decrease. Still, in relative terms, consumption shifts towards women, and we may ask whether this transfer is welfare improving. Of course, the paper leaves equally open the companion, normative question of whether the fairness of the intrahousehold distribution is improved by this new deal. The existence of a redistribution necessarily lead to revisit the risk sharing role of the household: is the resource transfer happening within the household really efficient, or are husband and wife engaged in a non-cooperative bargaining solution? A convincing answer to this interrogation would require to capture the possible added-worker effect, and thus model jointly the labor supply and consumption decisions. The identification of a full income sharing rule within this complex theoretical framework is yet to come in the collective model literature. Last, a last unsolved issue deals with the long term effects of the mancession. We may wonder whether the intra-household allocation changes occurring together with the current 'mancession' will persist when the Spanish economy – and more specifically its industry and construction sectors – will go back on track towards a recovery.

References

Anderberg, D., Rainer, H., Wadsworth, J., and Wilson, T. (2015). Unemployment and Domestic Violence: Theory and Evidence. *The Economic Journal, forthcoming.*

Banks, J., Blundell, R., and Lewbel, A. (1997). Quadratic Engel Curves and Consumer Demand. *Review of Economics and Statistics*, 79(4):527–539.

Bargain, O. and Donni, O. (2012). The Measurement of Child Costs: A Rothbarth-Type Method Consistent with Scale Economies and Parents Bargaining. *European Economic Review*, 56(2010-30):792–813.

Bentolila, S., Cahuc, P., Dolado, J. J., and Le Barbanchon, T. (2012). Two-Tier Labour Markets in the Great Recession: France Versus Spain. *The Economic Journal*, 122(562):155–187.

Blundell, R. and Robin, J. M. (1999). Estimation in Large and Dissaggregated Demand Systems: an Estimator for Conditionally Linear Systems. *Journal of Applied Econometrics*, 14(3):209–232.

Blundell, R. and Robin, J.-M. (2000). Latent Separability: Grouping Goods without Weak Separability. *Econometrica*, 68(1):53–84.

Bourguignon, F., Browning, M., and Chiappori, P.-A. (2009). Efficient Intra-household Allocations and Distribution Factors: Implications and Identification. *The Review of Economic Studies*, 76(2):503–528.

Browning, M., Bourguignon, F., Chiappori, P.-A., and Lechene, V. (1994). Income and Outcomes: A Structural Model of Intrahousehold Allocation. *Journal of Political Economy*, 102(6):pp. 1067–1096.

Browning, M., Chiappori, P.-A., and Lechene, V. (2010). Distributional Effects in Household Models: Separate Spheres and Income Pooling. *The Economic Journal*, 120(545):786–799.

Browning, M., Chiappori, P.-A., and Lewbel, A. (2013). Estimating Consumption Economies of Scale, Adult Equivalence Scales, and Household Bargaining Power. *The Review of Economic Studies*, 80(4):1267–1303.

Cherchye, L., Demuynck, T., De Rock, B., and Vermeulen, F. (2014). Household Consumption when the Marriage is Stable. Technical report, KU Leuven. DPS 14.08.

Cho, Y. and Newhouse, D. (2013). How did the Great Recession affect Different Types of Workers? Evidence from 17 Middle-Income Countries. *World Development*, 41:31–50.

Congregado, E., Golpe, A. A., and van Stel, A. (2011). Exploring the Big jump in the Spanish Unemployment Rate: Evidence on an 'Added-Worker' Effect. *Economic Modelling*, 28(3):1099–1105.

Couprie, H. (2007). Time Allocation within the Family: Welfare Implications of Life in a Couple. *The Economic Journal*, 117(516):287–305.

Goodman, W., Antczak, S., and Freeman, L. (1993). Women and Jobs in Recessions: 1969-92. *Monthly Labor Review*, pages 26–35.

Hoynes, H., Miller, D. L., and Schaller, J. (2012). Who Suffers During Recessions? *The Journal of Economic Perspectives*, pages 27–47.

ILO (2010). Women in Labour Markets: Measuring Progress and Identifying Challenges. Technical report, ILO.

Lechene, V. and Preston, I. (2011). Noncooperative Household Demand. *Journal of Economic Theory*, 146(2):504 – 527.

Lewbel, A. (1991). Cost of Characteristics Indices and Household Equivalence Scales. *European Economic Review*, 35(6):1277 – 1293.

Lewbel, A. (2003). Calculating Compensation in Cases of Wrongful Death. *Journal of Econometrics*, 113(1):115 – 128.

Lewbel, A. and Pendakur, K. (2008). Estimation of Collective Household Models with Engel curves. *Journal of Econometrics*, 147(2):350–358.

Lise, J. and Seitz, S. (2011). Consumption Inequality and Intra-household Allocations. *The Review of Economic Studies*, 78(1):328–355.

Mazzocco, M., Ruiz, C., and Yamaguchi, S. (2014). Labor Supply and Household Dynamics. *The American Economic Review*, 104(5):354–359.

McElroy, M. B. and Horney, M. J. (1981). Nash-Bargained Household Decisions: Toward a Generalization of the Theory of Demand. *International Economic Review*, 22(2):pp. 333–349.

OECD (2010). OECD Factbook 2010. Economic, Environmental and Social Statistics. Technical report, OECD.

Pendakur, K. (1999). Semiparametric Estimates and Tests of Base-Independent Equivalence Scales. *Journal of Econometrics*, 88(1):1 – 40.

Pissarides, C. A. (2013). Unemployment in the Great Recession. *Economica*, 80(319):385–403.

Pollak, R. A. (1991). Welfare Comparisons and Situation Comparisons. *Journal of Econometrics*, 50(12):31 – 48.

Sabarwal, S., Sinha, N., and Buvinic, M. (2011). How Do Women Weather Economic Shocks? What We Know. *Economic Premise, World Bank*, 46:1–6.

Sahin, A., Song, J., and Hobijn, B. (2010). The Unemployment Gender Gap during the 2007 Recession. Technical report, Current Issues in Economics and Finance, Federal Reserve Bank of New York.

