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WEALTH INEQUALITY AND HOUSEHOLD STRUCTURE: U.S. VS. SPAIN

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We study the link between culturally inherited household structure and wealth distribution in international comparisons using household data for the U.S. and Spain (the SCF and the EFF). We estimate counterfactual U.S. distributions relying on the Spanish household structure. Our results show that differences in household structure account for most of the differences in the lower part of the distribution between the two countries, but mask even larger differences in the upper part of the distribution. Imposing the Spanish household structure to the U.S. wealth distribution has little effect on the Gini coefficient and wealth top shares. However, this is the net result of reduced differences at the bottom and increased differences at the top. So there is distinct additional information in considering the whole distribution. Finally, we present results for the within-group differences between the two countries using quantile regressions and find a reversing pattern by age.

1. Introduction and Summary

Differences in wealth distribution across developed countries are large. The estimated share of the nation's wealth held by the top 1 percent of the population, an often cited distributional measure, may vary from 15 to 35 percent.[1] Documenting these differences is important in at least two different contexts.

Firstly, distributional comparisons of net worth are obviously of interest in the literature on inequality measurement. Such interest comes from the fact that real and financial marketable assets can be readily used for consumption smoothing and intergenerational transmission. The quality of wealth data based on household surveys available in many countries is such that international comparisons of wealth distributions are now feasible.[2]

Secondly, the nature of these differences may help to discriminate between alternative economic theories of the distribution of wealth. The literature on computable general equilibrium models has tried to develop theories of saving behavior that can endogenously produce the form of distribution encountered in wealth data, given household-specific shocks from an exogenous earnings process. Since the basic models fail to account for the facts, additional features have been

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[1]See, for example, the evidence in Davies and Shorrocks (2000).

[2]See Bover *et al.* (2005) for a comparison, based on micro data, between Italy, the U.K., the U.S., and Spain, using harmonized definitions of asset holdings from individual-country household surveys and the recently created LWS database with wealth surveys from many countries.

considered in the literature.[3] Understanding international differences can be important for establishing which of these features matter if the features themselves are associated with institutional differences across countries in, for example, business regulations, welfare programs, bequests, or taxation.

However, the influence of differences in household structure on cross-country comparisons may be important. For example, if two countries differ in the pattern of household formation by young adults, not only the age distribution of households will differ but also the distribution of household size and type. This raises the question to what extent the differences we observe in wealth distributions across countries persist for comparable households, and to what extent they are due to differences in household structure between countries. This is important to elucidate because wealth magnitudes in micro surveys are usually measured for households as opposed to individuals and the economic interpretation of the disparities in the distribution can be very different in one case and the other. From an equity point of view differences due to family demographics should probably be netted out.

Previous work on international comparisons treated households as homogeneous across countries (except when trying to equivalize wealth by the number of household members). This could be a good strategy when comparing countries such as the U.K. and the U.S. where demographic structure may be relatively similar.[4] However, for general cross-country comparisons, taking into account differences in household structure becomes a more important consideration.

Table 1 shows some characteristics of the wealth distributions for the U.S. and Spain, the countries we consider in this paper. It is noticeable that the sizeable differences in the summary measures for all households are considerably reduced when comparing more homogeneous demographic groups, as for example households with head aged 35 to 54 and living with a partner.

The age at which young people leave the parental home to establish their own household is one key reflection of long-standing differences in family systems between Western countries. Other indicators are the prevalence of lone parent households or of elderly persons living with their children. The sociological literature (see Reher, 1998) identifies two clearly different geographical areas regarding family systems, one where family ties are strong (mostly Mediterranean countries) and another where they are weak (Northern Europe and the U.S.). In the former, children tend to leave home coinciding with marriage and may save up until then, while in the latter they settle for an independent life as they reach maturity. These differences exist at least since the 17th century when the earliest data are available. According to the first modern censuses, in the mid 19th century in Northern Europe between 30 and 55 percent of 15–24-year olds of both sexes would leave the parental home and be servants with another family, while only 5 to 20 percent of them would do so in Southern Europe where family labor was much preferred. The

[3]For useful summaries on the recent literature, see, for example, Quadrini and Ríos-Rull (1997) and Cagetti and De Nardi (2006). For earlier models, see Atkinson and Harrison (1978) and Jenkins (1990).

[4]Nevertheless, Banks *et al.* (2003) condition on three age bands when conducting part of their U.K. vs. U.S. comparison. Hyslop and Maré (2005) point out the influence of changes in household types on the increase in income inequality in New Zealand since the 1980s.

TABLE 1

Summary Statistics for U.S. and Spanish Wealth Distributions, All and Selected Groups

	Gini	Median[1]	p75/p25	No. of Observations
All households				
U.S.	0.80	66	22.7	4,442
Spain	0.56	102	4.3	5,143
Households with head aged 35 to 54				
U.S.	0.77	79	13.6	1,994
Spain	0.54	114	3.8	1,717
Households with head aged 35 to 54 and couple				
U.S.	0.74	118	8.1	1,427
Spain	0.52	121	3.6	1,293
Households with head aged 35 to 54, couple, one child <16				
U.S.	0.74	121	8.1	297
Spain	0.50	118	3.5	417
All households, using square root equivalence scale ($\sqrt{\text{no.}}$ of hh members)				
U.S.	0.80	45	22.5	4,442
Spain	0.56	62	4.3	5,143
All households, per capita (scaling by no. of hh members)				
U.S.	0.81	31	22.5	4,442
Spain	0.58	37	4.5	5,143

Note: [1]In thousands of euros.
Source: United States: Survey of Consumer Finances (SCF) 2001.
Spain: Spanish Survey of Household Finances (EFF) 2002.

factors shaping up these differences could be partly traced to the Germanic vs. Muslim and Oriental influence, the Reformation in contrast to Catholicism, and the earlier and more profound effects of the Industrial Revolution in Northern Europe (Reher, 1998). Moreover, despite the fact that some convergence has occurred lately, a clear divide remains. In Table 2 we report, for several Western countries, the proportion of single person households, of lone parents families, and of 25–29-year olds still living with their parents. The divide between Northern and Southern countries is clear, and at the extremes we observe Sweden with 44 percent of single-person households and Spain with 16.9 percent.

In this paper we argue that the prevailing family systems in each country are important to understand differences in wealth inequality between countries. We study the implications of the differences in family structure for the comparison of wealth inequality between two countries, one with weak family ties, the U.S., and another with strong family ties, Spain. Moreover, for these two countries we have quite comparable wealth micro data (the SCF2001 and the EFF2002, respectively). We believe this approach could be useful more generally when comparing wealth data across countries.

We take cross-country differences in family structure as given. If these differences were endogenously determined to first-order by differences in wealth, our results, though still valid from a descriptive point of view, would be less informative. Marriage and divorce decisions are known to be influenced by economic motives (Becker, 1973). Moreover, recent work by Guner and Knowles (2004) has

TABLE 2

Household Types: Indicators for Some Western Countries

	% of Single Person Households (1990/1991)[1]	% of Lone Parent Families (of fam. with children <18) (1989/1991)[2]	% Aged 25–29 Still Living with Parents (1994)[3]	
			Men	Women
Sweden	44.0	22.3	–	–
Denmark	38.1	22.0	–	–
Netherlands	37.7	18.1	–	–
Germany	37.7	15.7	28.8	12.7
U.K.	30.0	19.4	20.8	10.8
U.S.	29.2	23.5	15.6	8.8
France	29.2	11.9	22.5	10.3
Italy	23.7	–	66.0	44.1
Greece	21.1	–	62.6	32.1
Spain	16.9	8.6	64.8	47.6

Source:

[1]Reher (1998) from Eurostat for Europe using census and register data; CPS U.S. Census Bureau from Census 2000. More recent figures on single person households are available for European countries from the ECHP which provide figures broadly comparable to those shown.

[2]Fernández-Cordón and Tobío-Soler (1998) from INSEE; Bureau of Labor Statistics.

[3]Fernández-Cordón (1997) from Eurostat for Europe; CPS U.S. Census Bureau.

considered a general equilibrium model of the joint determination of marriage, divorce, and household savings, and compares its predictions with those from more traditional macro models with exogenous marriages. In contrast, the motivation of this paper is in conditioning on slow-moving aspects of household structure, possibly generated by values or social norms. We emphasize the more exogenous fact that young adults leaving their parents' home at a later age in Spain (and other Southern type countries) than in the U.S. implies that certain types of households are rarer in Spain (and others more abundant). It is noteworthy that strong family ties may override to a large extent divorce outcomes. Reher (1998) describes that despite an increase in lone-parenthood in recent years everywhere due to divorce and teenage pregnancies, there continue to be important differences in the levels between North- and South-type countries. In Spain, around 30 percent of all lone mothers with children co-reside with their own mother (Reher, 1998), while only 15 percent of single mothers live with their parents in the U.S. (London, 1998).[5]

To assess the impact of household structure on the differences in wealth distribution between the U.S. and Spain, we estimate non-parametrically the counterfactual distribution that would have prevailed in the U.S. had the demographic characteristics of households been those prevailing in Spain. Following DiNardo *et al.*'s (1996) study for earnings, we assess graphically what part of the differences are attributable to differences in household structure for the entire wealth distribution.[6] However, in contrast to DiNardo *et al.*, our main instrument

[5]These figures refer to 1991 for both countries.

[6]Morissette *et al.* (2006) for Canada and D'Ambrosio and Wolff (2006) for the U.S. use this methodology to study the changes in wealth inequality over time.

of analysis is the evaluation of counterfactual cumulative distribution functions rather than counterfactual densities.[7] An advantage of relying on conditional distributions is that one avoids having to choose a smoothing method. It is well known that density estimation is sensitive to the smoothing method adopted. This is particularly relevant in the case of wealth distributions, which often have a marked spike at zero because a non-negligible proportion of the population has no wealth. The presence of spikes increases the sensitivity of density estimations to the smoothing method used.

