Unemployment and Consumption Near and Far Away from the Mediterranean*

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Abstract

We study the insurance mechanisms employed by households to absorb unemployment shocks using comparable data for four countries: Italy, Spain, Great Britain, and the US. Results on family transfers when the male household head becomes unemployed suggest that family networks are the weakest in Britain, while unemployment benefits there are instead the most generous across the four countries. Despite these differences, food consumption losses induced by unemployment of the male household head are similar across countries. These findings are consistent with the view that family support and the Welfare State substitute each other in mitigating the consequences of unemployment shocks.

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1 Introduction

Households can resort to many devices to mitigate the welfare effects of negative labor income shocks. Running down wealth, particularly liquid assets, is a natural way. Alternatively, they can use a variety of insurance mechanisms, such as credit markets and public programs designed to offset income shocks. Indeed, Dynarski and Gruber (1997) found that US households are fairly well able to smooth consumption in the face of household heads' earnings shocks, with government programs and self-insurance playing roughly equal roles. Households also receive help from other households, usually belonging to the same family.

The role of the family as an insurance device has been recognized for a long time. In the economic literature, Kotlikoff and Spivak (1981) characterized the family as an incomplete market for annuities. The family has advantages over market mechanisms in that it has lower monitoring costs and it can relieve the standard problems plaguing market insurance: adverse selection –if participation in family insurance networks is nearly mandatory and outsiders are excluded– and moral hazard –through good information about family members. On the other hand, the pool of participants sharing risk is typically small and enforcement cannot be chanelled through formal institutions, although altruism and moral suasion can help overcome this limitation.

There is a recent strand of literature analyzing insurance devices in less developed countries, in particular transfers between households within villages (Townsend 1994, for India; Udry 1994, for Nigeria; Albarran and Attanasio 2003, for Mexico). There has been less work on developed countries, except for a different literature that has studied family transfers with a view to testing competing economic models of the family: altruism, gift exchange, and "family constitutions". There is a long literature for the US (Cox 1987, Laitner 1997) and more recently for France and Italy (Cigno *et al.* 2004).

Family networks are typically seen as being strong in Southern Europe, especially when compared to Northern Europe, as has been often noted in the sociological literature, not only historically but also nowadays. Reher (1998) distinguishes between Western countries where family ties are weak –Scandinavia, the British Isles, the Low Countries, Germany, Austria, and the United States– from those where they are strong, namely the Mediterranean. However, hard evidence is difficult to obtain from standard statistical sources. In this paper we examine how households cope with important labor income shocks, namely those arising from unemployment, focusing in particular on the role of the family. We follow a cross-country approach, by using comparable data sets and the same empirical methods to compare two Mediterranean countries, Italy and Spain, with a Northern European country, Great Britain, and another Anglo-Saxon country, the United States.

We start by revising the stylized facts on each country, finding that unemployment benefits are significantly more generous in Britain than in the other three countries. This motivates our examination with microeconomic data of the operation of family networks through the response of transfers from other households to the event of the household head becoming unemployed. We find, in Section 2, that financial transfers within extended families are more frequent in Italy, Spain, and the US than in Britain, that they kick in more often in the former countries when unemployment hits a household, and that they are less likely if unemployment benefits received by the household increase in comparison with the preceding year. This is consistent with the view that whenever the Welfare State fails to mitigate the consequences of unemployment, the role of family support is stronger.

This finding motivates the second part of the paper, where we analyze the impact of unemployment shocks on household welfare, captured by food consumption losses. In Section 3 we estimate the consumption losses induced by head of household unemployment, finding no significant differences across the four countries studied. We check in Section 4 that this finding is not a consequence of the fact that food is a necessity and its consumption cannot be reduced. We also check that the similarity of consumption losses stands when we take into account the relative importance of the head of household's income and changes in the female head's and children's labor supply. This suggests that the recourse to the family allows households to reach similar levels of insurance as those attainable through other channels. Section 5 contains our conclusions.

2 Family matters

Households can resort to a variety of insurance mechanisms in order to mitigate the welfare effects of negative income shocks. However, private insurance markets usually do not cover the risk of unemployment (due, among other reasons, to moral hazard problems) and so public programs, particularly unemployment benefits, typically represent the most important insurance channel against that risk. For instance, Dynarski and Gruber (1997) find that government programs are the largest source of offsetting income against head of household earnings shocks arising from unemployment in the US. (Other mechanisms include personal bankruptcy laws, see Fay, Hurst, and White 2002, and insurance within the firm, see Guiso, Pistaferri, and Schivardi 2005). Thus we start this section by providing aggregate information on the availability of benefits across our four countries of interest: Italy, Spain, Great Britain, and the United States.

In Table 1 we compare benefits systems across countries for 1990-95. Unemployment Insurance (UI) supports involuntarily unemployed jobseekers who contributed when employed. The UI replacement rate (benefits divided by the previous wage) in the first 6 months is lowest in Italy and highest in Spain, but the Spanish ranking is misleading because benefits fall with unemployment length. Considering jointly UI and Unemployment Assistance (UA) –which supports the unemployed with no UI contributions or who have exhausted UI–, Italy and Spain are seen to provide no benefits at all after two years in unemployment, while the UK appears more generous than the US (cols. 1-2). People who are ineligible for both UI and UA can often rely on Social Assistance (SA), which provides a minimum income. As duration increases, the net replacement rate encompassing UI, UA, SA, housing benefits, and family benefits becomes higher in the UK than in Italy or the US, with Spain giving little protection (cols. 3-4).

(Insert Table 1 here)

Maximum benefit duration presents a similar picture: the number of months over which a worker can get the equivalent of the maximum replacement rate is much higher in the UK than in the remaining countries. Unemployment income depends on benefit duration also through actual unemployment duration. The table shows why differences in benefit regimes for long spells matter: the proportion of ongoing spells lasting for more than one year is higher, and the share of the unemployed receiving benefits much lower, in Italy and Spain than in the UK (cols. 5-7).

While the rankings vary depending on the measure observed, it is clear that unemployment benefits are more generous in the UK than elsewhere. If insurance mechanisms are substitutes, we would expect family networks to play a larger role in the other three countries. To assess this conjecture, we turn to household data.

In order to assess the role of family networks, we propose to quantify the response of family transfers to unemployment shocks. For this purpose we employ a set of longitudinal household surveys from the four countries: the Bank of Italy Survey of Household Income and Wealth (SHIW), the Spanish Continuous Family Expenditure Survey (ECPF), the British Household Panel Survey (BHPS), and the US Panel Study of Income Dynamics (PSID). See the Appendix for details on the samples extracted from these data sets.