Sierminska, E. and Takhtamanova, Y. (2010). Job Flows, Demographics and the Great Recession. CEPS/INSTEAD Working Paper Series 2010-41, CEPS/INSTEAD.

Wall, H. (2009). The Effects of Recessions Across Demographic Groups. Technical report, Federal Reserve Bank of Saint Louis.

Zamora, B. (2011). Does Female Participation affect the Sharing Rule? *Journal of Population Economics*, 24(1):47–83.

Zhang, S. (2014). Wage Shocks, Household Labor Supply, and Income Instability. *Journal of Population Economics*, 27(3):767–796.

6 Appendix

TABLE 7 – Regional unemployment 2006-2011, by region

	Unemployment rate, men						Unemployment rate, women						Men/women unemployment ratio					
Region	2006	2007	2008	2009	2010	2011	2006	2007	2008	2009	2010	2011	2006	2007	2008	2009	2010	2011
Andalucia	9.27	9.56	15.26	24.05	26.70	28.69	17.67	17.41	21.20	26.85	29.17	31.97	0.52	0.55	0.72	0.90	0.92	0.90
Aragon	3.87	3.91	6.63	13.41	14.19	16.88	7.86	7.17	8.20	12.59	15.94	17.32	0.49	0.55	0.81	1.07	0.89	0.97
Asturias	6.80	6.48	6.77	12.64	15.15	18.31	12.28	10.87	10.61	14.34	16.81	17.29	0.55	0.60	0.64	0.88	0.90	1.06
Baleares	5.23	6.02	9.69	19.25	20.79	22.35	8.06	8.70	10.78	16.25	19.30	21.28	0.65	0.69	0.90	1.18	1.08	1.05
Canarias	9.47	8.62	16.13	25.58	29.15	29.40	14.60	12.92	18.75	26.55	27.91	29.13	0.65	0.67	0.86	0.96	1.04	1.01
Cantabria	4.27	4.45	5.88	11.55	12.67	16.09	9.43	7.97	8.84	12.56	15.00	14.30	0.45	0.56	0.67	0.92	0.84	1.13
Castilla y Leon	5.24	4.81	6.92	12.08	14.22	15.69	12.26	10.45	13.36	16.53	17.86	18.37	0.43	0.46	0.52	0.73	0.80	0.85
Castilla Mancha	5.33	5.22	9.29	17.19	19.22	20.88	14.69	11.56	15.29	21.36	24.15	26.09	0.36	0.45	0.61	0.80	0.80	0.80
Cataluna	5.24	5.59	9.04	17.21	18.59	19.77	8.14	7.63	8.70	14.98	16.51	18.43	0.64	0.73	1.04	1.15	1.13	1.07
Valencia	6.41	6.90	10.91	20.97	22.93	23.70	11.03	11.27	13.43	20.49	22.76	24.33	0.58	0.61	0.81	1.02	1.01	0.97
Extremadura	9.94	9.27	11.09	17.52	20.57	23.03	18.56	18.50	21.69	25.03	26.30	27.89	0.54	0.50	0.51	0.70	0.78	0.83
Galicia	5.98	5.72	7.35	11.66	14.59	16.56	11.27	9.88	10.22	13.36	16.19	18.06	0.53	0.58	0.72	0.87	0.90	0.92
Madrid	4.64	4.92	7.85	13.91	15.57	16.40	8.32	7.84	9.51	13.79	16.14	16.27	0.56	0.63	0.83	1.01	0.96	1.01
Murcia	5.97	6.00	11.95	21.63	23.19	25.01	10.82	9.92	13.14	18.49	22.44	24.96	0.55	0.60	0.91	1.17	1.03	1.00
Navarra	4.15	3.14	5.64	9.99	11.50	12.57	7.02	6.90	8.42	11.94	12.40	13.49	0.59	0.46	0.67	0.84	0.93	0.93
Pais Vasco	5.74	4.97	5.73	11.08	10.29	11.76	9.03	7.85	7.77	11.65	11.18	13.04	0.64	0.63	0.74	0.95	0.92	0.90
La Rioja	5.11	4.05	6.47	12.59	12.85	16.65	7.57	8.32	9.91	12.72	15.87	17.91	0.68	0.49	0.65	0.99	0.81	0.93

Own calculation from INE data: Tasas de paro por distintos grupos de edad, sexo y comunidad autnoma