Furthermore, from the estimated counterfactual U.S. distribution, we easily derive summary counterfactual distribution measures and compare them to the actual measures for the U.S. and Spain. Using them, we can decompose the difference between the two countries in measures of position and dispersion into a part due to differences in household composition and another part holding household composition constant. We also compare concentration measures. One interesting result of the paper is that imposing the Spanish household structure to the U.S. wealth distribution has little effect on summary measures of inequality. However, this is the net result of reduced differences at the bottom and increased differences at the top. There is therefore distinct additional information in considering the whole distribution. We check the robustness of our results by controlling for education.

Finally, it is also of interest to study in some detail the distributional differences between the U.S. and Spain for given household types. To do so, we present saturated quantile regressions pooling the data for the two countries. We also provide plots of the within-groups wealth distributions for the different household types we consider in the paper.

The paper is organized as follows. In Section 2 we present the methodological framework. We discuss counterfactual wealth distributions from descriptive and structural perspectives, the role of conditioning, and weighting estimation methods. In Section 3 we describe the data and the classification of household types that we adopt. We also discuss the important role of oversampling for our exercise. In Section 4 we derive the counterfactual U.S. wealth distribution using the Spanish structure of households, and compare it graphically with the two factual distributions. To further characterize the differences between the two countries we also look at portfolio composition and debt. In Section 5 we summarize the differences in the three distributions using measures of position, dispersion, and concentration, and quantify how much of the differences are due to household composition. We also identify which particular types of households contribute most to the estimated compositional differences. In Section 6 we present alternative specifications, and check the robustness of our results to alternative household classifications and to controlling for education. We also present the counterfactual distribution for Spain and discuss the results. In Section 7 we provide information about the differences in wealth distribution for given household types. Section 8 concludes.

[7]As noted by Biewen (2001), it is also much easier to implement.

2. Objectives and Methodology

Accounting Decompositions

Let us consider two populations, 1 and 2 (Spain and the U.S.), and J household types. The wealth distributions of each country satisfy:

$$F_1(r) = \sum_{i=1}^{J} E_1[1(W \le r)|z = j]\, p_1(j)$$

$$F_2(r) = \sum_{i=1}^{J} E_2[1(W \le r)|z = j]\, p_2(j)$$

where $p_\ell(j) = \Pr_\ell(z = j)$, $1(.)$ represents the indicator function, and $E_\ell(.)$ denotes an expectation taken with respect to some distribution in population $\ell = 1,2$.

In a comparison between $F_1(r)$ and $F_2(r)$ the following accounting decomposition may be useful:

$$(1) \qquad F_2(r) - F_1(r) = \sum_{i=1}^{J} \{E_2[1(W \le r)|z = j] - E_1[1(W \le r)|z = j]\}\, p_1(j) + \sum_{i=1}^{J} E_2[1(W \le r)|z = j][p_2(j) - p_1(j)].$$

The first term is the part of the difference that can be attributed to differences in conditional wealth distributions, whereas the second is due to differences in the relative importance of each group between the two populations.

The first term is obtained by subtracting $F_1(r)$ to an artificial *cdf*, which combines the conditional *cdf*s of the second population with the group weights of the first population. Namely,

$$(2) \qquad F_2^C(r) = \sum_{i=1}^{J} E_2[1(W \le r)|z = j]\, p_1(j).$$

An object such as $F_2^C(r)$ is a useful descriptive summary for comparing the conditional distributions between the two populations because the weights are kept constant. Beginning with DiNardo *et al.* (1996), it has been much used in conjunction with nonparametric or flexible estimates.

Exogeneity and Wealth Functions

Under the assumption of exogeneity of household structure in population 2, $F_2^C(r)$ is a counterfactual *cdf* in the sense that if population 2 had the same household structure as population 1, then the distribution of wealth would be $F_2^C(r)$. In a similar symmetric sense, $F_1^C(r)$ would be a counterfactual *cdf* under the assumption of exogeneity of household structure in population 1.

To better understand the meaning of exogeneity in this context, we need to relate the conditional *cdf*s of wealth $F_1(r|z)$ and $F_2(r|z)$ to models of cross-sectional wealth determination at each population, W_1 and W_2 say. By construction:

$$F_1(W_1|z) = U_1$$

$$F_2(W_2|z) = U_2$$

where U_1 and U_2 are uniformly distributed variables independent of z. The random vector z includes age and descriptors of household type, such as the presence of children, or whether the household is formed by a couple or a single adult. Inverting the relationships we obtain the quantile functions:

$$W_1 = q_1(z, U_1)$$

$$W_2 = q_2(z, U_2).$$

If we think of these functions as economic models of individual wealth determination, some points can be noted. First, the unobservable component will include initial physical and human wealth, and accumulated income, health, and other shocks. Second, we would expect household type to have a structural effect on the ability to save and the process of wealth accumulation, through the operation of within-household insurance, and externalities in consumption and time devoted to non-market work (e.g. single parent households or households with external dependants as opposed to intact couples or parents with adult children living at home). Thirdly, the functions q_1 and q_2 will not only depend on preference and discount parameters, but also on institutional differences between the two populations, which are constant in a cross-section (such as access to credit and taxation). Thus even if preferences across populations are the same we would expect q_1 and q_2 to differ.

Here we argue that household structure is not just the result of many independent individual decisions, but something that is heavily influenced by prevailing practices and custom, as evidenced by large differences observed across countries, and therefore to some extent predetermined to individuals. Also, divorce decisions are not independent of expected re-marriage opportunities.

There are nevertheless a host of reasons why family structure may be influenced by unobservable wealth determinants. Our claim is just that we do not expect cross-country variation in household structure to be driven by reverse causality from wealth shocks. Therefore, we regard a statistical decomposition such as (1) as being not only useful from a statistical-descriptive point of view, but also as capable of providing suggestive evidence about the link between household structure and wealth inequality.

$F_2^C(r)$ is the *cdf* of a re-weighted population of country 2 households to make them comparable to country 1. In contrast, $F_1^C(r)$ is a *cdf* of a population of country 1 households re-weighted for comparability with country 2. In this paper we take the U.S. as the reference, but also report some results with Spain as the reference for comparisons. Arguably, the U.S. has a larger diversity of household types than Spain. This situation makes it easier to obtain a subsample (or a re-weighted sample) of U.S. households that are comparable to Spanish households than the opposite.

Conditioning

Even if our focus is on household structure, potential selection or endogeneity problems may be mitigated by further conditioning on observables. In doing so a distinction should be made between conditioning and "counterfactualizing." Suppose we consider counterfactual distributions of the form (2) for given values of some conditioning variable X:

$$F_2^C(r|X) = \sum_{i=1}^{J} E_2[1(W \leq r)|z = j, X]\, p_1(j|X).$$

Next we can consider a marginal counterfactual *cdf* by integrating with respect to the distribution of X in population 2:

$$F_2^{C_X}(r) = \int F_2^C(r|X)\, dG_2(X).$$

Under the assumption of conditional exogeneity of household structure in population 2, $F_2^{C_X}(r)$ is a counterfactual *cdf* in the sense that if population 2 had the same conditional household structure given X as population 1, then the distribution of wealth would be $F_2^{C_X}(r)$. Note that $F_2^{C_X}(r)$ is different from a counterfactual distribution with respect to both z and X.

The rationale for conditioning rather than counterfactualizing can be illustrated in a simple example as follows. Suppose we are interested in the effect of Z (household structure) on W (wealth) controlling for X (education) using a linear regression from population 1: $E_1(W|Z,X) = \alpha Z + \beta X$. Thus, factual and counterfactual average wealth for education level X satisfy $E_{1,1}(W|X) = \alpha E_1(Z|X) + \beta X$ and $E_{1,2}(W|X) = \alpha E_2(Z|X) + \beta X$, respectively. To examine the effect of Z on unconditional average wealth while keeping education constant, we can construct a counterfactual mean wealth by averaging $E_{1,2}(W|X)$ in population 1: $E_2^C(W) = \alpha E_1[E_2(Z|X)] + \beta E_1(X)$. By conditioning on X we make sure that α is not a spurious effect due to correlation between Z and X. Finally, note that a counterfactual average wealth with respect to both Z and X is $E_2^{CC}(W) = \alpha E_2[E_2(Z|X)] + \beta E_2(X)$. However, unlike $E_2^C(W)$, $E_2^{CC}(W)$ is not a counterfactual average associated to *only* changing household structure but to a joint change in household structure and education.

Estimating Counterfactual Distributions

Before discussing estimation of counterfactual distributions, we point out a generalized version of the law of iterated expectations, which considerably simplifies the calculation of counterfactual quantities by avoiding conditional distributions. First note that by the law of iterated expectations we have

$$F_2(r) = E_2\{E_2[1(W \leq r)|z]\} = E_2[1(W \leq r)].$$

This simply means that in order to calculate F_2(r) there is no need to calculate a weighted average of the conditional distributions, but that a simple average of individual indicators suffices. A similar simplification is available for calculating counterfactual distributions. We have[8]

$$F_2^C(r) = E_1\{E_2[1(W \leq r)|z]\} = E_2\left[1(W \leq r)\frac{p_1(z)}{p_2(z)}\right].$$

The random variable $R(z) = p_1(z)/p_2(z)$ is sometimes called a Radon–Nikodym derivative of population 1 with respect to population 2.[9]

Thus, given a sample $\{W_i, z_i\}_{i=1}^{N_2}$ from population 2 and an estimated weight function $\hat{R}(z) = \hat{p}_1(j)/\hat{p}_2(j)$, we can consider estimates of the counterfactual distribution of the form[10]

$$\hat{F}_2^C(r) = \frac{1}{N_2}\sum_{i=1}^{N_2} 1(W_i \leq r)\hat{R}(z_i).$$

Reweighting is a familiar approach in the treatment effects literature. Its use in a decomposition analysis has been suggested in Barsky *et al.* (2002).