Several theories of the family are capable of explaining the reasons for family transfers: altruism (Becker 1974), payment for services rendered (Cox 1987), and "family constitutions", i.e. a set of unwritten rules constraining the actions of family members (Cigno 1993). While our favorite interpretation is in terms of mutual insurance between households belonging to the same family, all of these theories predict a higher probability of transfers to households suffering negative economic shocks. We do not need to distinguish between them, however, since they are not inconsistent with insurance and, in particular, they provide a mechanism for the enforceability of insurance through family networks. For instance, Cigno *et al.* (1998)

cannot reject the strategic self-interest model, which explains transfers as part of self-enforcing intergenerational credit agreements, using the SHIW dataset.

We restrict the sample to households in which the male head was fully employed during year $t-1$. Within this sample, our measure of shocks is the number of months of unemployment of the male head of household in year t, which we label ΔU_{it}. As indicated by Dynarski and Gruber (1997), these are arguably individuals for whom the transition from employment to unemployment is more likely to be unanticipated, and the number of months provides a measure of shock size.

Income data reported in household surveys are well known to contain very sizable measurement error, significantly larger than that affecting consumption, which renders the results using income variables less reliable (see Deaton 1992, 138-139). For this reason, we think it more advisable to examine receipt of transfers or benefits rather than actual reported amounts. Receipt of benefits in our data appears to correspond with the country-wide information in Table 1: the share of households with an unemployed head which received income in the form of unemployment benefits is higher in Great Britain, 79%, than in Italy, 27%, with the US and Spain in between, 57% and 66% respectively (due to data limitations, in the US this variable reflects all public transfers, except pensions, rather than only unemployment benefits).

Regarding the role of family ties, the fraction of these households declaring to have received financial help from relatives in other households is equal to 9% in Italy, 5% in Spain, 3% in the US, and 1% in Great Britain, suggesting stronger family ties in the two Mediterranean countries (see the Appendix for details). The US figure may seem low vis-à-vis the 14% of households receiving transfers reported by Cox (1987) for 1981 using the President's Commission on Pension Policy survey. In our earliest wave, 1981, and not restricting the sample to households with a male head fully employed at $t-1$, our PSID figure is 5%. Cox's figure should be higher than ours because it includes any private transfer, not only those from relatives as ours, and also because it includes gifts in kind rather than money transfers alone as ours.[1] Additionally, in our samples, the fraction of households living in an inherited home is 11% in Italy and 6% in Great Britain, which is compatible with the view that intergenerational transfers are more frequent within Italian than within British extended families.

More formally, we estimate a probit model for the probability that a household receives a money transfer from other households. We include standard controls for changes in family composition –the numbers of children below 14 years old and aged 14-19 years old, the number of other members, and the fraction of females in the household–, the male household head's age and years of schooling, and a dummy for lagged home ownership (see the Appendix for definitions and Tables A.2 to A.5 for descriptive statistics). Yearly dummies capture aggregate shocks.

Table 2 reports, in percentage form, the marginal probability effect of ΔU_{it}, estimated separately for each country. Again there is a significant difference between Italy, Spain, and the US, on the one hand, and Britain (where the effect is not significant), on the other. The ratio of the marginal probability effect to the average fraction of households who receive a transfer is equal to 15.7% in the US, 15.3% in Italy, 8.1% in Spain and roughly nil in Britain. This result suggests not only that in the first three countries financial transfers within extended families are more frequent, but also that they kick in more often when unemployment hits a household.

(Insert Table 2 here)

The table also shows the marginal probability effect on transfers of an increase in real unemployment benefits received by the household from the preceding year. We examine this variable because the household may have been receiving benefits at $t-1$ and measure it through a 0-1 dummy variable to minimize the impact of measurement error. Our estimates indicate that family transfers are less likely if benefits increase, though the effect is only significant in Italy and the US.

The foregoing evidence is consistent with the view that whenever the Welfare State fails to mitigate the consequences of unemployment, the role of family support is stronger. We can also quote some informal evidence suggesting that family ties are stronger in Mediterranean countries than in Great Britain, with the US often showing up in between. For instance, in the latter countries, members of the same family are often scattered all over the country while in the former countries they are more likely to live in the same area. Recorded regional migration rates support this observation: in the 1980s and 1990s, the average fraction of the population changing region in one year was around 0.5% in Italy and Spain vs. 1%-1.5% in Britain and 2.8% in the US (OECD 1990, Maclennan *et al.* 1998). More direct evidence on physical distance between family members is offered by our surveys' samples: the fraction of households in which relatives other than parents and children are present is equal to 18% in Spain, 8% in Italy, 6% in the US, and 4% in Britain.

Additionally, in the Mediterranean countries children tend to remain close to their parents even when they have formed new households. In Italy, 45% of all married Italians aged up to 65 live within a single kilometer of at least one parent after marrying (ISTAT 1999, 102). While lacking similar information, we suspect that the corresponding numbers would be much lower in the Anglo-Saxon countries. Moreover, in Italy and Spain children wait longer before leaving the parental home. In 1995 the fraction of youngsters between 25 and 29 years of age living with their parents was 59% in Spain and 56% in Italy, while in Great Britain it was 17% (Eurostat 1997, see also Becker *et al.* 2005). This feature is also corroborated in our samples: the average age of the children living in the household is 18 years old in Italy, 15 y.o. in Spain, 11 y.o. in the US, and 9 y.o. in Britain.

Note that the nexus of causality between the roles of the Welfare State and the family is not obvious. One could argue that the greater generosity of the welfare system in Great Britain is a response to the weakness of family networks or, alternatively, that the latter retreated when the Welfare State was strengthened.

An important issue is whether Welfare State insurance increases the total level of insurance available to households or it merely crowds out the insurance provided by family networks. Attanasio and Rios-Rull (2000), Albarran and Attanasio (2003), Di Tella and MacCulloch (2002), and Krueger and Perri (2005) present models where such crowding out happens, sometimes fully or even more than fully. As to empirical evidence, the literature indicates that private transfers fall when public transfers increase (see Cox 1987). More recently, Albarran and Attanasio (2002) find that households who received money from the PROGRESA public program in Mexico are less likely to receive private transfers; and Schoeni (2002) estimates that in the US unemployment benefits displace family support by as much as 24-40 cents per dollar.

Since public programs are less generous in Italy, Spain, and the US than in the UK, it follows that consumption smoothing of income shocks should also be harder there. However, as we have seen, in the former countries households experiencing unemployment shocks are more likely to receive transfers from their family. Thus, it is interesting to try to measure the overall degree of insurance that households can obtain from tapping all sources. This is carried out in the next section.

3 Unemployment and consumption

In order to approximate the degree of insurance available to households, we need a measure of changes in household welfare due to unemployment shocks. Following the literature, we focus on consumption. After briefly discussing the underlying theoretical framework that can help us understand the results, we present our estimates for the consumption impact of unemployment shocks.