Table 8 – Regional relative price 2006-2011, by region

Region	Food						Vice						Clothing					
	2006	2007	2008	2009	2010	2011	2006	2007	2008	2009	2010	2011	2006	2007	2008	2009	2010	2011
Andalucia	1.067	1.093	1.120	1.134	1.096	1.070	1.025	1.073	1.075	1.194	1.317	1.421	0.951	0.929	0.898	0.890	0.867	0.840
Aragon	1.060	1.079	1.114	1.122	1.091	1.068	1.027	1.068	1.068	1.177	1.301	1.407	0.945	0.927	0.895	0.889	0.872	0.849
Asturias	1.050	1.079	1.110	1.122	1.095	1.079	1.025	1.071	1.072	1.182	1.298	1.393	0.981	0.967	0.933	0.925	0.904	0.877
Baleares	1.054	1.080	1.104	1.113	1.073	1.047	1.012	1.054	1.050	1.146	1.245	1.330	0.954	0.938	0.909	0.898	0.879	0.854
Canarias	1.027	1.058	1.087	1.117	1.084	1.064	1.144	1.102	1.115	1.188	1.196	1.229	0.916	0.896	0.856	0.856	0.841	0.813
Cantabria	1.063	1.082	1.106	1.120	1.081	1.054	1.024	1.066	1.064	1.174	1.282	1.375	0.945	0.923	0.890	0.883	0.863	0.837
Castilla y Leon	1.063	1.086	1.117	1.132	1.105	1.084	1.026	1.070	1.072	1.185	1.302	1.403	0.958	0.941	0.910	0.905	0.885	0.859
Castilla Mancha	1.060	1.080	1.110	1.128	1.090	1.066	1.028	1.079	1.082	1.204	1.330	1.433	0.942	0.923	0.892	0.887	0.866	0.837
Cataluna	1.052	1.074	1.099	1.119	1.092	1.071	0.999	1.043	1.045	1.140	1.238	1.320	0.983	0.966	0.936	0.925	0.903	0.877
Valencia	1.056	1.087	1.118	1.134	1.101	1.079	1.025	1.072	1.073	1.187	1.312	1.417	0.968	0.952	0.922	0.915	0.897	0.871
Extremadura	1.043	1.068	1.101	1.117	1.083	1.058	1.048	1.097	1.101	1.228	1.359	1.469	0.947	0.927	0.894	0.887	0.864	0.834
Galicia	1.046	1.064	1.093	1.113	1.084	1.070	1.001	1.054	1.050	1.145	1.241	1.321	0.957	0.942	0.908	0.901	0.881	0.853
Madrid	1.052	1.071	1.092	1.101	1.069	1.046	1.025	1.078	1.077	1.189	1.309	1.411	0.930	0.913	0.883	0.871	0.851	0.825
Murcia	1.067	1.106	1.132	1.143	1.104	1.070	1.011	1.047	1.049	1.163	1.278	1.371	0.964	0.940	0.910	0.903	0.883	0.856
Navarra	1.059	1.076	1.104	1.112	1.082	1.057	1.013	1.057	1.065	1.172	1.286	1.382	0.989	0.976	0.948	0.942	0.927	0.902
Pais Vasco	1.070	1.097	1.125	1.145	1.120	1.103	1.015	1.055	1.057	1.151	1.257	1.345	0.932	0.919	0.889	0.878	0.861	0.838
La Rioja	1.060	1.075	1.095	1.101	1.069	1.045	1.013	1.055	1.059	1.167	1.283	1.378	0.939	0.922	0.892	0.884	0.870	0.844

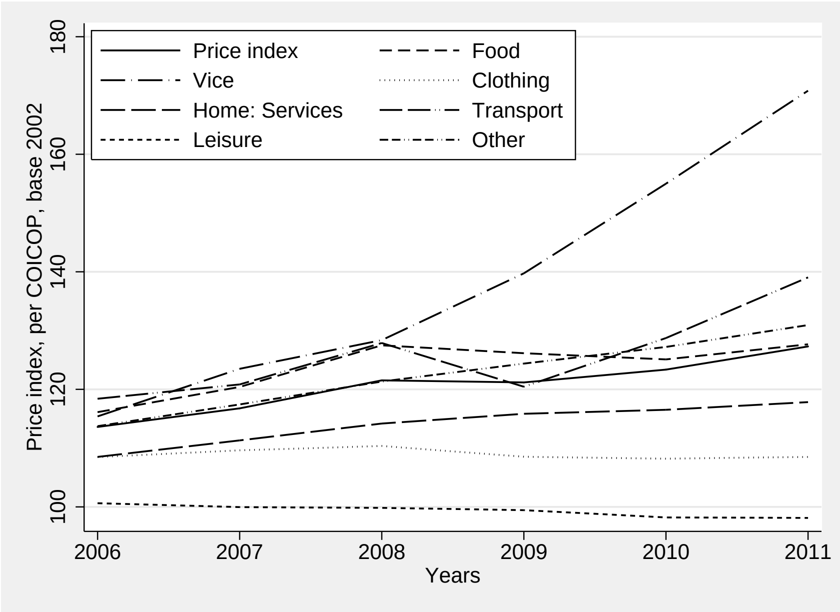
Table 9 – Regional relative price 2006-2011, by region (continued)

Region	Leisure						Personal Care						Transport					
	2006	2007	2008	2009	2010	2011	2006	2007	2008	2009	2010	2011	2006	2007	2008	2009	2010	2011
Andalucia	0.870	0.833	0.794	0.794	0.766	0.737	0.996	0.997	0.991	1.029	1.031	1.026	1.004	0.986	0.997	0.944	0.988	1.033
Aragon	0.888	0.863	0.822	0.825	0.796	0.771	1.008	1.015	1.004	1.044	1.055	1.056	1.010	0.995	1.003	0.953	0.992	1.032
Asturias	0.861	0.829	0.792	0.794	0.769	0.746	1.005	1.015	1.017	1.053	1.060	1.051	1.010	0.992	1.000	0.951	0.986	1.024
Baleares	0.869	0.842	0.808	0.815	0.800	0.766	1.013	1.016	1.009	1.051	1.057	1.058	1.018	1.003	1.021	0.969	1.008	1.054
Canarias	0.882	0.846	0.802	0.810	0.789	0.758	0.999	1.001	0.978	1.014	1.015	0.998	1.056	1.048	1.077	1.013	1.070	1.134
Cantabria	0.891	0.863	0.836	0.850	0.828	0.805	1.007	1.012	0.998	1.036	1.042	1.034	1.016	1.004	1.021	0.959	1.010	1.062
Castilla y Leon	0.863	0.827	0.789	0.790	0.763	0.738	0.999	1.001	0.994	1.031	1.034	1.034	1.002	0.989	0.997	0.944	0.982	1.022
Castilla Mancha	0.862	0.827	0.785	0.786	0.766	0.737	0.997	0.998	0.987	1.022	1.025	1.021	1.023	1.011	1.024	0.968	1.011	1.056
Cataluna	0.894	0.866	0.834	0.835	0.806	0.781	1.015	1.027	1.024	1.066	1.079	1.083	0.997	0.978	0.990	0.932	0.968	1.008
Valencia	0.870	0.835	0.799	0.804	0.779	0.752	1.010	1.013	1.006	1.040	1.043	1.038	1.005	0.988	0.999	0.944	0.983	1.024
Extremadura	0.874	0.834	0.790	0.787	0.768	0.738	1.022	1.022	1.008	1.043	1.048	1.043	1.015	0.993	1.001	0.945	0.982	1.029
Galicia	0.895	0.868	0.829	0.832	0.805	0.778	0.989	0.990	0.981	1.015	1.018	1.011	1.018	1.006	1.020	0.963	1.009	1.053
Madrid	0.884	0.856	0.823	0.826	0.799	0.775	1.010	1.023	1.022	1.064	1.072	1.073	1.026	1.011	1.028	0.977	1.020	1.059
Murcia	0.866	0.828	0.800	0.803	0.773	0.746	1.019	1.017	1.014	1.048	1.048	1.049	0.999	0.974	0.990	0.934	0.980	1.034
Navarra	0.906	0.880	0.846	0.856	0.830	0.809	1.013	1.029	1.031	1.073	1.091	1.102	0.996	0.981	0.988	0.934	0.971	1.012
Pais Vasco	0.880	0.846	0.814	0.817	0.795	0.773	0.999	1.005	1.001	1.032	1.037	1.036	1.002	0.987	0.997	0.942	0.979	1.019
La Rioja	0.895	0.872	0.841	0.839	0.805	0.792	1.027	1.032	1.034	1.075	1.093	1.097	0.982	0.969	0.978	0.926	0.966	1.003