If z_i is discrete and $\hat{p}_2(j) = N_2^{-1}\sum_{i=1}^{N_2} 1(z_i = j)$, then $\hat{F}_2^C(r)$ is numerically identical to

$$\sum_{i=1}^{J} \widehat{\Pr}_2(W \leq r|z = j)\hat{p}_1(j)$$

where $\widehat{\Pr}_2(W \leq r|z = j)$ denotes an empirical *cdf* conditioned on $z = j$. However, alternative estimators of $R(z)$ are possible, including those constructed from complementary datasets.

The point here is that explicit calculation of $\widehat{\Pr}_2(W \leq r|z = j)$ is not needed. Aside from a (substantial) computational simplification, this is relevant in our context because surveys of household wealth are expensive and therefore sample size is typically not very large. As a result conditional distributions may be difficult to estimate with sufficient precision. Although outside the scope of this paper, estimates of $R(z)$ from complementary datasets offer the possibility of considering a larger number of cells than would be feasible from wealth surveys alone. Complementary datasets would be representative samples from populations 1 and 2, which contain information on household characteristics (z) but not necessarily on wealth.

In our context the reweighting formulation is critical to our ability to obtain bootstrap standard errors of characteristics of counterfactual wealth distributions

[8]When z is discrete, as in our case, $p_\ell(j)$ is a probability mass function, but the same result holds for continuous z reinterpreting $p_\ell(j)$ as a density function.

[9]Note that $E_2[R(z)] = 1$ and $Cov_2[W, R(z)] = E_2^C(W) - E_2(W)$.

[10]In this section we abstract from issues related to stratification, which are discussed in the next two sections.

without an excessive computational burden. The reweighting formula also makes clear that a household group with few observations may have an impact on inference about counterfactual marginal distributions, but not because of lack of precision in conditional distributions (which are not used in the calculation) but only through sample error in estimated weights.

In the case of the counterfactual *cdf* $F_2^{C_X}(r)$ obtained by re-weighting conditionally on X the result is

$$F_2^{C_X}(r) = E_2\left[1(W \le r)\frac{p_1(z|X)p_2(X)}{p_2(z|X)p_2(X)}\right] = E_2[1(W \le r)R(z, W)]$$

where $R(z, W) = p_1(z|X)/p_2(z|X)$, or using Bayes formula

$$R(z, W) = R(z)\frac{p_1(X|z)p_2(X)}{p_2(X|z)p_1(X)}.$$

Note that if we were counterfactualizing with respect to both z and X then the appropriate change of measure would be

$$R^*(z, W) = R(z)\frac{p_1(X|z)}{p_2(X|z)}.$$

In our application when conditioning on educational categories, we are able to estimate $R(z, W)$ from cell-specific sample frequencies. However, again flexible estimators from semiparametric models of $p_\ell(X|z)$ or complementary datasets would be possible alternatives.

3. Data and Demographic Groups

The data come from the U.S. Survey of Consumer Finances (SCF) 2001 and the new Spanish Survey of Household Finances (EFF) 2002.[11] The focus of both surveys is to collect rich information on household assets and debts together with socioeconomic variables relative to households and their members. An important feature for a wealth survey that they have in common is that the wealthy are oversampled. We construct comparable assets and debt definitions from the variables in both surveys.

The Data

In its current form the SCF has been carried out in the U.S. by the Board of Governors of the Federal Reserve System since 1989 every three years. The sample has a dual frame design. One part is drawn using a standard area probability

[11]See Aizcorbe *et al.* (2003) for an overview of the 2001 SCF and Kennickell (2005) for details on the SCF sample and methodology. For a full description of the 2002 EFF, see Bover (2004).

design. The other is selected using income tax returns in order to oversample wealthier households through a mapping from income components to an estimate of wealth. Both parts of the sample are then joined and overall sample weights estimated. The SCF collects detailed information on all household assets. These include: owner occupied residence, other real estate, businesses, vehicles, all types of financial assets (bank accounts, fixed income securities, listed and unlisted shares, managed accounts and trust fund equity, mutual funds, pension schemes), and debts. It also collects income, demographics, employment history and more recently some consumption measures. The questionnaire is conducted through a computer assisted personal interview and the median length is approximately 75 minutes.

The EFF was first conducted in 2002 by the Banco de España and every three years since then. From its start it was designed to obtain wealth magnitudes that could be compared to the ones in the SCF but the methods and questionnaire were adapted to fit Spanish circumstances. In Spain there is a wealth tax and it is on individual wealth tax files information that the EFF oversampling is based. The wealth tax information is linked to census information providing the EFF with a unique population frame for its sample. The questionnaire takes into account extended families living together, thus increasing the number of questions on labor status and income. On the other hand, household wealth in Spain is predominantly in real estate and therefore there is a less detailed financial asset categorization in the EFF as compared to the SCF. In particular, bank accounts, fixed income securities, listed and unlisted shares, mutual funds, and pension schemes are all covered in the EFF, but questions on each particular asset as opposed to asset category are collected only for mutual funds and pension schemes.[12] Detailed questions about all sorts of debts are included in the questionnaire, except education debt, which is non-existent in Spain, and credit card debt, which is only included in the EFF since 2005 when this type of debt started to be used by households. The CAPI questionnaire for the EFF takes approximately 60 minutes (median) to be completed.

Finally, a relevant issue for comparability of results is how item non-response is handled.[13] In that respect both surveys provide multiple imputations based on similar imputation techniques and programs to impute missing values.

Measure of Wealth

The wealth measure we use throughout this paper is net worth defined as non-human assets minus debts. Assets include financial assets, pension wealth, main residence and other real estate wealth, business equity, vehicles and jewels, and other comparable valuables.[14] All assets (including small businesses) are valued at market prices. Debts include all kinds of outstanding debts. All

[12]Managed accounts and trust fund equity in the 2002 EFF would be included in the "Other financial assets" category.

[13]Item non-response refers to the household failing to answer some questions in an otherwise completed interview.

[14]Spanish Social Security pension provisions and U.S. employer-sponsored defined-benefit plans are excluded from this measure. Defined contribution pension plans are covered in both surveys.

TABLE 3

Household Types in the U.S. and Spain, by Demographic Characteristics (%)

	U.S.	Spain
Couple vs. single		
Couples	60	71
Single male	14	10
Single female		
No children	18	17
With children	8	2
Age of household head		
<25	6	2
25≤ <35	17	12
35≤ <55	43	42
≥55	34	44
Presence of children under 16		
No	66	69
Yes	34	31

Source: SCF 2001 and EFF 2002.

monetary amounts are expressed in 2002 euros and have been adjusted for inflation in the U.S. and for purchasing power parity for 2002.[15]

This is a measure of marketable wealth, as opposed to conceptually wider measures that would include human wealth or Social Security type pension entitlements. In contrast with income, marketable wealth comparisons provide information on differences in consumption smoothing possibilities over the life-cycle (especially when households are subject to liquidity constraints) and in the scope for intergenerational transfers and inheritances.[16]

We checked that our results are not driven by potential differences in the definition of the unit of analysis in the two surveys. To this end we experimented with alternative definitions, taking into account the information provided for the U.S. on the wealth of household members who are outside the primary economic unit but share the same residence. Our results are unchanged.

Demographic Groups

In Table 3 we see that in the data used we observe the differences pointed out in the previous section: more single person households in the U.S. (40 vs. 29 percent), more lone parent households, in particular in the case of single female parents (8 percent in the U.S. vs. 2 percent in Spain, the percentage of single male parents being very small in both countries). Moreover, the larger proportion of households headed by young people in the U.S. is also clear.

To characterize the structure of households in both countries, we consider 16 types of households which differ in the age of the household head, marital status,

[15]The figures adopted are 1.6 percent for the U.S. inflation in 2002 and 0.743 for the U.S.–Spain purchasing power parity in 2002. If instead of adjusting for purchasing power parity we adjust only for the exchange rate, the differences between Spain and the U.S. become smaller (larger) when U.S. wealth is below (above) Spanish wealth. The reason is that no allowance is made for higher U.S. prices.

[16]See the early emphasis on the marketability of wealth by the Royal Commission on the Distribution of Income and Wealth (1975) in Britain.

gender of the head of household in case of single households, and presence of children. The choice of groups is based on the differences in households structure between the two countries, as explained in the previous section, making sure that a sufficient number of observations is available for each group in each country. The 16 groups considered may be found in Table 4, which shows for each group its population share in both countries (columns 1 and 2) and the number of observations available in our data (columns 5 and 6). In Section 6 we present alternative characterizations for robustness analysis. In particular, for the over 54s we distinguish households with grown-up offspring living with their parents. We also consider a characterization merging households under 25 with those aged 25 to 34 as a way of avoiding groups with small sample sizes.

In this paper we take the differences in the mix of groups to reflect mainly differences in household formation and structure, but differentials in gender mortality across countries could also be thought to affect the share of single women households among those over 54. However, if we take for example the death rates of those born between 1930 and 1939 (i.e. aged 63 to 72 in 2002) at 63, male death rates are higher than female rates by a larger amount in Spain (0.0090) than in the U.S. (0.0068).[17] Therefore, gender mortality differences could not be behind the higher share of single women among households aged over 54 in the U.S. (29.7 percent) as compared to Spain (27.4 percent).