3.1 Framework of analysis

We follow the approach to testing the risk sharing model as spelled out in Cochrane (1991) and recently pursued by Blundell *et al.* (2005). The first step is standard. Start by assuming that the household maximizes an isoelastic utility function in consumption, C_{it}, which is additively separable with respect to leisure (with labor being supplied inelastically):

$$\underset{\{C\}}{Max} \;\; E_t \sum_{j=0}^{T} \beta^{-1}(C_{i,t+j}^{\beta} - 1)e^{Z'_{i,t+j}\theta} \tag{1}$$

where $Z'_{i,t+j}\theta$ includes taste shifters and discount rate heterogeneity, i denotes households, t denotes years, T is the household's final period, known with certainty, and E is the expectations operator. Assuming that the only source of uncertainty is labor income (including transfers) and that $\ln C_t$ is normally distributed, we get the well-known appproximation to the Euler equation (see Deaton 1992):

$$\Delta \ln C_{it} \simeq \Delta Z'_{it}\theta + \Omega_{it} + \eta_{it} \tag{2}$$

where Ω_{it} captures the slope in consumption arising from interest rates, impatience, or precautionary saving, Δ is the first difference operator, and $E_{t-1}\eta_{it} = 0$. Note that if consumption and leisure were non-separable in utility (Browning and Meghir 1991), under isoelastic preferences and a unit elasticity of substitution between consumption and leisure, the Euler equation would be augmented with expected changes in leisure (see Attanasio and Weber 1993, 1995, for evidence on the UK and the US, respectively).

Blundell *et al.* (2005) further assume that labor income net of demographic effects is the sum of a permanent component following a martingale with a serially uncorrelated shock, ζ_{it}, and a transitory component following a moving average process with random term ε_{it}. Under these conditions, making a logarithmic approximation to the household's intertemporal budget constraint and using a Taylor expansion, they express the change in consumption as:

$$\Delta \ln C_{it} \simeq \Delta Z'_{it}\theta + \Omega_{it} + \phi_t\zeta_{it} + \psi_t\varepsilon_{it} \tag{3}$$

Eq. (3) encompasses the cases of perfect credit markets ($\phi_t = \psi_t = 0$), no insurance ($\phi_t = \psi_t = 1$), and the permanent income hypothesis (quadratic utility) with self-insurance through saving, for which those authors show that $\phi_t = \psi_t/\delta_t = \pi_{it}$, where δ_t is an age-increasing known weight and π_{it} the share of future labor income in current human and financial wealth (Blundell *et al.* 2005, Appendix A.1). In the intermediate cases, $0 < \phi_t < 1$ and/or $0 < \psi_t < 1$, deviations of consumption growth from its predictable component result from income shocks times a coefficient of partial insurance. Our empirical strategy is inspired in this insight. It is worth recalling that the perfect insurance case has been repeatedly rejected in the data, see Cochrane (1991), Udry (1994), Attanasio and Davis (1996), Hayashi *et al.* (1996), or Stephens (2001); while Altug and Miller (1990), Mace (1991), and Townsend (1994) are exceptions.

We estimate the relationship between the growth rate of household consumption and changes in the number of months of unemployment of the male head of household, U_{it}, as captured by parameter γ in the following regression:

$$\Delta \ln C_{it} = \alpha + \Delta Z_{it}'\theta + \lambda_t + \gamma \Delta U_{it} + \xi_{it} \tag{4}$$

where Z_{it} denotes demographic variables, λ_t denotes a linear combination of time dummies, and ξ_{it} is a random term. Note that Blundell *et al.* (2005) also obtain a random component in eq. (3) because they further decompose Ω_{it} into a cohort-specific slope term and a random individual deviation from that slope.

As in Section 2, we restrict the sample to households whose male head is continuously employed during $t-1$, for whom unemployment at t is more likely to be unanticipated. Moreover, focusing on male household heads mitigates potential endogeneity problems due to the joint determination of consumption and labor supply decisions. In our specification, we focus on an important transitory and idiosyncratic shock to labor income, namely going from no unemployment at $t-1$ to some unemployment at t, while permanent shocks are captured by ξ_{it}, and the time dummies capture aggregate shocks. We therefore interpret γ as a measure of the degree of partial insurance to these shocks, attained through formal or informal mechanisms. We should note that γ will also capture saved costs of going to work, which we have no way of identifying. This will not matter if they are similar across countries.

Our approach echoes Castillo *et al.* (2000), which compares the difference between the consumption *levels* of employed and unemployed workers in Portugal and Spain. Related work for the US, based like ours on longitudinal data, are found in Dynarski and Sheffrin (1987), Dynarski and Gruber (1997), and Stephens (2001, 2004). However, as far as we know, we are the first to compute the consumption growth effects of unemployment shocks with comparable household data for several countries and a common framework.

Browning and Crossley (2001, 2004) study the effect of unemployment shocks on income through unemployment benefits in Canada. Benefits are measured accurately, from administrative records, and they are projected on a rich set of variables (e.g. previous employment history) to approximate so-called potential benefits. Survey data availability and quality preclude this option in our case.

3.2 The painfulness of unemployment shocks

Let us now turn to the empirical evidence. We first describe our empirical specification more in detail and present estimates of consumption growth equations with unemployment shocks. Subsequently we refine our measure and reexamine the results. Lastly, we perform a few robustness checks on the empirical specification.

As just indicated, we estimate household food consumption losses due to the male head becoming unemployed, measured by the estimates –obtained separately for each of country– of parameter γ in eq. (4), slightly rewritten as:

$$\Delta \ln C_{it} = \alpha + \Delta X_{it}'\omega + W_{it}'\phi + \lambda_t + \gamma \Delta U_{it} + \xi_{it} \tag{5}$$

Throughout the paper the term "male (female) household head" refers simply to the male (female) member of the heading couple (or single head). The demographic variables are meant to control for taste shifters and discount rate heterogeneity. One set of controls (X_{it}) affects the level of consumption, which includes the numbers of children below 14 years old and aged 14-19 years old, the number of other members, and the fraction of females in the household. Further controls (W_{it}) affect the consumption profile, including the years of schooling and the age of the male household head, and, to capture household wealth, a dummy for lagged home ownership (Hurst and Stafford 2004, show that US households that experience an unemployment shock and have limited access to liquid assets are more likely to refinance their home mortgage). For our specification to be valid these variables should be orthogonal to preference shocks and to consumption measurement error.

We analyze food consumption because it is the only measure available in our British and US data sets. This is useful, in that food expenditures closely track food consumption flows, while this is not the case for durable goods (we are implicitly imposing separability in utility of food and other goods' consumption). Moreover, food is a necessity and so it is a key component of welfare. On the other hand, food should be preferentially smoothed in comparison with other items, and thus its changes may understate changes in household well-being. Nevertheless, Browning and Crossley (2004) show that with irreversibility (i.e. no second-hand markets for durables) and under liquidity constraints, agents absorb most of temporary, moderate earnings cuts by strongly reducing durables purchases, but they instead cut non-durables significantly when facing larger income shocks. As to empirical evidence, for the US Dynarski and Gruber (1997) find that food consumption reductions due to earnings losses arising from unemployment of the household head are about half the size of the reductions in expenditures on other goods. For Canada, Browning and Crossley (2004) compare the response of food and clothing, finding the same 50% lower response of food on average, but a 30% larger response of food than of clothing in households where the lost job provided more than 60% of household income. In any event, our results would be strengthened if the differential degree of smoothing of food vis-à-vis other goods was similar across countries. In a previous version we found that the response of *total* consumption to unemployment shocks is similar in Italy, Spain, and West Germany. We omit this evidence here, since in Germany consumption is measured as income minus saving, and therefore the results are not comparable to those for the other countries.