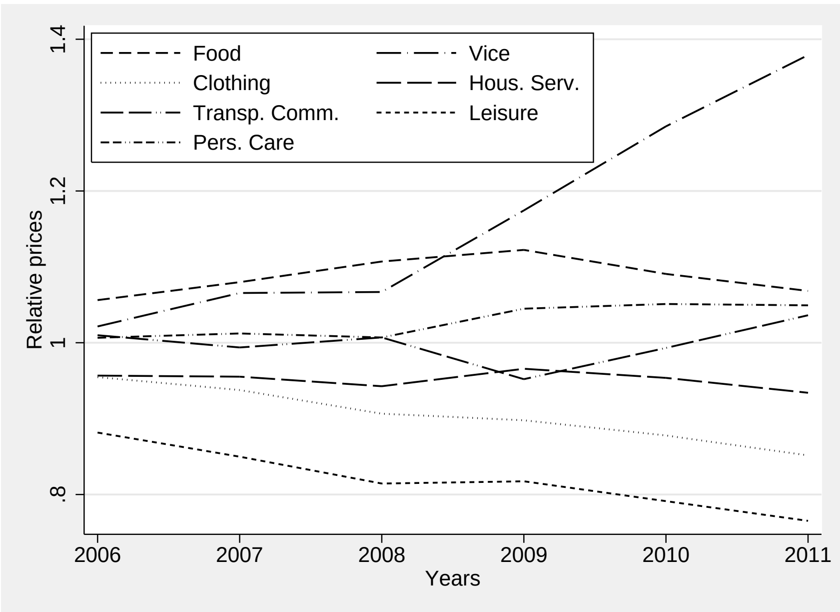
Table 10 – Summary statistics on household heads and life partners aged 20-44, by household type, 2006-2011

	Single men			Single women			Couples		
	No child	With child	Other	No child	With child	Other	No child	With child	Other
Women									
Age (in years)				34.72	37.76	34.85	31.59	35.87	34.01
				(5.92)	(5.14)	(6.41)	(4.98)	(4.77)	(6.11)
Primary education				0.04	0.12	0.13	0.05	0.10	0.15
				(0.19)	(0.32)	(0.34)	(0.21)	(0.30)	(0.36)
Secondary education 1				0.14	0.33	0.28	0.18	0.29	0.35
				(0.35)	(0.47)	(0.45)	(0.38)	(0.45)	(0.48)
Secondary education 2				0.20	0.27	0.28	0.23	0.23	0.27
				(0.40)	(0.44)	(0.45)	(0.42)	(0.42)	(0.45)
Superior education				0.62	0.29	0.31	0.55	0.38	0.22
				(0.49)	(0.45)	(0.46)	(0.50)	(0.49)	(0.41)
Income				1308.90	1084.78	1025.55	993.69	720.05	628.53
				(639.74)	(634.80)	(572.65)	(639.63)	(720.34)	(638.49)
Active				0.99	0.95	0.98	0.94	0.76	0.77
				(0.11)	(0.22)	(0.15)	(0.24)	(0.43)	(0.42)
Employed				0.92	0.78	0.88	0.82	0.62	0.62
				(0.40)	(0.53)	(0.33)	(0.43)	(0.53)	(0.52)
Unemployed				0.08	0.18	0.10	0.12	0.15	0.15
				(0.28)	(0.39)	(0.30)	(0.33)	(0.35)	(0.35)
Men									
Age (in years)	35.23	36.11	33.47				33.32	37.66	36.17
	(5.70)	(6.60)	(6.18)				(4.98)	(4.51)	(5.70)
Primary education	0.07	0.13	0.18				0.07	0.12	0.17
	(0.25)	(0.33)	(0.38)				(0.25)	(0.32)	(0.37)
Secondary education 1	0.25	0.37	0.29				0.26	0.33	0.39
	(0.43)	(0.48)	(0.45)				(0.44)	(0.47)	(0.49)
Secondary education 2	0.24	0.19	0.24				0.25	0.23	0.24
	(0.42)	(0.40)	(0.43)				(0.43)	(0.42)	(0.43)
Superior education	0.45	0.31	0.29				0.43	0.33	0.20
	(0.50)	(0.46)	(0.46)				(0.50)	(0.47)	(0.40)
Income	1366.28	1190.24	1109.38				1367.27	1481.58	1214.86
	(698.79)	(598.76)	(543.83)				(668.56)	(754.25)	(712.64)
Employed	0.88	0.90	0.89				0.91	0.91	0.86
	(0.33)	(0.30)	(0.31)				(0.29)	(0.28)	(0.35)
Unemployed	0.12	0.10	0.11				0.09	0.09	0.14
	(0.33)	(0.30)	(0.31)				(0.29)	(0.28)	(0.35)
Household									
Homeowner w/o loan	0.14	0.53	0.21	0.12	0.23	0.20	0.09	0.20	0.24
	(0.35)	(0.50)	(0.41)	(0.32)	(0.42)	(0.40)	(0.28)	(0.40)	(0.43)
Homeowner	0.62	0.80	0.40	0.61	0.63	0.41	0.74	0.82	0.63
	(0.49)	(0.40)	(0.49)	(0.49)	(0.48)	(0.49)	(0.44)	(0.39)	(0.48)
Rural area	0.21	0.24	0.19	0.12	0.17	0.12	0.20	0.25	0.25
	(0.41)	(0.43)	(0.40)	(0.33)	(0.37)	(0.33)	(0.40)	(0.43)	(0.43)
Madrid-Barcelona	0.11	0.11	0.19	0.13	0.10	0.15	0.10	0.08	0.13
	(0.31)	(0.31)	(0.39)	(0.33)	(0.31)	(0.36)	(0.31)	(0.28)	(0.33)
Wage ratio							0.40	0.29	0.31
							(0.21)	(0.24)	(0.27)
Observations	1978	548	656	1354	2132	685	5543	21211	1665