The U.S. is ethnically and culturally more heterogeneous than Spain. It is well known that race and religious attitudes correlate with demographic variables such as divorce rates or the number of children. Differences in demographic structures across countries may well be associated with ethnic, religious, or cultural differences. However, we believe there is a more direct association between wealth accumulation and household structure, operating, for example, through household economies of scale or household dissolution. Establishing a link between household structure and cultural or ethnic diversity is outside the scope of this paper.

Equivalence Scales

A related but separate concern is the use of equivalence scales for normalizing household wealth. Equivalizing wealth would require an intertemporal theoretical framework, an avenue which is not pursued here. In this paper we aim to estimate to what extent the demographic structure of households accounts for the differences we observe in wealth distributions, as opposed to trying to approximate personal wealth distributions (for an attempt on the latter, see Sierminska and Smeeding, 2005). We argue that the differences in household structure we consider are not just a question of size of household. In Table 1 we also report results when normalizing household wealth using a square root equivalence scale or per capita wealth. As we can see, these standardizations reduce the difference in a measure of position like the median (although by less

[17]Death rates by cohorts from the *Human Mortality Database*.

TABLE 4

Information on the 16 Household Groups Considered

	Percentage in Population		Median Net Wealth[1]		No. of Observ. in the Sample		% of Owner-Occupiers		
	U.S. (1)	Spain (2)	U.S. (3)	Spain (4)	U.S. (5)	Spain (6)	U.S. (7)	Spain (8)	Counterf. U.S. (9)
Overall	100	100	65.8	101.9	4,442	5,143	67.7	81.9	74.9
Age < 25									
1. Couple	2.4	0.6	5.8	12.0	78	18	21.0	41.7	
2. Single male	1.4	0.6	2.0	3.2	52	20	3.9	49.2	
3. Single female	1.8	0.4	0.3	6.5	57	18	11.7	49.4	
25 ≤ Age < 35									
Couple									
4. No children	3.4	4.0	34.5	71.0	121	98	56.4	79.5	
5. Children	6.9	5.4	26.0	70.2	242	149	63.8	73.9	
6. Single male	2.6	1.7	9.7	62.6	94	62	35.2	55.6	
Single female									
7. No children	1.9	1.1	6.1	30.4	72	47	25.4	53.3	
8. Children	2.4	0.3	1.8	10.8	89	10	25.1	59.6	
35 ≤ Age < 55									
Couple									
9. No children	12.0	12.0	118.6	130.0	560	486	81.4	83.4	
10. Children	16.0	20.9	117.5	116.1	867	807	83.3	83.3	
11. Single male	5.2	3.6	36.5	78.5	215	163	54.3	67.0	
Single female									
12. No children	5.4	3.9	25.0	108.1	203	190	51.2	78.9	
13. Children	4.2	1.3	11.7	68.4	149	71	48.6	65.9	
Age ≥ 55									
14. Couple	19.7	28.2	220.9	122.4	1,102	1,937	89.3	90.5	
15. Single male	4.4	3.8	85.0	86.1	191	284	75.4	77.1	
16. Single female	10.2	12.1	60.7	78.6	350	783	67.1	82.6	

Note: [1]In thousands of euros.
Source: SCF 2001 and EFF 2002.

TABLE 5

Precision of Wealth Distribution Measures: Oversampling vs. Random Sampling

	p10[1]	p25[1]	p50[1]	p75[1]	p90[1]	(p75 – p25)/p25	% of Wealth Held by Top 50%	20%	10%	5%	1%
U.S.											
Point estimate	0.04	9.7	65.8	221.1	562.7	21.7	97.1	82.2	69.0	56.9	32.1
Standard error with											
Oversampling	0.06	0.5	2.1	5.0	14.2	1.1	0.1	0.4	0.5	0.6	0.5
Random sample	0.08	0.8	2.9	7.4	24.5	1.7	0.2	1.3	2.2	3.0	4.0
Spain											
Point estimate	6.4	43.2	101.9	185.7	330.2	3.3	86.4	58.6	41.8	29.5	13.2
Standard error with											
Oversampling	1.0	2.0	2.8	3.3	10.3	0.2	0.5	1.0	1.3	1.5	1.6
Random sample	0.6	1.3	2.0	2.1	7.3	0.1	0.9	2.6	3.6	4.4	5.3

Notes: [1]In thousands of euros.

than in comparisons of demographically comparable households), but not the difference in measures of inequality.[18]

The Critical Role of Oversampling in International Wealth Comparisons

In Table 5 (second and fifth row) we report standard errors for most of the distribution measures we calculate in the paper. As we mentioned when describing the data, an important common feature of the SCF and the EFF is that in both surveys the wealthy are oversampled. This sampling feature is crucial for the precision of some wealth distribution statistics routinely reported. To illustrate this point we also report bootstrap standard errors that would have resulted from randomly sampling the U.S. and the Spanish populations (third and sixth row).[19] As can be seen, for some of the statistics the difference in precision is very substantial. For example, the 95 percent confidence interval for the percentage of wealth held by the top 1 percent of the population in the absence of oversampling is almost as large as the international variation in this figure of 20 percentage points reported in Davies and Shorrocks (2000).[20,21] In the absence of oversampling we believe international comparisons should place the emphasis on less extreme points of the distribution like quartiles or interquartile ranges.[22]

[18]Other normalizations, like wealth divided by number of adult members of the household (as chosen by Davies *et al.*, 2006) provide similar results.

[19]A total of 999 random samples are drawn from the U.S. population which was obtained by applying the SCF population weights to the SCF sample.

[20]Interestingly, the differences in the point estimates between Spain and the U.S. are large enough to be significantly different even if one uses the random sample standard error estimate for the SCF.

[21]The share of the top 1 percent for Spain is lower than the one reported in Davies *et al.* (2007) (13.2 vs. 18.3 percent), who reproduce estimates by Alvaredo and Saez (2007). However, Alvaredo and Saez estimates are based on wealth tax returns filed only by approximately the richest 4 percent. Owner occupied main residence, which is a very significant and widespread asset of Spanish households, falls below the wealth tax threshold for the majority of households. This may give rise to some overestimation of top shares based on this source.

[22]Cowell and Flachaire (2007) examine the statistical performance of inequality indices, including the Gini coefficient, and show that these are very sensitive to the presence of extreme values.

4. Counterfactual U.S. Wealth with Spanish Household Structure

Estimation of the Counterfactual U.S. Distribution

We first estimate the U.S. empirical wealth distribution using population weights ψ_i as follows:

$$\hat{F}_{US}(r) = \frac{1}{\sum_{i=1}^{N^{US}} \psi_i^{US}} \sum_{i=1}^{N^{US}} \psi_i^{US} 1\left(w_i^{US} \le r\right)$$

where N is the sample size. Similarly, we evaluate the empirical *cdf* of wealth for Spain, $\hat{F}_{SP}$.

The counterfactual U.S. distribution, i.e. the U.S. within-groups distribution with the Spanish household structure, can be calculated as the weighted average[23]

$$\hat{F}_{US}^{SP}(r) = \frac{1}{\sum_{i=1}^{N^{US}} \psi_i^{US}} \sum_{i=1}^{N^{US}} \psi_i^{US} 1\left(w_i^{US} \le r\right) \hat{R}_i$$

where

$$\hat{R}_i = \frac{\widehat{\Pr}_{SP}\left(z = z_i^{US}\right)}{\widehat{\Pr}_{US}\left(z = z_i^{US}\right)}, \quad \widehat{\Pr}_{US}(z = j) = \frac{1}{\sum_{i=1}^{N^{US}} \psi_i^{US}} \sum_{i=1}^{N^{US}} \psi_i^{US} 1\left(z_i^{US} \le j\right), \text{ etc.}$$

where $j = 1, \ldots, J$ denotes the different household types (in this case $J = 16$, see Table 4).

One aim of this paper is to evaluate to what extent the larger wealth inequality observed in the U.S. relative to Spain is due to differences in the structure of households between the two countries. To this end, we study whether the differences between the U.S. and Spain are reduced or amplified when the Spanish distribution is compared with a counterfactual U.S. distribution with the same structure of households.

An important component of household wealth, which differs markedly across countries, is owner-occupied housing. An illustrative and interesting example of the previous general method is to look at differences in the proportion of owner occupied housing. In the U.S., 68 percent of households own their main residence, while 82 percent do so in Spain. However, the differences across different types of households are substantial. In the U.S., house-ownership varies from 4 percent for single males aged under 25 to 89 percent for couples over 55. When weighting the U.S. shares of owner-occupiers for each household type (column 7 in Table 4) by the Spanish population probabilities for each group type (column 2 in Table 4), the counterfactual U.S. percentage of the population owning their main residence goes

[23]The counterfactual *cdf* can also be calculated as an explicit weighted average of conditional empirical *cdf*s $\hat{F}_{US}^{SP}(r) = \Sigma_{j=1}^{J} \widehat{\Pr}_{US}(w \le r \mid z = j)\widehat{\Pr}_{SP}(z = j)$.