The Appendix provides information on the data sets and descriptive statistics on the variables included in the analysis. Remarkably, the fraction of households with an unemployed male head, going from 2% to 3%, is quite low in all countries (Tables A.2 to A.5). This may seem surprising in the case of Spain. However, over the sample period for this country, the unemployment rate of male household heads was 10.5 percentage points lower than the national unemployment rate.

As indicated by Dynarski and Gruber (1997), if ΔU_{it} is measured with error then the estimate of γ will be biased downward. Thus, in principle, differences in measurement error across countries could explain differences in the estimates of γ. We deal with this potential problem using the instrumental variable (IV) method. In all our data sets two different types of questions provide information on the employment status of male household heads. One type allows us to construct the number of months of unemployment during year t. The other gives information on the employment status at a precise date (see the Appendix). From the latter we construct a dummy variable M_{it} equal to 1 in case of unemployment and 0 otherwise. Its first difference, ΔM_{it}, is evidently correlated with our variable of interest, ΔU_{it} (which is confirmed by F-tests for the inclusion of the IV in the first stage, see Staiger and Stock 1997, reported in the tables). However, under the fairly reasonable assumption that the measurement error in ΔU_{it} is not correlated with the measurement error in ΔM_{it}, the latter is a good instrument for the former in order to reduce bias in the estimate of γ.

Table 3 reports the IV estimates of γ in eq. (5) in percentage form, as well as OLS estimates. Using the test proposed by Arellano (1993) to check whether the OLS and IV estimates differ significantly, we cannot reject the null hypothesis of no difference. This result suggests that measurerement error does not cause a significant bias in our estimates. The estimates are similar across countries, though in the Mediterranean countries losses appear to be smaller and less statistically significant than in the Anglo-Saxon ones, ranging from a 1.3% decrease in yearly household food consumption for a one-month increase in unemployment of the male household head in Britain to 0.2% in Spain. How significant are these differences? For no pairwise comparison can we reject statistically the equality of γ across countries at conventional significance levels. Thus, food consumption appears to react in similar ways in all countries. It is worth noting that these conclusions do not change if households with self-employed males heads are excluded from the analysis.

(Insert Table 3 here)

4 Robustness

We have so far established that food consumption losses induced by the unemployment of the male household head are similar across countries, and we would like to argue that this similarity results from the fact that family support and the welfare state substitute each other in mitigating the consequences of unemployment shocks. There are, however, other competing explanations and in this section we explore the extent to which our interpretation is robust with respect to these alternatives.

4.1 Is food consumption compressible?

It could be argued that food consumption is a necessity and as such cannot be compressed. If this were true, the similarity of food consumption changes induced by unemployment across countries would just be the consequence of the fact that within each country families have little room to change their consumption habits. Under this interpretation, families hit by unemployment would not be able to reduce food consumption but only consumption of other goods. Our evidence, however, suggests that in the countries we consider, the margins for compression of food consumption are wide and do not constrain the changes induced by unemployment shocks.

The first row of Table 4 reports the 10th percentile of the distribution of annual real food consumption changes between $t-1$ and t in each country, after controlling for changes in demographic characteristics (the X variables in eq. (5)). Looking, for example, at Italy in the first column, households at the 10th percentile of the distribution decrease their food consumption by 51.0%. The households in which the male head becomes unemployed at t, after having been fully employed at $t-1$ (second row), decrease their food consumption by 8.8% and such a decrease corresponds to the 47th percentile of the distribution of consumption changes. These statistics indicate that in Italy there is a wide margin for a compression of food consumption beyond what is induced by the unemployment shock we analyze. The same conclusion can be reached also for the other countries, although there are cross-country differences in the variability of the distributions. We conclude that the similarity across countries of the food consumption changes induced by unemployment is not a consequence of the fact that food is a necessity and its consumption cannot be compressed.

(Insert Table 4 here)

4.2 The importance of the male head's income

The number of months of unemployment during the current year only captures in part the size of the unemployment shock. We therefore wish to check the robustness of the finding in Table 3 by refining the measure of the shock.

Table 5 reveals that the average number of months of unemployment of male household heads in our sample is actually higher in the Mediterranean than in the Anglo-Saxon countries, though Spain and Britain are close. These figures suggest that unemployment entails a larger shock near the Mediterranean, especially in Italy. But it could be that the male head's labor income before becoming unemployed represents a lower fraction of total household income there. As indicated before, income data reported in household surveys are noisy, but it is worth checking, since the number of household members is higher in Italy and Spain (see Tables A.2 to A.5). Table 5 shows that there are no big differences across the three European countries, but indeed in Italy the male head's income importance is somewhat lower than in Britain and the US.

(Insert Table 5 here)

To probe this issue, we follow Dynarski and Gruber (1997) and estimate the household income effects of changes in the male head's labor income as follows:

$$\Delta Y_{it} = \vartheta + \Delta X_{it}' \varrho + W_{it}' \kappa + \lambda_t + \mu \Delta Y L_{it}^h + \varpi (\Delta Y L_{it}^h \Delta U_{it}) + \rho \Delta U_{it} + \nu_{it} \qquad (6)$$

where Y_{it} denotes total household income and YL_{it}^h the male head's labor income, which excludes any benefits. We also include ΔYL_{it}^h interacted with ΔU_{it}, and the controls in eq. (5).

The first line of Table 6 captures the main effect: the loss of labor income – excluding benefits– due to unemployment. It confirms the lower impact of changes in the male head's labor income on total household labor income in Italy and Spain. The third line suggests that when the male head becomes unemployed, other mechanisms alleviate the loss of income (though significantly only in Spain and the US). We can observe two channels for this mechanism. First, transfers from relatives – found in Section 2 for all countries bar Britain. Second, real unemployment benefits received by the family respond positively and significantly to ΔU_{it} in all countries except for Italy (this is found in an equation like (6) with unemployment benefits in the left hand side). Lastly, the coefficients on the interacted variables are similar across countries.

(Insert Table 6 here)

Thus, like Browning and Crossley (2001), we correct our measure of unemployment shocks by interacting the male head's months of unemployment with his lagged income importance, which yields an *adjusted unemployment shock*, ΔU_{it}^{*}. The average values across countries are shown in the third line of Table 5.