Figure 3 – Domestic prices, by good category, base 2002



(a) Price index



(b) Relative price

Figure 4 – Relative price, clothing

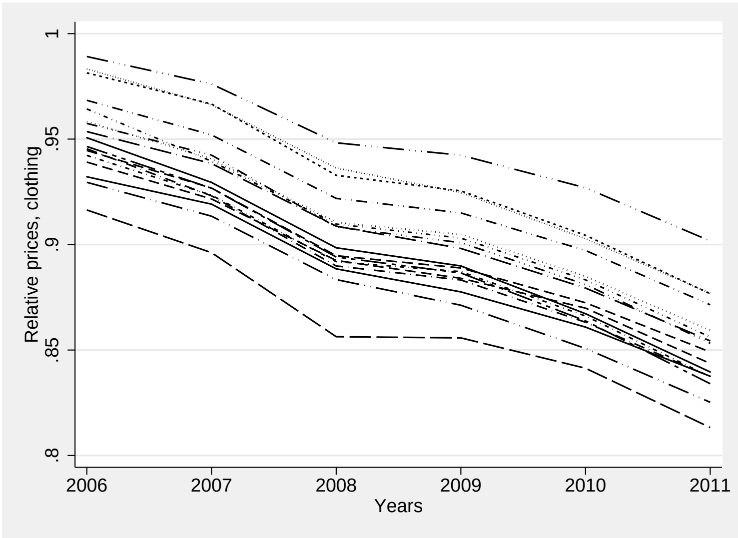


Figure 5 – Engel curves, kernel weighted local polynomial smoothing

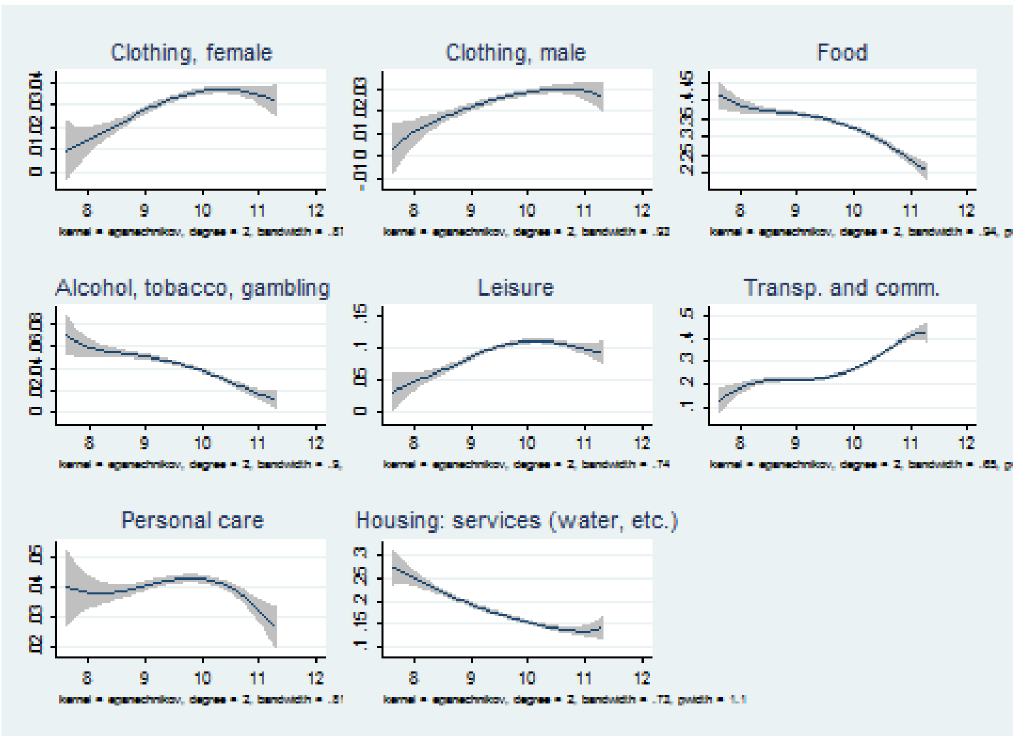


Table 11 – Non linearities in budget shares of assignable goods

	Linear	Quadratic	
		W/o Wu-Hausman	Wu-Hausman
Clothing share, single women			
Log yearly expenditure	0.00743***	0.223***	0.228***
	(2.60)	(3.78)	(3.88)
Squared log yearly expenditure		-0.0115***	-0.0106***
		(-3.66)	(-3.35)
Wu-Hausman residual			-0.0266***
			(-3.32)
Observations	1354	1354	1354
Clothing share, single men			
Log yearly expenditure	0.00975***	0.156***	0.154***
	(4.79)	(3.87)	(3.81)
Squared log yearly expenditure		-0.00779***	-0.00797***
		(-3.63)	(-3.70)
Wu-Hausman residual			0.00617
			(1.07)
Observations	1978	1978	1978
Clothing share, women in couple			
Log yearly expenditure	0.00549***	0.152***	0.152***
	(5.18)	(6.25)	(6.25)
Squared log yearly expenditure		-0.00743***	-0.00708***
		(-6.03)	(-5.72)
Wu-Hausman residual			-0.00804***
			(-2.78)
Observations	5543	5543	5543
Clothing share, men in couple			
Log yearly expenditure	0.00612***	0.0799***	0.0799***
	(6.14)	(3.48)	(3.48)
Squared log yearly expenditure		-0.00374***	-0.00351***
		(-3.22)	(-3.01)
Wu-Hausman residual			-0.00522*
			(-1.91)
Observations	5543	5543	5543

Note: $^{+}$ P-values of differences, * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Additional controls : age, tertiary education dummy, rural area dummy, main city dummy, dummy for home and car ownership, year fixed effects.