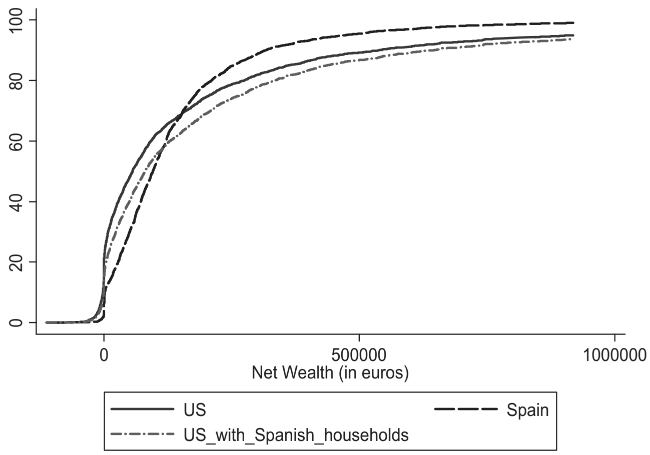

Figure 1. Empirical Wealth Distributions

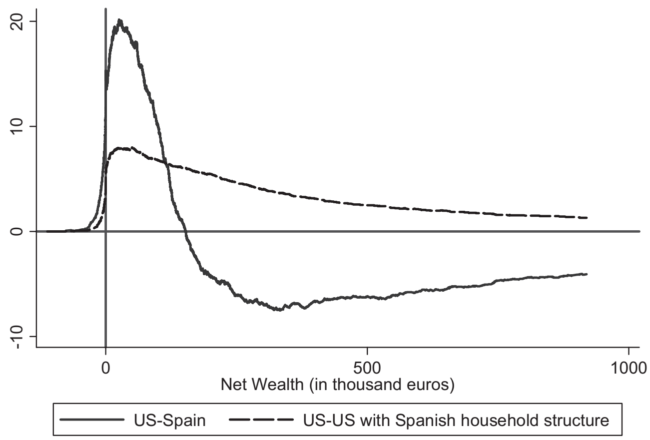

Figure 2. Difference between the Empirical Distribution Functions

up to 75 percent. Therefore, half of the difference in the proportion of owner-occupied housing between the U.S. and Spain could be attributed to differences in the types of households prevailing in both countries.

The empirical cumulative distribution functions for the U.S., Spain, and the counterfactual U.S. are plotted in Figure 1. The differences between the U.S. and Spain distributions and between the U.S. and the counterfactual U.S. are shown in Figure 2.[24] Household wealth in the U.S. is lower than in Spain up to approximately the 67th percentile. At this point the two distributions cross and the situation is reversed.

[24]The figures reflect wealth values up to 99 percent of the Spanish wealth distribution for the scale to be visually meaningful.

TABLE 6

Financial Assets: Wealth Share[1] and Participation Rates[2] (%)

	U.S.	Spain	U.S. with Spanish Mix of Households
Financial assets share	41.0	12.0	41.8
Percentage of households holding financial assets			
All financial assets (excluding bank accounts)	71.0	35.2	73.9
Stocks	21.7	12.5	24.3
Mutual funds	21.5	7.2	24.2
Fixed-income securities	18.9	1.9	20.6
Pension schemes	61.6	24.1	65.1

Notes:

[1]Wealth in financial assets (including bank accounts and deposits, stocks, mutual funds, fixed-income securities, and pension schemes) over wealth (including debts).

[2]Percentage of households holding various types of financial assets (excluding bank accounts and deposits).

These figures make it clear that there are considerably more households with zero or very low wealth in the U.S. as compared to Spain. However, differences in household structure explain a large part of the difference in wealth distributions, as the counterfactual U.S. distribution reveals. Indeed, the difference between the U.S. and Spain is greatly reduced when looking at the difference between Spain and the U.S. counterfactual up to approximately the 50th to 60th percentiles. For the first part of the distribution the counterfactual U.S. lies between the U.S. and the Spanish *cdf*s. In contrast, for the upper half of the distribution counterfactual U.S. wealth is higher than both the U.S. and the Spanish *cdf*s. This indicates that if the structure of households in the U.S. was as in Spain, the differences in household wealth between the two countries would be even larger than those observed for the upper half of the distribution. The likely explanation is that in Spain there is a higher frequency of households of the type that in the U.S. are high wealth (e.g. couples over 54, as we will see later).

To further characterize the difference between the two countries we look at portfolio composition and debt. The proportion of owner occupied housing by groups and the counterfactual U.S. rate are presented in Table 4. In Table 6 we report overall rates on the proportion of wealth invested in financial assets and the percentage of households that own financial assets (other than bank accounts and deposits).[25] The difference in the rate of home ownership between Spain and the U.S. is greatly reduced when Spain is compared to the counterfactual U.S. The same occurs when comparing debt shares but to a lesser extent when comparing percentage of indebted households.[26] In contrast, the share of financial assets in household wealth and the percentage of households holding various financial assets for the counterfactual U.S. are even higher than for the U.S.

[25]Bank accounts and deposits are held by 91 percent of households in the U.S. and 98 percent in Spain.

[26]Tables providing information about debt share and percentage of indebted households are available from the author upon request.

In the Appendix we provide plots for the three estimated wealth densities. These are dependent on the smoothing assumptions adopted but similar conclusions emerge.

5. Counterfactual Summary Measures

In this section we provide some measures to summarize the differences in the overall distributions. Counterfactual measures are used to decompose the total differences. Bootstrap standard errors are also provided.[27]

Differences in Measures of Position and Dispersion along the Distribution when only Household Structure Differs

From the previously estimated counterfactual distribution the calculation of percentiles is straightforward (e.g. the median, p50, is the smallest value of r for which $\hat{F}_{US}^{SP}(r) \le 0.5$. In the first three columns of Table 7 we report various measures of position and dispersion for the three distributions. In columns 4 to 9 (Table 7) we decompose the differences between the U.S. and Spain for the previous summary measures in the following way:

$$m_{SP} - m_{US} = \left(m_{SP} - m_{US}^{SP}\right) + \left(m_{US}^{SP} - m_{US}\right)$$

(m representing any of those measures). The first term reflects the difference in wealth for the same household composition and the second the differences when only household composition changes.

The figures in Table 7 reflect what was anticipated from looking at the graphs of the three cumulative distribution functions (Figure 1). Firstly, we can quantify to what extent applying the Spanish marginal probabilities to the within-groups U.S. wealth distribution reduces the observed differences in wealth distribution between the U.S. and Spain for the first part of the distribution (up to approximately the 60th percentile). For example, the percentage of households with zero or negative net worth for the U.S. would go from 9.6 to 6.4 percent. The role of household composition is the largest around the median where the U.S. median would increase from the actual 65,800 to the counterfactual 91,600, much closer to the Spanish 101,900 value when changing only household composition, reducing the difference with the U.S. by 71.5 percent. Furthermore, household composition accounts for 55 percent of the difference in inter-quartile range. In greater detail, it is more relevant for the difference between the median and the lower quartile (63 percent) than for the difference between the median and the upper quartile (13 percent).

However, for the upper part of the distribution the situation is reversed. Differences in household structure between the U.S. and Spain mask differences in the distribution of wealth between the two countries, which are larger under a common household composition. These differences are the largest around the 75th percentile. At that point, the Spanish magnitude is smaller than the U.S., which in

[27]Bootstrap standard errors are estimated using the 999 replicate weights available for the EFF and the SCF that allow taking into account stratification and clustering.

TABLE 7

Counterfactual Distribution Measures

	Distribution Measures			Decomposing the Differences						Alternative Specifications				
				Total Difference		Diff. Same hh Composition		Diff. Only hh Comp. Changes						
	U.S. m_{US} (1)	Spain m_{SP} (2)	Countf. U.S. m_{US}^{SP} (3)	$m_{SP} - m_{US}$ (4)	% (5)	$m_{SP} - m_{US}^{SP}$ (6)	% (7)	$m_{US}^{SP} - m_{US}$ (8)	% (9)	Control Educat. (10)	13 Groups (11)	18 Groups Adult ≥ 18 (12)	18 Groups Adult ≥ 25 (13)	Counterfactual Spain (14)
% households net worth ≤ 0	9.6	1.4	6.4	–8.2	100	–5.0	61.0	–3.2	39.0	7.1	6.6	6.4	6.3	1.7
	(0.3)[2]	(0.2)	(0.3)	(0.4)		(0.4)		(0.2)		(0.4)	(0.3)	(0.5)	(0.5)	(0.3)
p10[1]	0.04	6.4	1.7	6.3	100	4.6	73.4	1.7	26.6	1.3	1.5	2.0	1.9	2.9
	(0.1)	(1.0)	(0.2)	(1.0)		(1.0)		(0.2)		(0.3)	(0.2)	(0.6)	(0.6)	(0.6)
p25[1]	9.7	43.2	22.6	33.5	100	20.6	61.4	12.9	38.6	18.8	22.0	26.3	26.5	33.6
	(0.5)	(2.0)	(1.1)	(2.1)		(2.1)		(1.0)		(1.3)	(1.3)	(1.9)	(1.9)	(1.9)
Median[1]	65.8	101.9	91.6	36.1	100	10.3	28.5	25.8	71.5	84.6	90.8	96.1	98.6	91.6
	(2.1)	(2.8)	(2.7)	(3.4)		(3.6)		(2.2)		(3.4)	(2.6)	(6.7)	(6.5)	(2.7)
Mean[1]	299.8	160.4	367.3	–139.4	100	–206.9	148.4	67.5	–48.4	327.2	365.9	346.3	331.0	145.5
	(4.5)	(4.8)	(7.6)	(6.5)		(8.4)		(5.6)		(10.3)	(7.6)	(18.6)	(19.1)	(4.4)
p75[1]	221.1	185.7	282.9	–35.4	100	–97.1	274.5	61.7	–174.5	249.5	280.9	277.6	272.4	171.4
	(5.0)	(3.3)	(7.8)	(5.9)		(8.2)		(7.0)		(8.5)	(8.1)	(12.7)	(13.5)	(4.4)
p90[1]	562.7	330.2	664.0	–232.6	100	–333.8	143.5	101.3	–43.5	599.6	664.0	628.0	601.3	306.2
	(14.2)	(10.3)	(15.8)	(17.4)		(18.5)		(13.8)		(18.5)	(16.1)	(36.3)	(39.5)	(8.4)
$\frac{p75 - p25}{p25}$	21.7	3.3	11.5	–18.4	100	–8.2	44.6	–10.2	55.4	12.3	11.8	9.5	9.3	4.1
	(1.1)	(0.2)	(0.6)	(1.2)		(0.6)		(1.0)		(0.7)	(0.7)	(0.8)	(0.8)	(0.3)
$\frac{p50 - p25}{p25}$	5.7	1.4	3.0	–4.3	100	–1.6	37.2	–2.7	62.8	3.5	3.1	2.6	2.7	1.7
	(0.3)	(0.1)	(0.2)	(0.3)		(0.2)		(0.3)		(0.2)	(0.2)	(0.3)	(0.3)	(0.1)
$\frac{p75 - p50}{p50}$	2.3	0.8	2.1	–1.48	100	–1.28	86.5	–0.2	13.5	1.9	2.1	1.9	1.7	0.9
	(0.1)	(0.0)	(0.1)	(0.1)		(0.1)		(0.1)		(0.1)	(0.1)	(0.2)	(0.2)	(0.0)
$\frac{p90 - p50}{p50}$	7.5	2.2	6.2	–5.3	100	–4.0	75.5	–1.3	24.5	6.1	6.3	5.5	5.1	2.3
	(0.3)	(0.1)	(0.2)	(0.3)		(0.2)		(0.2)		(0.3)	(0.2)	(0.5)	(0.5)	(0.1)
Gini	0.80	0.56	0.78							0.78	0.78	0.77	0.76	0.58
	(0.003)	(0.011)	(0.003)							(0.004)	(0.004)	(0.01)	(0.01)	(0.011)
% held top 1%	32.1	13.2	30.0							30.4	30.1	28.0	26.8	13.5
	(0.5)	(1.6)	(0.6)							(0.6)	(0.6)	(1.8)	(2.0)	(1.5)
% held top 5%	56.9	29.5	55.3							55.2	55.4	54.3	52.7	30.3
	(0.6)	(1.5)	(0.6)							(0.7)	(0.7)	(1.8)	(2.0)	(1.4)
% held top 10%	69.0	41.8	67.1							67.0	67.2	65.1	64.2	43.0
	(0.5)	(1.3)	(0.6)							(0.6)	(0.6)	(1.7)	(1.9)	(1.3)