Table 7 presents evidence on the food consumption effects of unemployment based on the estimation of eq. (5) but replacing ΔU_{it} by ΔU_{it}^{*}. Focusing again on the IV estimates, the findings in Table 3 are confirmed: once the male head's income importance is taken into account, there is still a similar impact of unemployment shocks across all countries, though somewhat smaller in Italy and Spain. Again, for no pairwise comparison can we reject statistically the equality of γ across countries. (We again check whether the OLS and IV estimates differ significantly and we cannot reject the null hypothesis of no difference.)

(Insert Table 7 here)

To get a sense of magnitudes, we have computed the impact of ΔU_{it}^{*} for a male head with 6 months of unemployment and with a 50% importance of lagged income.

The estimated consumption loss is around 1% in Spain, 5% in Italy, 7% in the US, and 9% in Britain. We believe that the effect for Spain is underestimated, because the Spanish data are collected quarterly, rather than annually as in the other countries. Thus the instrumental variable, which refers to a given point in time, should be less able to correct measurement bias than in the other countries. In this case the OLS estimate is higher, 2%, but this should still be downward biased if there is measurement error. This finding echoes the results for Spanish households by Albarran (2000), who finds little consumption effects of the variance of the household-specific component of income risk but a significant effect of the cohort component.

It is not straightforward to compare our results with those in the literature, since other researchers use different regressors and units. For the US, using the PSID and the Consumer Expenditure Survey, Blundell *et al.* (2005) find that changes in transitory labor income lead to a 5-6% reduction in non-durable consumption. More closely, Dynarski and Gruber (1997) estimate the response of consumption to earnings variation using a 0-1 dummy variable for unemployment as an instrument. They find that changes in the head's earnings reduces food consumption by 6-8%. Possibly the most comparable results are obtained with the PSID by Stephens (2001), who reports a 9% drop in annual food consumption for US households whose head is displaced from his job.

4.3 Added worker effects: spouses and children

When the male head becomes unemployed, other members of the household may react so as to smooth income. In particular, the female head can increase her labor supply. This is often called the *added worker effect.* This possibility represents a problem for our interpretation of the evidence because, given the low female employment rate in Italy and Spain when compared with the other two countries, the smaller consumption losses observed in these countries could result from female heads starting to work. In the Anglo-Saxon countries this option is more restricted because most females work already, so that they can only raise their hours of work. Indeed, the employment to working-age population ratio for females aged 25-54 years old in the 1980s was equal to 40% in Italy, 31% in Spain, 64% in the UK and 65% in the US (OECD 1992).

In the same spirit as for total household income, we analyze the correlation between the two spouses' labor incomes by running the following regression:

$$\Delta YL_{it}^{w} = \vartheta' + \Delta X_{it}'\varrho' + W_{it}'\kappa' + \lambda_t' + \mu'\Delta YL_{it}^{h} + \varpi'(\Delta YL_{it}^{h}\Delta U_{it}) + \rho'\Delta U_{it} + \nu_{it}' \quad (7)$$

where YL_{it}^{w} denotes the female head's labor income and the rest of the specification is as in eq. (6).

Table 8 shows that in Italy changes in the male and female head's labor incomes are positively correlated, the event of the male head becoming unemployed brings forth higher female labor income (insignificantly), and the interaction is negative. In the other countries none of these effects is significant, which is consistent with a large literature that has found no strong effects of the husband's unemployment on the labor supply of the wife (Cullen and Gruber 2000; an exception is Stephens 2002).

While some sign of a female head reaction is observed at least in Italy, in no country do we find evidence of a reaction of children, whose labour supply does not seem to change when the male head of the household becomes unemployed. Evidence on this is omitted to save space but is available from the authors. Becker *et al.* (2005) find evidence that children tend to abandon the parental home when their father suffers an increase in job insecurity, more specifically when he expects to become unemployed. This pattern is taken care of in this paper by including changes in the size of the household as controls in our statistical models.

(Insert Table 8 here)

Even if there is no evidence of a relevant added worker effect, with the possible exception of Italy, we have reestimated equation (5) including as controls for female labour supply three indicator variables for transitions from working to not working, not working to working, and working in both periods (leaving female heads not working in either period as the reference). The labour supply of children is instead controlled for by the inclusion of the change in the number of working children in the household. The results, presented in Table 9, are essentially the same as in Table 7. The null hypothesis that the consumption effect of unemployment is the same can be rejected in only one case, namely for Spain *vs.* Britain (with a p-value of 0.02), and in this case it is rejected in the sense of suggesting that the painfulness of unemployment shocks is larger in the northern country. If anything the sign of this difference supports the existence of a stronger insurance role of families in Spain, where the welfare state is less developed.

(Insert Table 9 here)

4.4 Other factors

Even if private markets do not usually offer contracts against the risk of unemployment, the structure of financial markets does matter, because the possibility of smoothing consumption in the face of unemployment shocks will in general depend on the presence of credit constraints. Unfortunately our data are not rich enough to assess the extent to which subjects in our samples have access to credit. In the aggregate, over the 1980s and 1990s, households made more use of consumer and mortgage credit in the US and the UK than in Italy and Spain (Jappelli and Pagano 1993, Maclennan *et al.* 1998). This fact suggests that buffering income shocks through financial markets was harder in the latter countries. And it helps explain the higher household saving ratios in Italy, 28%, and Spain, 13%, vis-à-vis the UK, 9%, and the US, 8%, over our full sample period, 1981-1996 (source: *OECD Economic Outlook* database).

We might therefore have expected that, for this reason, unemployment shocks would have a higher consumption impact in the Mediterranean countries. That the response is similar across countries may be the result of the higher reliance on family networks, as found in Section 2. But this cannot be concluded in the absence of household data on credit availability. In our consumption equation we have approximated collateral by the home ownership dummy. For unemployed workers without collateral, it is usually very difficult to access credit in all countries, and so it is not obvious that this particular group of people will have a harder time smoothing shocks in countries with less developed financial markets.

Lastly, our estimates may be subject to measurement error in unemployment, arising from the underground economy, which is larger in the Mediterranean than in the Anglo-Saxon countries. Johnson *et al.* (1998) estimate it to represent 20% of GDP in Italy, 16% in Spain, 7% in the UK, and 14% in the US (1990-93). The limited information we have does not suggest this to be a large enough source of error, however. A survey carried out in Spain in 1985 (see Muro *et al.* 1988) found that only 4% of employed heads of household were both working underground and officially counted as unemployed. The reason is that two thirds of the heads of household with an underground job also had another, registered job, while 15% were either disabled workers or retirees. We also believe that the underground economy is likely to be a complementary and not alternative device to the extended family: the family is likely to be a key channel of access to the underground economy.

5 Conclusions

In this paper we have presented evidence on the operation of the insurance mechanisms employed by households in order to mitigate the effect of unemployment shocks using comparable household survey data and the same empirical techniques for a set of four countries, namely Italy, Spain, Great Britain, and the US.

We started by quantifying the importance of family networks through the response of family transfers to the event of the household head becoming unemployed. We found that it is stronger in Italy, Spain, and the US than in Britain. This evidence becomes particularly interesting when we take into account that unemployment benefits are more generous in Britain than in the other three countries.