TABLE 12 – Estimated coefficients of the simple model – Budget share equations of men and women

	Simple model, specification [1]		Simple model, specification [4]	
	Budget share for men clothing	Budget share for women clothing	Budget share for men clothing	Budget share for women clothing
Constant	-0.8030	-1.4473	-0.7993	-1.4537
	(0.1557)	(0.2392)	(0.1551)	(0.2400)
Characteristics				
Aged over 35	-0.0036	0.0027	-0.0037	0.0025
	(0.0024)	(0.0055)	(0.0024)	(0.0055)
University degree	0.0025	0.0030	0.0026	0.0030
	(0.0028)	(0.0047)	(0.0028)	(0.0047)
Log scaled exp.	0.1693	0.2802	0.1684	0.2816
	(0.0327)	(0.0503)	(0.0326)	(0.0505)
Sq.log scaled exp.	-0.0089	-0.0138	-0.0089	-0.0139
	(0.0018)	(0.0027)	(0.0018)	(0.0027)
Rural resident	0.0002	-0.0087	0.0003	-0.0088
	(0.0028)	(0.0042)	(0.0028)	(0.0042)
Madrid Barcelona resident	0.0044	0.0149	0.0043	0.0149
	(0.0043)	(0.0076)	(0.0043)	(0.0076)
House owner	0.0034	0.0030	0.0035	0.0033
	(0.0040)	(0.0075)	(0.004)	(0.0075)
Car owner	-0.0056	-0.0126	-0.0056	-0.0126
	(0.0037)	(0.0051)	(0.0037)	(0.0051)
Regional relative prices				
Rel.price of clothing	0.0557	0.1254	0.0560	0.1256
	(0.0267)	(0.0370)	(0.0267)	(0.0370)

Note: The parameters of the budget share correspond to specification [1].

TABLE 13 – Estimated coefficients of the complete model – Budget share equations of men and women

	Clothing		Food		Vice		Leisure		Transport		Personal care	
	Men	Women	Men	Women	Men	Women	Men	Women	Men	Women	Men	Women
Constant	-0.0723	-1.1264	-4.2224	-1.1956	0.8252	-0.3532	-2.8399	-1.7409	6.1244	3.3281	-0.3133	-1.1548
	(0,3362)	(0,3854)	(1,0246)	(1,023)	(0,4647)	(0,3738)	(0,5488)	(0,5921)	(1,2671)	(1,1725)	(0,226)	(0,318)
Characteristics												
Aged over 35	-0.0038	0.0038	0.0035	-0.0017	0.0106	-0.0039	-0.0041	0.0116	-0.0193	-0.0285	0.0014	0.0040
	(0,002)	(0,0029)	(0,0061)	(0,0074)	(0,0028)	(0,0031)	(0,0036)	(0,0041)	(0,0066)	(0,0069)	(0,0014)	(0,0021)
University degree	0.0011	0.0041	-0.0063	-0.0020	-0.0269	-0.0174	0.0394	0.0191	-0.0093	0.0047	0.0022	-0.0024
	(0,0019)	(0,0026)	(0,006)	(0,0063)	(0,0029)	(0,0026)	(0,0038)	(0,0034)	(0,0068)	(0,0058)	(0,0014)	(0,0019)
Log scaled exp.	0.1750	0.2542	0.9175	0.4081	0.0769	0.1312	0.5014	0.4751	-1.4611	-1.0812	0.0620	0.1925
	(0,0293)	(0,0335)	(0,1044)	(0,1162)	(0,047)	(0,0392)	(0,0477)	(0,0624)	(0,1377)	(0,1536)	(0,0221)	(0,0324)
Sq.log scaled exp.	-0.0093	-0.0123	-0.0526	-0.0264	-0.0063	-0.0091	-0.0249	-0.0242	0.0817	0.0597	-0.0031	-0.0100
	(0,0017)	(0,0018)	(0,0058)	(0,0062)	(0,0026)	(0,0021)	(0,0026)	(0,0033)	(0,0078)	(0,0085)	(0,0012)	(0,0017)
Rural resident	0.0005	-0.0035	0.0109	-0.0192	0.0074	-0.0044	-0.0117	0.0023	0.0228	0.0030	-0.0061	-0.0014
	(0,0023)	(0,0023)	(0,0092)	(0,0093)	(0,0041)	(0,0038)	(0,0047)	(0,0051)	(0,0095)	(0,0099)	(0,0017)	(0,0022)
Madrid Barcelona resident	0.0077	0.0144	-0.0128	-0.0014	0.0017	-0.0133	0.0043	0.0283	0.0150	-0.0332	-0.0039	0.0042
	(0,0036)	(0,0038)	(0,011)	(0,0111)	(0,005)	(0,0046)	(0,0071)	(0,0075)	(0,0116)	(0,0116)	(0,0025)	(0,0031)
House owner	0.0025	-0.0027	0.0081	-0.0133	-0.0055	-0.0039	0.0071	0.0060	-0.0114	-0.0127	0.0018	0.0019
	(0,0029)	(0,0031)	(0,0101)	(0,0106)	(0,0046)	(0,0043)	(0,0057)	(0,0065)	(0,01)	(0,01)	(0,0021)	(0,0028)
Car owner	-0.0030	-0.0083	-0.0678	-0.0431	0.0004	-0.0063	-0.0206	0.0067	0.1292	0.0856	-0.0049	-0.0057
	(0,0022)	(0,0023)	(0,0075)	(0,0075)	(0,0034)	(0,0032)	(0,0043)	(0,0043)	(0,0074)	(0,0073)	(0,0015)	(0,0019)
Regional relative prices												
Clothing	0.0000	0.0970	-0.0209	0.0804	-0.0275	-0.0573	0.0649	0.0633	0.0058	-0.2484	-0.0261	0.1740
	(0,0662)	(0,0748)	(0,1957)	(0,1744)	(0,0798)	(0,0671)	(0,1085)	(0,1111)	(0,223)	(0,1997)	(0,0412)	(0,0604)
Food	-0.2295	0.0066	0.3082	0.0037	-0.2061	0.0221	0.0792	-0.1733	0.0834	0.5640	0.0097	0.0828
	(0,1077)	(0,1232)	(0,3112)	(0,2948)	(0,1308)	(0,1148)	(0,171)	(0,1874)	(0,3631)	(0,3383)	(0,0695)	(0,0991)
Vice	-0.0063	0.0481	-0.1200	0.0516	0.0462	-0.0030	0.0715	-0.1372	-0.0572	0.0361	-0.0178	0.0088
	(0,0289)	(0,0315)	(0,0851)	(0,076)	(0,0364)	(0,0315)	(0,0475)	(0,0488)	(0,0962)	(0,0892)	(0,0189)	(0,026)
Leisure	-0.0998	0.0580	-0.1030	0.0137	-0.1330	0.0425	0.2283	-0.4348	-0.0799	0.5710	0.0124	-0.0833
	(0,0729)	(0,0831)	(0,217)	(0,1912)	(0,093)	(0,0788)	(0,1235)	(0,1258)	(0,2397)	(0,2322)	(0,0493)	(0,0679)
Transport	-0.2127	-0.1131	0.2542	-0.0193	-0.2973	-0.0321	0.0341	-0.0759	0.3361	0.6228	0.0528	0.0506
	(0,0814)	(0,0934)	(0,2453)	(0,227)	(0,1036)	(0,0869)	(0,1358)	(0,1494)	(0,2826)	(0,2608)	(0,0532)	(0,075)
Personal care	-0.1442	-0.1773	0.2815	-0.1231	-0.3458	-0.0120	-0.0013	0.2300	0.2718	0.2238	0.0072	0.0523
	(0,0747)	(0,0822)	(0,2194)	(0,2046)	(0,0985)	(0,0809)	(0,1281)	(0,1294)	(0,2423)	(0,2197)	(0,0503)	(0,0682)