Notes:
[1]In thousands of euros except columns 5, 7, and 9.
[2]Standard errors in parentheses.

turn is smaller than the U.S. counterfactual. If it were not for the difference in household composition (columns 8 and 9 in Table 7) the difference between Spain and the (counterfactual) U.S. would be 2.75 times the actual U.S. vs. Spain difference. These differences diminish further up in the distribution, as the corresponding values for the 90th percentile show.

Types of Households that Make the Compositional Difference

In what follows we try to learn more about where these differences come from, namely which particular types of households among the 16 considered are behind these estimated compositional differences. To this end in Table 8 we vary the proportion of household types in the U.S. one type at a time. Specifically, we divide households into two types: the group of interest and the rest. Then we see how U.S. wealth at various percentiles (p25, median, and p75) would change if only the proportion of households in the U.S. of that particular type changed to the Spanish one. Thus, for each group j we obtain counterfactual medians (and p25, p75) from distributions of the form:

$$\hat{F}^{SP}_{US[j]}(r) = \widehat{\Pr}_{US}(w \leq r | z = j)\widehat{\Pr}_{SP}(z = j) + \widehat{\Pr}_{US}(w \leq r | z \neq j)\widehat{\Pr}_{SP}(z \neq j) \quad (j = 1, \ldots, J).$$

The results in the table show that it is mostly (i) couples aged 55 and over, followed by (ii) very young single women and couples (<25), (iii) single women under 55 with children, and (iv) couples aged 35 to 55 with children, which are responsible for the changes in the counterfactual U.S. distribution. For example, if we single out the group of single female households with children aged between 25 and 34 vs. the rest and change their relative weights in the U.S. population (2.4 and 97.6 percent, see Table 4) by the Spanish weights (0.3 and 99.7 percent), the U.S. median would increase by 4,100 euros. In the case of couples aged under 25, the increase in the U.S. median would be 3,800. Households in (ii) and (iii) have typically low wealth in both countries (see, for example, the median by groups in columns 3 and 4 of Table 4); the higher incidence of these household types in the U.S. as compared to Spain is responsible for a large part of the estimated increase in counterfactual U.S. wealth relative to factual U.S. wealth. In contrast we see that the low incidence of couple households over 55 in the U.S. (19.7 percent) compared to Spain (28.2 percent) and of couples with children aged 35 to 55 (16 vs. 20.9 percent) pushes down the U.S. quantiles, proportionately more at the median and above. These are typically rich households and if their share in the U.S. were to be the one prevailing in Spain, the U.S. median would go up by 10,900 and 3,800 euros, respectively, and the U.S. 75th percentile by 28,400 and 6,900 (see Table 8).

Summary Inequality Measures: Gini, and Share of Wealth Held by Top Percentiles

Commonly reported summary measures are the Gini coefficient and selected ordinates of the Lorenz curve, namely the share in total wealth of the richest x percent (where, for example, x equals 10, 5 or 1).

The Lorenz curve is given by:

TABLE 8

Difference Due to Household Composition, by Household Groups[1]: Varying One Group at a Time

	p25	Diff. with[2] U.S. p25	p50	Diff. with U.S. p50	p75	Diff. with U.S. p75
Age < 25						
Couple	11.0	1.3	69.7	3.8	227.8	6.7
Single male	10.7	1.0	67.6	1.8	223.1	1.9
Single female	11.2	1.5	68.8	3.0	225.3	4.2
25 ≤ Age < 35						
Couple						
No children	9.7	–0.03	65.7	–0.15	220.5	–0.6
Children	9.9	0.2	67.9	2.0	224.3	3.2
Single male	10.1	0.4	66.4	0.6	222.1	1.0
Single female						
No children	10.2	0.4	66.6	0.7	222.4	1.3
Children	11.3	1.6	70.0	4.1	228.0	6.8
35 ≤ Age < 55						
Couple						
No children	9.7	0	66.0	0.1	221.1	0
Children	11.2	1.5	69.6	3.8	228.0	6.9
Single male	9.7	0	66.3	0.4	222.1	1.0
Single female						
No children	10.0	0.3	66.9	1.1	223.2	2.1
Children	11.0	1.2	69.6	3.8	228.0	6.9
Age ≥ 55						
Couple	13.0	3.3	76.7	10.9	249.5	28.4
Single male	9.7	–0.02	65.8	0	221.1	0
Single female	9.8	0.05	65.7	–0.1	220.5	–0.6

Notes:
[1]In thousands of euros. Standard errors are provided in Table A1.
[2]Memo from Tables 7 and 8 (when varying all groups at the same time):
$p25_{US} = 9.7$, $p25_{US}^{SP} = 22.6$, $p25_{US}^{SP} - p25_{US} = 12.9$
$p50_{US} = 65.8$, $p50_{US}^{SP} = 91.6$, $p50_{US}^{SP} - p50_{US} = 25.8$
$p75_{US} = 221.1$, $p75_{US}^{SP} = 282.9$, $p75_{US}^{SP} - p75_{US} = 61.7$
(note that in the case of quantiles the sum of the differences for each group is not equal to the overall difference).

$$L(F(r)) = \frac{E(W|w \le r)F(r)}{\mu} \equiv \frac{H(r)}{\mu}$$

where $\mu = E(W)$.

The Gini coefficient is defined as the ratio of the areas on the Lorenz curve diagram:

$$G = 1 - 2\int_0^1 L(p)\,dp \equiv 1 - 2\frac{E[H(W)]}{\mu}.$$

The counterfactual U.S. Lorenz curve can be calculated as the empirical counterpart to:

$$L_{US}^{SP} = \frac{H_{US}^{SP}(r)}{\mu_{US}^{SP}}$$

where

$$H_{US}^{SP}(r) = E_{US}[1(W \le r)WR(z)],$$

$$\mu_{US}^{SP} = E_{US}[WR(z)],$$

and the weighting factor is $R(j) = \Pr_{SP}(z = j)/\Pr_{US}(z = j)$.

Similarly, the counterfactual U.S. Gini coefficient is given by:

$$G_{US}^{SP} = 1 - 2\frac{E_{US}\left[H_{US}^{SP}(W)R(z)\right]}{\mu_{US}^{SP}}.$$

Note that to evaluate the cumulative net wealth share for the U.S. counterfactual, the U.S. population weight factor for each household has to be corrected by the relative number of households in the group for Spain relative to the U.S., i.e. $\frac{\sum_{i=1}^{N^{SP}} 1\left(z_i^{SP} = j\right)}{\sum_{i=1}^{N^{US}} 1\left(z_i^{US} = j\right)}$.