This highlights the interest of finding an overall measure of insurance availability for each country. Following this motivation, in the second part of the paper we have estimated the food consumption losses induced by unemployment of the male head of household with our household surveys, finding no significant differences across the four countries studied.

While not providing airtight proof, these findings are consistent with the view that whenever markets and the Welfare State fail to mitigate the consequences of unemployment, the role of family support is stronger. Thus, our empirical results on interhousehold transfers and the consumption effects of unemployment shocks suggest that family networks represent an important device that allows households to insure against labor market risk.

Appendix: Database description

Our evidence is based on four longitudinal household surveys: the Bank of Italy Survey of Household Income and Wealth (SHIW), the Spanish Continuous Family Expenditure Survey (ECPF), the British Household Panel Survey (BHPS), and the US Panel Study of Income Dynamics (PSID).

Publicly available surveys usually report employment status and demographic characteristics of family members but only few of them contain information also on household consumption and intra-family transfers. We have chosen the surveys mentioned above precisely because they offer this additional information. Unfortunately, however, their design and the questions they ask differ substantially in some cases. Therefore, our attempt to extract comparable data sets for each country faces some constraints and the outcome suffers from several shortcomings. Yet we believe that our pooled data set provides sufficiently comparable and interesting information from the viewpoint of our research objectives.

Time structure of the surveys

A first potentially important comparability problem results from the fact that the temporal design of the surveys differs across countries. In Britain and the US the surveys take place at a yearly frequency. The BHPS exists since 1991 and we are able to use all the waves up to 1995. The PSID exists instead since 1968, but we decided to restrict the analysis only to the 1980s and 1990s. Within this period, the information needed for our purposes is available only in the 1980-86 and 1989-92 waves. The SHIW exists since the seventies but it has a panel structure with sufficient information only in 1991, 1993, and 1995. Since this structure imposes a one-year gap, we repeated the estimation of the change in consumption also allowing for such a gap in the other countries (except Spain, where the data do not allow for it) and the results were qualitatively the same as those reported in the text. Note also that the Italian sample has a partially rotating structure: some households are interviewed in all three years while others are interviewed only in a couple of years. Significantly more divergent is the design of the ECPF, which is a survey with a quarterly rotating structure. So, for Spain we have information on households observed for eight consecutive quarters in the period 1986-96. For the comparison between Spain and the other countries we have annualized the quarterly Spanish observations. In this way we obtained, for each Spanish household, two observations corresponding to two consecutive periods of four quarters each. Whenever Spain is analyzed we include in the estimation a set of dummies for the quarter in which a household begins to be observed.

Variables extracted from the surveys

The comparability of the information on unemployment, consumption, and demographic variables in the four countries is another issue of potential concern for the interpretations of the results presented in this paper. From each survey we extracted the following information.

1. *Unemployment.* As an indicator of the extent to which a household is affected by unemployment we use the number of months during which the male head is unemployed in each year. One household member answers for all members, which may introduce measurement error. It is reassuring to observe that the wording of the questions concerning employment status are very similar across countries. However, the definition of unemployment implied by these questions (i.e. not employed and searching at the time of the interview) is not necessarily equivalent to the official country-specific definitions. In all our data sets there exists a second type of employment status questions which gives information on whether the male head was unemployed at a precise date during a year. We exploit these questions to construct the instrument used in the IV estimations (see Section 3).

2. *Household consumption.* We would have liked to obtain indicators of total, durable, and non-durable consumption for all countries. Unfortunately, while the ECPF and the SHIW do contain that breakdown, the BHPS and the PSID offer information only on food expenditures. A sensible cross-country comparison is therefore possible only for food expenditures. These are deflated by the Consumer Price Index.

3. *Demographics.* This is the category of variables in which we encounter less comparability problems given the objective nature of the variables on which we focus. These are: the number of members, broken down into three groups: children aged less than 14 years old, children between 14 and 25 years old, and adults (male and/or female heads plus other members older than 25 years old); the fraction of females in the household; an indicator for the presence also of a female head; and the age and education of parents.

4. *Transfers.* It is admittedly difficult to obtain for different countries comparable measures of tranfers received by households from relatives. We did our best using the following information contained in the original data sets.

   Italy For each member of the household the questionnaire in a typical year of the survey asks: "In the year ... did ... (name of the member) receive scholarships, gifts or cash from relatives or friends not living in the house, alimony, or other income?". In case of a positive answer the questionnaire further asks to select the specific kind of transfer among:
   d1. Scholarship?
   d2. Gifts or cash from relatives or friends not living in the house?
   d3. Alimony?
   d4. Other?
   We considered the answer d2 as indicating the existence of tranfers from relatives.

   Spain The questionnaire of a typical year asks if the household received: Other irregular income different from Labor income, Capital income, Benefit income, Pension income, Self-employment income, Other regular income, and Non-classifiable income. We considered the existence of this kind of Other irregular income as indicating the existence of tranfers from relatives.

Britain The questionnaire asks the follwing question: "I am going to show you four cards listing different types of income and payments. Please look at this card and tell me if, since September 1st last year, you have received any of the types of income or payments shown, either just yourself or jointly?". We considered the answer "Payments from a family member not living here" as indicating the existence of tranfers from relatives.

United States The data contain three variables that indicate the "Amount of help received from relatives" respectively by the head of the household, the spouse, and all the other family members in a given year. The documentation says that "The values for this variable ... represent the amount of financial help received from relatives in whole dollars". We considered a positive amount for any of these variables as indicating the existence of transfers from relatives.

With the possible exception of Spain, where the information is less clean, the other datasets offer indicators that measure explicitly the extent of help received by households from relatives.

5. *Income.* We use the male head's labor income (including self-employment income if applicable), the female's labor income identically defined, household unemployment benefits, and total household income. In the US unemployment benefits are not observed and are therefore replaced by public transfers (excluding pensions).

Observations extracted from the surveys

From the original samples we select the observations used in the analysis on the basis of three sets of criteria. First, we keep the households in which a male head is present and for which our analysis is less likely to suffer from evident potential confounding factors. This implies excluding households in which the identity of the male head changes from year to year. These filters leave us with a sample of 104,206 household-year observations for the four countries, which we consider as our starting sample.

The second set of criteria requires the exclusions of all the observations for which one of the variables used in our analysis is missing or clearly wrong (e.g. negative or null consumption). These filters reduce the sample size by approximately 3%, leaving us with 100,997 observations. No country appears to be evidently more prone to loss of observations in this selection step.

The third set of criteria aims at eliminating outliers with respect to household income, which are likely due to misreporting. We drop within each country the top and bottom 1% of the real income distribution. As a result, the sample is further reduced to 99,114 observations, which implies an additional 2% loss. This (unbalanced) panel of household-years observations contains information on 32,714 households in the four countries.