Note: The parameters of the budget share correspond to specification [7].

Table 14 – Sample comparison

	Large sample				Restricted sample			
	Three goods		Eight goods		Three goods		Eight goods	
	Estimate	St. Err.	Estimate	St. Err.	Estimate	St. Err.	Estimate	St. Err.
Estimated share								
Pre 2009	0.4642	0.0793	0.5924	0.0486	0.4932	0.0836	0.6418	0.0537
Pre 2009 (1)	0.4940	0.0802	0.6247	0.0517	0.5408	0.0847	0.6739	0.0557
Post 2009 (2)	0.4956	0.0800	0.6261	0.0517	0.5416	0.0846	0.6746	0.0558
Post 2009	0.5033	0.0834	0.6088	0.0499	0.5172	0.0887	0.6344	0.0544
Parameters								
Constant	-0.0038	1.4502	-6.4232	22.2238	-2.1315	1.6919	-16.9838	23.4249
Reg. unempl. ratio	0.3831	0.1717	0.4365	0.2604	0.6105	0.2136	0.4557	0.2513
Age $\geq$ 35	0.1419	0.0854	0.1273	0.1286	0.1475	0.1050	0.1398	0.1553
High ed.	0.0200	0.0713	0.0567	0.1087	-0.0264	0.0826	0.0311	0.1269
Active, wife			-0.0371	0.0610			-0.0504	0.0533
Prices	YES		YES		YES		YES	
Region FE	YES		YES		YES		YES	
Estimated scales								
Wife: scale economies	0.6450	0.2006	0.5298	0.1093	0.6433	0.2177	0.5086	0.1120
Husband: scale economies	1.8011	0.6723	1.3048	0.2041	1.2618	0.4614	1.6734	0.3357
Parameters, wife								
Constant	1.2274	1.3975	-18.8175	15.7451	-0.0464	1.6615	-24.5118	17.8583
Age	0.0827	0.1765	0.1309	0.1176	0.1644	0.2024	0.1624	0.1403
Education	-0.0201	0.1661	0.0394	0.1047	-0.0331	0.1854	0.0428	0.1235
Rural	0.1411	0.1921	-0.0504	0.0913	0.1331	0.2035	-0.0183	0.0811
Madrid-Barcelona	-0.1603	0.1885	-0.0786	0.1139	-0.1157	0.2009	-0.0825	0.1002
Car owner	-0.3113	0.1377	-0.2550	0.0696	-0.3257	0.1486	-0.2428	0.0610
Home owner	0.1101	0.1830	0.1089	0.0924	0.0546	0.2050	0.1071	0.0813
Regional relative price	YES		YES		YES		YES	
Parameters, husband								
Constant	-2.6722	1.7205	7.0775	8.1043	-4.5251	1.9809	13.8204	7.7519
Age	-0.0749	0.2148	-0.0183	0.0484	-0.1062	0.2462	-0.0566	0.0493
Education	-0.1939	0.2118	-0.1611	0.0507	-0.2999	0.2383	-0.1902	0.0517
Rural	-0.0508	0.2512	0.1103	0.0745	-0.1826	0.3015	0.0792	0.0742
Madrid-Barcelona	0.2978	0.3460	0.2178	0.1086	0.3597	0.4006	0.2116	0.1022
Car owner	0.3669	0.2328	0.1127	0.0803	0.3326	0.2705	0.0839	0.0739
Home owner	-0.3151	0.2983	-0.1670	0.0955	-1.6952	0.5882	-0.1226	0.0899
Regional relative price	YES		YES		YES		YES	

Note: (1) Share computed with pre-2009 averages but post-2009 regional unemployment ratio. (2) Share computed with post-2009 averages but pre-2009 regional relative prices. Unless specified, the demographic characteristics are women's. Men to women unemployment ratio. Women to men wage ratio.