In Figure 3 the Lorenz curves for the U.S., Spain, and counterfactual U.S. wealth distributions are plotted. As expected, the curve for the Spanish distribution is nearer the line of perfect equality than the U.S. curve. The Lorenz curve for the counterfactual U.S. distribution is distinctly nearer the perfect equality curve than the U.S., but closer to the U.S. curve than to the Spanish one. Although too

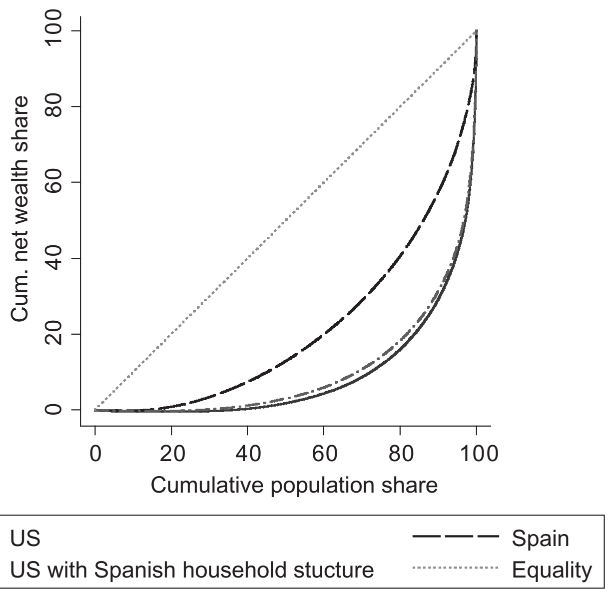

Figure 3. Lorenz Curves

small to be noticeable in the graph, some negative values for the cumulative net wealth shares are observed (the minimum being –0.15 for Spain and –0.44 for the U.S.) given the existence of negative values for net wealth.[28] Some Lorenz curves ordinates for the three distributions, namely the percentage of total wealth held by some top percentiles, can be found in Table 7. Values of the Gini coefficient are 0.80 for the U.S., 0.56 for Spain, and 0.78 for the U.S. counterfactual.[29]

We see that in contrast to what we found with position and dispersion measures along the distribution, Gini and wealth top shares do not vary as much between the U.S. and the counterfactual U.S. distributions. However, this is the net result of reduced differences at the bottom and increased differences at the top. This shows that very distinct information has been gained by comparing the entire wealth distribution as opposed to the commonly reported Gini coefficient and wealth top shares.[30]

Sampling design has a large impact on the statistical precision of Gini coefficients. In parallel with the calculations reported in Table 5, we obtained bootstrap standard errors for the U.S. Gini coefficient (0.8) using the SCF with oversampling and an equivalent random sample. The former is 0.003 and the latter is almost five times larger (0.014). The Gini coefficient for Spain is 0.56 with a bootstrap standard error with oversampling of 0.011.

Generalized Lorenz Curve

Although wealth is less unequally distributed in Spain than in the U.S. (and than in counterfactual U.S.), since the U.S. and counterfactual U.S. means exceed the Spanish mean we cannot say which distribution is to be preferred. Therefore, we briefly analyze the Generalized Lorenz curve defined as

$$H(r) = E(W|w \leq r)F(r).$$

That is, the Lorenz curve multiplied by the mean or, equivalently, the cumulative mean wealth at each point of the cumulative population share. While Lorenz types of criteria ignore the size of overall wealth, this is not the case for the Generalized Lorenz.

Figure 4 contains the Generalized Lorenz curves for the three distributions. When size is taken into account, the Generalized Lorenz curve for the counterfactual U.S. distribution lies between the Spanish and the U.S. curves for 75 percent of the population but surpasses both for the top 25 percent. Furthermore, since the Generalized Lorenz curve for the U.S. and the counterfactual U.S. distributions do not cross, there is unambiguous social welfare ordering in favour of the counterfactual U.S. as compared to the U.S. When comparing Spain to the U.S. or the

[28] Jenkins and Jäntti (2005) contains a useful discussion on how to apply the methods commonly used to summarize income distributions to the study of wealth distributions given the peculiarities of wealth (i.e. non-negligible zero and negative values etc.).

[29] In the presence of negative values the Gini coefficient is not bounded by one. Chen *et al.* (1982) propose a normalization.

[30] Other measures of inequality, such as generalized entropy indices, could be informative about where in the distribution differences occur. However, a similar ranking of distributions would be obtained from all indices of relative inequality given the Lorenz dominance relations among the three distributions.

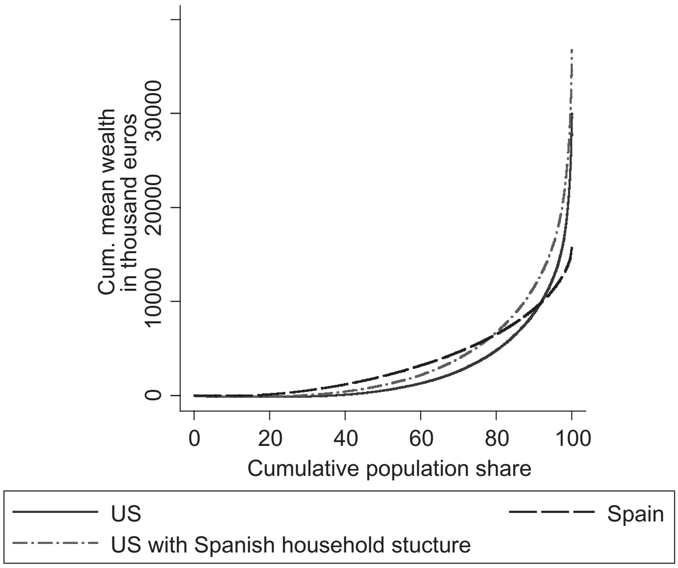

Figure 4. Generalized Lorenz Curves

counterfactual U.S. there are trade-offs between gains for the lower percentiles and losses for the wealthier, given the observed crossing of the curve with the other two. In the case of Spain vs. the U.S. there are gains for 90 percent of the population in Spain as compared to the U.S. and losses for the wealthier 10 percent.

The parallelism we observe between our Generalized Lorenz curves comparisons and *cdf* comparisons is to be expected given the equivalence of Generalized Lorenz dominance and second-order dominance of *cdf*s.

6. Extensions

Controlling for Education

We study the sensitivity of results to conditioning on education. It is natural to condition on education since it is an observable component of initial human capital and permanent income. The counterfactual U.S. distribution that uses the Spanish household structure conditioned on education is calculated as follows:

$$\hat{F}_{US}^{SP(X)}(r) = \frac{1}{\sum_{i=1}^{S^{US}} \psi_i^{US}} \sum_{i=1}^{S^{US}} \psi_i^{US} 1\left(w_i^{US} \leq r\right) \hat{Q}_i$$

where

$$\hat{Q}_i = \hat{R}_i \frac{\widehat{\Pr}_{SP}\left(X = X_i^{US} \middle| z = z_i^{US}\right) \widehat{\Pr}_{US}\left(X = X_i^{US}\right)}{\widehat{\Pr}_{US}\left(X = X_i^{US} \middle| z = z_i^{US}\right) \widehat{\Pr}_{SP}\left(X = X_i^{US}\right)}.$$

We distinguish between two educational categories (having a college degree or not) so that X is a binary indicator. The results are reported in column 10 of Table 7. As we can see, controlling for education does not alter qualitatively the conclusions on the role of household structure.

Alternative Demographic Group Classifications

We check the robustness of our results to variations in household type classification. First, we consider a 13 groups classification where there are no separate groups for households younger than 25, so that the first age category is for households aged less than 35. The motivation for considering fewer groups is to avoid groups with small sample sizes. Results presented in Table 7 (column 11) are not significantly different to the ones obtained with the 16 groups classification.

Second, we analyze an 18 groups classification where households over 54 (specifically couples and lone women) are further divided according to whether they include at least one adult-aged child. The objective is to reflect the difference across the two countries in the proportion of households where more than two adult generations live together, which mirrors the difference in the number of young households. In the U.S., 16.9 percent of households are formed by couples aged over 54 with no adult children living with them, while those with adult children amount to 2.8 percent. In Spain the percentages are reversed (0.4 percent without and 27.8 percent with). Lone mothers without (with) adult children constitute 9.2 percent (1.1 percent) of households in the U.S. but 0.06 percent (12 percent) in Spain. However, this classification produces small cells in our sample for Spain (three observations for single females without adult children). These figures are obtained defining adult children as those aged 18 or more. We also consider an alternative definition with children adulthood starting at 25. The population percentages do not change much.[31] We evaluate the counterfactual U.S. distributions for these two 18 groups household classification. The results are reported in Table 7 (columns 12 and 13). We can see that the results are qualitatively unchanged. There are some small differences, however. These counterfactual U.S. distributions are slightly closer to the Spanish distribution relative to the counterfactual based on the main 16 groups classification. Quantiles in the first part of the distribution are somewhat higher and lower in the upper part. Looking at wealth quantiles of the various groups involved (not reported) reveals that lone women with adult children in the U.S. are wealthier than those without in the first half of the distribution, while couples with adult children are less wealthy than those without in the upper part. This coupled with the larger percentage of households with adult children in Spain would explain this slight compression of the U.S. counterfactual when using the 18 group classifications.

Counterfactual Spain

Finally, we present various statistics of the counterfactual Spanish distribution. The results can be found in column 14 of Table 7. The counterfactual dis-

[31]Namely, 17.7 and 2 percent for couples in the U.S. and 3.2 and 25 percent in Spain; 9.4 and 0.9 percent for single women in the U.S. and 0.3 and 11.8 percent in Spain.

tribution for Spain evolves below the factual one. For the first part of the distribution (and up to approximately the 60th percentile) the difference between the two countries is reduced. In particular, comparing the U.S. (Table 7, column 1) to the counterfactual for Spain reduces 65 percent of the difference between the two countries at the first decile and 29 percent at the first quartile and at the median. For the upper half, the counterfactual Spanish distribution is lower than both the U.S. and the Spanish *cdf*s, therefore increasing further the factual differences in *cdf*s. For example, at the 75th percentile the difference between the U.S. and the (counterfactual) Spain would be 1.4 the actual U.S. vs. Spain difference if it were not for the difference in household composition. These differences are reduced for higher up quantiles as the differences in the 90th percentile show.