Using this panel, we can construct 61,206 within-household yearly first-differenced observations. The sample finally used is obtained from these first differenced observations, with the additional restriction that the male household head be fully employed during year $t-1$. With this further restriction we obtain the 53,757 first differenced observations. Table A.1 describes the time structure of this sample, while Tables A.2 to A.5 report country-specific descriptive statistics of the variables used in the regressions and some other relevant variables in levels (referring to the second time observation of each difference).

Table A.1. Time structure of the data
(household-year observations)

year	Italy	Spain	Britain	US	Total
1981	0	0	0	3313	3313
1982	0	0	0	3294	3294
1983	0	0	0	3200	3200
1984	0	0	0	3366	3366
1985	0	0	0	3498	3498
1986	0	0	0	3633	3633
1988	0	351	0	0	351
1989	0	721	0	0	721
1990	0	703	0	3809	4512
1991	0	729	0	3864	4593
1992	0	710	2814	3704	7228
1993	2554	748	2681	0	5983
1994	0	724	2637	0	3361
1995	2647	727	2630	0	6004
1996	0	700	0	0	700
Total	5201	6113	10762	31681	53757

Table A.2. Descriptive statistics for the Italian panel
(5201 household-year observations)

Variable	Mean	St. Dev.	Min.	Max.
Δ real food consumption (%)	0.01	0.46	-3.27	2.72
Male head unemployed	0.02	0.13	0	1
Δ male head's months of unemployment	0.20	1.50	0	12
Number of adults	2.24	0.71	1	8
Δ number of adults	0.03	0.44	-4	4
Number of children $<$ 14 y.o.	0.53	0.85	0	5
Δ number of children $<$ 14 y.o.	-0.07	0.43	-3	3
Number of children 14–25 y.o.	0.58	0.85	0	6
Δ number of children 14–25 y.o.	-0.04	0.51	-3	2
Female rate	0.46	0.18	0	0.83
Δ female rate	-0	0.09	-0.75	0.75
Male head's age	53	13.49	24	91
Male head's education (years)	8.66	4.42	0	20
Home ownership	0.67	0.47	0	1
Wife present	0.93	0.26	0	1
Household receives transfers from relatives	0.09	0.28	0	1

Table A.3. Descriptive statistics for the Spanish panel
(6113 household-year observations)

Variable	Mean	St. Dev.	Min.	Max.
Δ real food consumption (%)	-0.03	0.28	-2.06	2.01
Male head unemployed	0.02	0.13	0	1
Δ male head's months of unemployment	0.15	1.06	0	12
Number of adults	2.35	0.72	1	7
Δ number of adults	0.01	0.33	-2	3
Number of children $<$ 14 y.o.	0.62	0.89	0	5
Δ number of children $<$ 14 y.o.	-0.05	0.32	-2	2
Number of children 14–25 y.o.	0.69	0.98	0	7
Δ number of children 14–25 y.o.	-0.01	0.41	-3	6
Female rate	0.48	0.17	0	0.86
Δ female rate	-0	0.06	-0.67	0.50
Male head's age	53.50	14.42	18	99
Male head's education (years)	6.34	3.82	0	17
Home ownership	0.83	0.38	0	1
Wife present	0.95	0.23	0	1
Household receives transfers from relatives	0.05	0.21	0	1

Table A.4. Descriptive statistics for the British panel
(10762 household-year observations)

Variable	Mean	St. Dev.	Min.	Max.
Δ real food consumption (%)	0.03	0.35	-2.86	2.17
Male head unemployed	0.03	0.16	0	1
Δ male head's months of unemployment	0.19	1.23	0	12
Number of adults	1.92	0.49	1	7
Δ number of adults	-0	0.27	-5	5
Number of children $<$ 14 y.o.	0.55	0.94	0	5
Δ number of children $<$ 14 y.o.	-0	0.30	-4	3
Number of children 14–25 y.o.	0.25	0.59	0	4
Δ number of children 14–25 y.o.	0	0.30	-3	2
Female rate	0.42	0.21	0	0.86
Δ female rate	-0	0.08	-0.80	0.80
Male head's age	49.51	16.09	19	93
Male head's education (years)	10.73	1.37	5	21
Home ownership	0.80	0.40	0	1
Wife present	0.83	0.37	0	1
Household receives transfers from relatives	0.01	0.09	0	1

Table A.5. Descriptive statistics for the US panel
(31681 household-year observations)

Variable	Mean	St. Dev.	Min.	Max.
Δ real food consumption (%)	-0	0.46	-5.86	5.05
Male head unemployed	0.02	0.14	0	1
Δ male head's months of unemployment	0.24	1.14	0	12
Number of adults	1.98	0.62	1	13
Δ number of adults	0	0.39	-7	6
Number of children $<$ 14 y.o.	0.83	1.10	0	8
Δ number of children $<$ 14 y.o.	0.01	0.41	-5	5
Number of children 14–25 y.o.	0.29	0.70	0	8
Δ number of children 14–25 y.o.	-0	0.35	-5	3
Female rate	0.43	0.21	0	0.89
Δ female rate	0	0.11	-0.80	0.80
Male head's age	43.14	15.31	17	95
Male head's education (years)	12.44	3.02	1	17
Home ownership	0.67	0.47	0	1
Wife present	0.84	0.36	0	1
Household receives transfers from relatives	0.03	0.18	0	1

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Endnotes

[1] Altonji and Villanueva (2005), using the 1988 PSID supplement, report that one third of American households give a transfer to their children. This is compatible with a much lower fraction of total households receiving transfers and refers to a set of parental households with a likely higher-than-average giving propensity (the youngest head is 48 years old, wealth must be observed, divorced parents are included, etc.).

Table 1: Unemployment benefit systems (1990-1995)[a]

	Replacement rates (% previous wage)				Maximum benefit duration (months)	Long-term unempl. (1990, %)	Unempl. benefit coverage (1991, %)
	Gross		Net				
	UI, first 6 mos.	Years 3 to 5	First month	60th month			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Italy	20	0.0	76	46	12	69.8	19
Spain	70	0.0	47	11	22	54.0	29
UK	38	15.7	67	76	44	34.4	62
US	50	4.7	59	51	12	5.6	n.a.

[a] Definitions and sources by column:
(1) Averages for the period 1989-1994. From Nickell (1997), Table 4. UI denotes unemployment insurance. (2) Measure with equal weights to replacement rates for years 3 to 5 in unemployment. From Blanchard and Wolfers (2000) Data Appendix, itself taken from OECD Database on benefits and entitlements. (3),(4) Replacement rates include unemployment, family, and housing benefits. Data for a married couple with two children and average production worker earnings. From OECD (1998), Tables 3.1 and 3.4. (5) Number of years over which a worker can get the maximum replacement rate. From Blanchard and Wolfers (2000) Data Appendix. (6) Unemployed for more than one year as a percentage of total unemployment. From OECD (1995), Table L. (7) Percentage of unemployed who report receiving benefits in the Labor Force Survey. From OECD (1994), Table 8.4.