Table 15 – Diff-in-diff estimates of the scale economies

	Model with 3 goods						Model with 8 goods		
Estimated	[1]	[2]	[3]	[4]	[5]	[6]	[7]	[8]	[9]
Estimated scales									
Wife: scale economies	0.6609	0.642	0.6669	0.6552	0.6349	0.6606	0.5518	0.5632	0.5717
	(0.2183)	(0.2165)	(0.2231)	(0.2176)	(0.2155)	(0.2225)	(0.1185)	(0.1205)	(0.122)
Husband: scale economies	2.1022	2.161	2.0786	2.1209	2.1855	2.0845	1.3881	1.398	1.3143
	(0.9153)	(0.955)	(0.916)	(0.927)	(0.9698)	(0.9237)	(0.2442)	(0.2602)	(0.2357)
Parameters, wife									
Constant	0.7062	0.6449	0.6904	0.7295	0.6699	0.6563	-25.9561	-24.236	-24.1246
	(1.5025)	(1.5095)	(1.5021)	(1.5109)	(1.5197)	(1.4991)	(17.1223)	(16.9373)	(16.8795)
Age $\geq$ 35	0.1251	0.1285	0.1234	0.1259	0.1299	0.1223	0.133	0.1315	0.1309
	(0.1899)	(0.1912)	(0.1896)	(0.1906)	(0.192)	(0.1899)	(0.1297)	(0.1291)	(0.1258)
High ed.	0.0303	0.0329	0.0295	0.0315	0.0343	0.0318	0.0296	0.0256	0.0216
	(0.1759)	(0.1768)	(0.1754)	(0.1764)	(0.1775)	(0.1752)	(0.1132)	(0.1125)	(0.109)
Rural	0.1243	0.117	0.1246	0.1249	0.118	0.1191	-0.0252	-0.0307	-0.0348
	(0.1958)	(0.1955)	(0.1953)	(0.1965)	(0.1965)	(0.1946)	(0.0855)	(0.0836)	(0.0853)
Madrid	-0.0366	-0.0286	-0.0389	-0.0348	-0.0263	-0.034	-0.0638	-0.0627	-0.0647
	(0.1842)	(0.1863)	(0.1836)	(0.1852)	(0.1877)	(0.1838)	(0.0996)	(0.097)	(0.099)
Car owner	-0.3107	-0.3205	-0.3089	-0.3118	-0.322	-0.3142	-0.2439	-0.2435	-0.2482
	(0.1446)	(0.1461)	(0.1439)	(0.1451)	(0.1468)	(0.144)	(0.0657)	(0.0641)	(0.0663)
Home owner	0.0969	0.1017	0.0952	0.0977	0.1028	0.0944	0.1135	0.1134	0.114
	(0.1933)	(0.1963)	(0.1925)	(0.1942)	(0.1975)	(0.1933)	(0.0879)	(0.0857)	(0.0876)
Prices	-1.0361	-0.9935	-1.0091	-1.0717	-1.0337	-0.9774	2.9725	2.6193	2.6974
	(1.6833)	(1.6904)	(1.6883)	(1.6949)	(1.7043)	(1.6843)	(3.5958)	(3.5471)	(3.5385)
Parameters, husband									
Constant	-1.5707	-1.5517	-1.5822	-1.5667	-1.5414	-1.6268	16.9766	15.6459	16.0298
	(1.956)	(1.9508)	(1.9639)	(1.9481)	(1.9402)	(1.9711)	(9.0087)	(8.8841)	(9.3359)
Age *geq* 35	-0.1374	-0.1376	-0.1371	-0.1368	-0.1367	-0.1338	-0.0695	-0.0676	-0.0692
	(0.2254)	(0.2243)	(0.226)	(0.2252)	(0.224)	(0.2263)	(0.0518)	(0.0517)	(0.0528)
High ed.	-0.231	-0.236	-0.2298	-0.2321	-0.2373	-0.231	-0.1619	-0.1671	-0.1584
	(0.219)	(0.2179)	(0.2195)	(0.2186)	(0.2174)	(0.2192)	(0.0544)	(0.0546)	(0.0553)
Rural	-0.0522	-0.0523	-0.0519	-0.0524	-0.0525	-0.0516	0.0935	0.103	0.1066
	(0.2588)	(0.257)	(0.2595)	(0.2582)	(0.2562)	(0.2593)	(0.0769)	(0.0771)	(0.0784)
Madrid	0.1584	0.1546	0.1601	0.1579	0.1543	0.1588	0.1964	0.1998	0.1997
	(0.3627)	(0.3628)	(0.3632)	(0.3622)	(0.3622)	(0.3643)	(0.1088)	(0.1086)	(0.112)
Car owner	0.3442	0.3503	0.3422	0.345	0.3513	0.3434	0.0867	0.0977	0.0987
	(0.2454)	(0.2444)	(0.246)	(0.245)	(0.2439)	(0.2458)	(0.0806)	(0.0817)	(0.0851)
Home owner	-0.2159	-0.2098	-0.2175	-0.2134	-0.2073	-0.2147	-0.1417	-0.1505	-0.1504
	(0.2986)	(0.2953)	(0.3)	(0.2974)	(0.2939)	(0.2994)	(0.0952)	(0.0958)	(0.099)
Prices	YES	YES	YES	YES	YES	YES	YES	YES	YES
Prices	YES	YES	YES	YES	YES	YES	YES	YES	YES

Note: Standard errors in parenthesis.