The results obtained comparing the U.S. to counterfactual Spain go qualitatively in the same direction as the ones obtained when comparing Spain to counterfactual U.S. Namely, taking into account household composition goes some way in reducing differences between the two countries in the first part of the wealth distribution, while it unveils even larger differences for the upper part. However, the extent of the differences accounted for by household composition are quite different. One reason for our focus on the U.S. counterfactual distribution is that the counterfactual experiment in the U.S. is of broader interest because the U.S. is a country of reference. Another, perhaps more important reason is that in Spain there is a low incidence of some types of households (households aged less than 25 and single females under 55), which renders the counterfactual for Spain less reliable than the U.S. counterfactual.[32]

7. Within-Group Differences

Comparing Within-Group Distributions Across Countries

Finally, we provide information about differences across countries in the wealth distribution for given household types. In Table 9 we present differences between the conditional quartiles and the conditional median of the two countries.[33] As a convenient way of obtaining standard errors for the differences, they were calculated as the coefficients in saturated quantile regressions pooling the two datasets.[34] The specification of these quantile regressions is the following:

$$Q_\tau(W|z_i) = \alpha_{1\tau}1(z=1) + \gamma_{1\tau}1(z=1)D_{SP} + \ldots + \alpha_{16\tau}1(z=16) + \gamma_{16\tau}1(z=16)D_{SP}$$

[32]There are 5.6 percent households aged less than 25 in the U.S. but only 1.6 percent in Spain, and there are 15.7 percent single-female households 55 or less but only 7 percent in Spain (see Table 4).

[33]Sample size limitations prevent us from exploring more extreme quantiles. Graphs of the conditional wealth distributions in the U.S. and Spain for each of the 16 types of households are available in Bover (2008).

[34]Since the quantile regressions are saturated, the calculations are equivalent to direct comparisons of quantiles between the two countries for each household type.

TABLE 9

Quantile Regressions for the Conditional Distributions[1]

	p25	p50	p75
Age < 25			
Couple	1.2	6.3	15.2
Single male	2.8	1.1	104.5**
Single female	4.6	6.2	52.3
25 ≤ Age < 35			
Couple			
No children	29.1**	36.5**	3.6
Children	20.9**	44.3**	56.4**
Single male	4.8	52.9**	52.0
Single female			
No children	1.3	24.3	76.8
Children	1.1	9.0	27.0
35 ≤ Age < 55			
Couple			
No children	26.8**	11.4	–61.2*
Children	18.5**	–1.4	–107.6**
Single male	11.6**	42.0**	–8.2
Single female			
No children	39.3**	83.1**	73.2**
Children	9.8	56.7**	55.2**
Age ≥ 55			
Couple	–14.5**	–98.5**	–301.9**
Single male	10.6	1.1	–32.3
Single female	22.0**	17.9**	–2.2

Notes:

[1]The coefficients reported reflect the difference of the Spanish conditional quantile with respect to the U.S. one for each of the 16 groups. In thousands of euros.

[2]*5% significance, **1% significance (taking into account stratification and clustering).

where $\tau = 0.25$, 0.50, and 0.75 and D_{SP} is a zero-one dummy for Spain. Pooled quantile regressions are estimated combining the U.S. and Spanish samples with each observation keeping its own country population weight.

In the table we report only the coefficients measuring the difference of the Spanish conditional quantiles with respect to the U.S. for each of the 16 groups (i.e. the γ's). Couples aged 25–34 with children have significantly higher wealth in Spain than in the U.S. at all quantiles considered; namely 20,900 euros at p25, 44,300 at the median, and 56,400 at p75. In contrast, couples over 54 have significantly less wealth in Spain than in the U.S. at all points of the distribution (i.e. 14,500 less at p25, 98,500 at the median, and 301,900 at p75). Interestingly, couples aged in between (i.e. aged 35 to 54) with children are better off in Spain in the first part of the distribution, worse off in the upper half, and not significantly different at the median.

Another group for which significant differences occur at all points of the conditional distribution are single females aged 35 to 54, especially those without children, who have significantly less wealth in the U.S. For other groups where differences in the conditional distributions between the two countries occur, these are more limited to certain parts of the distribution.

8. Concluding Remarks

In this paper we highlight the link between culturally inherited household structure and wealth distribution. To this end we compare two countries with very different family structures, the U.S. and Spain, using two high quality comparable datasets: the U.S. Survey of Consumer Finances 2001 and the Spanish Survey of Household Finances (EFF) 2002. We construct U.S. counterfactual wealth distributions and related summary measures as the basis for our analysis.

We find that for the first part of the distribution, controlling for household demographics explains a great deal of the observed difference between the U.S. and Spain. It accounts for 71 percent of the difference in the median between the two countries and for 55 percent of the difference in inter-quartile range. In contrast, for the upper part of the distribution, differences in family structure mask the extent of the differences between the two countries. Indeed, these differences become larger when the same household structure is assumed.

We also analyze actual and counterfactual measures of wealth inequality in the two countries. Interestingly, our results show that imposing the Spanish household structure to the U.S. wealth distribution has little effect on the Gini coefficient and wealth top shares. However, this is the net result of reduced differences at the bottom and increased differences at the top, highlighting that relevant distinct information may be missed if the entire distribution is not considered.

As an illustrative example of the importance of differences in household structure, we calculate the percentage of owner-occupied housing that would prevail in the U.S. if the demographic structure of households was similar to the one in Spain. We estimate it to be 75 percent, which lies between the 68 percent of the U.S. and the 82 percent of Spain.

We identify the main groups of households that are behind the differences between the counterfactual and the actual U.S. distributions. These are: (i) couples aged 55 and over; (ii) very young single women and couples (aged < 25); (iii) single women under 55 with children; and (iv) couples aged 35 to 54 with children. For example, if the percentage of households with a couple older than 54 in the U.S. was the one prevailing in Spain (i.e. 28.2 percent instead of 19.7 percent), the U.S. median would increase by 10,900 euros, and the 25th and 75th percentiles by 3,300 and 28,400 euros, respectively.

Looking at comparable household groups, the main feature that emerges is how differences between the U.S. and Spain in household wealth change over the life-cycle for a large group of the population, namely couples (with children when young), giving rise to an interesting reversing pattern.[35] In the U.S. they are significantly worse off at all quartiles when young (aged 25–34), significantly better off at all quartiles when old (over 54), and worse off in the first part of the distribution but better off in the upper part when aged in between (i.e. aged 35 to 54).

[35]Given the cross-section nature of our data we cannot distinguish between life-cycle and cohort effects.

Appendix: Counterfactual U.S. Density and Density Differences

We provide plots for the three estimated wealth densities. These are derived as the difference between consecutive points in the cumulative distribution and using the smoothing Stata defaults for width and kernel (i.e. Epanechnikov). Figure A1 displays the densities and Figure A2 directly the differences in densities.

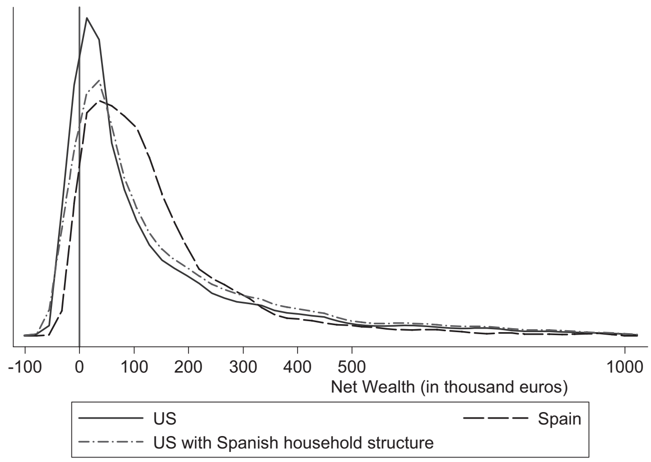

Figure A1. Estimated Densities

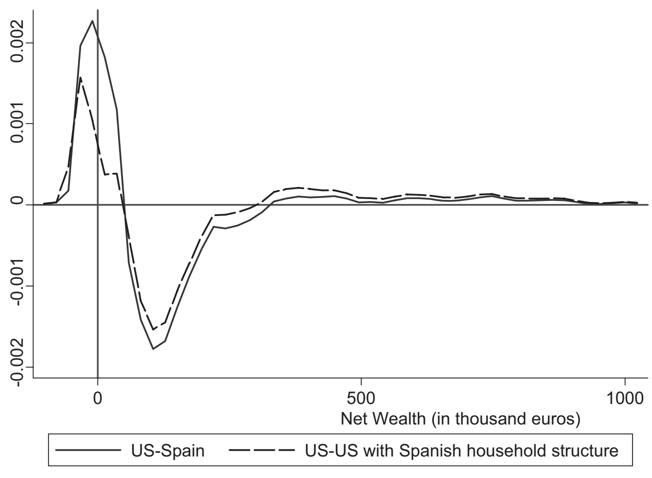

Figure A2. Difference between the Estimated Densities

TABLE A1

Standard Errors for Estimates in Table 8 (Difference Due to Household Composition, by Household Groups: Varying One Group at a Time)

	p25	Diff. with U.S. p25	p50	Diff. with U.S. p50	p75	Diff. with U.S. p75
Age < 25						
Couple	0.6	0.3	2.3	0.8	4.3	2.0
Single male	0.6	0.3	2.2	0.6	4.8	1.6
Single female	0.6	0.3	2.2	0.7	4.4	1.9
25 ≤ Age < 35						
Couple						
No children	0.5	0.1	2.1	0.4	5.3	1.5
Children	0.6	0.1	2.3	0.8	4.8	2.1
Single male	0.6	0.2	2.3	0.6	4.8	1.5
Single female						
No children	0.6	0.2	2.3	0.7	4.9	1.5
Children	0.5	0.3	2.2	0.8	4.2	2.0
35 ≤ Age < 55						
Couple						
No children	0.5	0.1	2.1	0.7	5.2	1.1
Children	0.6	0.3	2.5	1.0	4.4	2.4
Single male	0.5	0.0	2.1	0.4	5.0	1.3
Single female						
No children	0.6	0.2	2.2	0.6	4.7	1.9
Children	0.6	0.3	2.3	0.8	4.1	2.1
Age ≥ 55						
Couple	0.7	0.5	2.2	1.5	6.6	5.2
Single male	0.5	0.1	2.1	0.2	5.0	0.5
Single female	0.5	0.1	2.1	0.3	5.1	1.3

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