Table 2: Change in the probability of a transfer (%) and male head's months of unemployment[a]

	Italy	Spain	Britain	US
Months of unemployment	1.4	0.4	0.0	0.5
	(0.2)	(0.2)	(0.1)	(0.1)
Dummy for change in unemployment benefits	-11.6	-1.4	-0.1	-1.0
	(3.8)	(1.1)	(0.2)	(0.4)
No. of observations	5201	6113	10762	31681

[a] For each country the table reports, in percentage form, probit estimates of the effect of the number of months of unemployment experienced by the male head of the household, ΔU_{it}, and of a 0-1 dummy variable for a positive change in real unemployment benefits received by the household, on the probability that the household receives a transfer from relatives. A different regression is estimated for each country. The sample is restricted to households in which the head was never unemployed at $t-1$. The regression also includes the changes in the number of children aged less than 14, in the number of children aged between 14 and 19, in the number of other household members, and in the fraction of females in the household, as well as the levels of the age and years of schooling of the male head, and a dummy for home ownership (lagged). Year dummies (quarter dummies for Spain) are included as well. Descriptive statistics for the household samples used in the regressions are given in Tables A.2 to A.5. Robust standard errors, adjusted for within-household serial correlation, are reported in parentheses.

Table 3: Food consumption change (in %) and male head's months of unemployment[a]

	Italy	Spain	Britain	US
IV	-0.8	-0.2	-1.3	-1.6
	(0.5)	(0.4)	(0.6)	(1.1)
OLS	-0.8	-0.4	-0.6	-0.9
	(0.5)	(0.3)	(0.3)	(0.3)
F-test on IV in first stage	164.0	102.9	65.7	58.1
p-value of test for OLS = IV	0.91	0.41	0.15	0.53
No. of observations	5201	6113	10762	31681

[a] The table reports, in percentage form, OLS and IV estimates of the coefficient γ in equation (4). A different regression is estimated for each country. The dependent variable $\Delta \ln C_{it}$ is the change in the log of food consumption of each household. γ is the coefficient attached to variable ΔU_{it}, which measures the change in the number of months of unemployment experienced by the male head of the household. The sample is restricted to households in which the head was never unemployed at $t-1$. The IV estimates use as instrument an indicator of the unemployment status of the male head, constructed on the basis of a question different from the one used to compute ΔU_{it}. F-tests for the inclusion of the IV in the first stage are reported (see Staiger and Stock 1997). The test for the equality of OLS and IV estimates is the one proposed by Arellano (1993) and we report the p-value. The regression also includes the changes in the number of children aged less than 14, in the number of children aged between 14 and 19, in the number of other household members, and in the fraction of females in the household, as well as the levels of the age and years of schooling of the male head, and a dummy for home ownership (lagged). Year dummies (quarter dummies for Spain) are included as well. Descriptive statistics for the household samples used in the regressions are given in Tables A.2 to A.5. Robust standard errors, adjusted for within-household serial correlation, are reported in parentheses.

Table 4: Evidence on the compressibility of food consumption[a]

	Italy		Spain		Britain		US	
	Centile	Order	Centile	Order	Centile	Order	Centile	Order
P10	-51.0	10th	-35.4	10th	-36.2	10th	-47.4	10th
Became unemployed	-8.8	47th	-4.3	46th	- 0.4	43th	-4.8	43th

[a] The first row reports the 10th percentile (P10) of the distribution of percent changes in food consumption between $t-1$ and t for each country. The second row reports the average percent change in food consumption for households in which the male head became unemployed at t (and was fully employed at $t-1$). It also reports the order of the percentile to which this figure corresponds in the distribution of consumption changes of each country.

Table 5: Descriptive statistics on unemployment shocks and the importance of the male head's labor income[a]

	Italy	Spain	Britain	US
Male head's months of unemployment for heads with positive months	11.47	5.54	5.26	3.20
(Std. deviation)	(1.52)	(3.26)	(3.80)	(2.82)
Male head's lagged income importance	0.42	0.51	0.48	0.61
(Std. deviation)	(0.35)	(0.40)	(0.36)	(0.34)
Male head's months of unemployment $\times$ lagged income importance	0.08	0.10	0.09	0.15
(Std. deviation)	(0.81)	(0.73)	(0.70)	(0.76)
No. of observations	5201	6113	10762	31681

[a] For each country the table reports, the mean and standard deviation of the male head's months of unemployment for those with some unemployment at t, the importance of his labor income in total family income at $t-1$, and the product of the two variables (with the former variable not restricted to heads with some unemployment at t). The sample is as described in the footnote to Table 3.

Table 6: The impact of changes in the male head's labor income and unemployment shocks on total household income[a]

	Italy	Spain	Britain	US
Change in male head's labor income	0.85	0.74	0.95	0.90
	(0.03)	(0.05)	(0.01)	(0.02)
Change in male head's labor income	-0.008	-0.006	-0.008	-0.010
× number of months of unemployment	(0.009)	(0.013)	(0.007)	(0.005)
Male head's months of unemployment	86.25	20.70	64.30	126.34
	(130.18)	(10.21)	(57.03)	(50.82)
No. of observations	5201	6113	10762	31681

[a] The table reports country-specific OLS estimates of the coefficients in equation (6). The dependent variable ΔY_{it} is the change in total household income. The regressors are the change in the male head's labor income, ΔYL_{it}^{h}, the change in the number of months of unemployment experienced by the male head, ΔU_{it}, and the product of the two variables, ΔU_{it}^{*}. Income variables are in real terms and in national currency (for Italy in thousand lira, for Spain in thousand pesetas). The sample and other regression characteristics are as described in the footnote to Table 3. Robust standard errors, adjusted for within-household serial correlation, are reported in parentheses.

Table 7: Food consumption change (in %) and importance-adjusted male head's months of unemployment[a]

	Italy	Spain	Britain	US
IV	-1.6	-0.3	-2.9	-2.4
	(1.1)	(0.7)	(1.3)	(1.7)
OLS	-1.9	-0.8	-1.2	-2.2
	(1.0)	(0.5)	(0.6)	(0.5)
Estimated effects (IV)	-4.7	-0.9	-8.7	-7.2
F-test for IV	109.0	81.3	50.2	57.9
p-value of test for OLS = IV	0.57	0.25	0.15	0.90
No. of observations	5201	6113	10762	31681

[a] The table reports, in percentage form, OLS and IV estimates of the coefficient γ in equation (4) with ΔU_{it}. replaced by ΔU^*_{it}. A different regression is estimated for each country. The dependent variable $\Delta \ln C_{it}$ is the change in the log of food consumption of each household. γ is the coefficient attached to variable ΔU^*_{it}, which measures the change in the number of months of unemployment experienced by the male head of the household multiplied by his labor income importance at $t-1$. The sample is restricted to households in which the head was never unemployed at $t-1$; hence ΔU^*_{it} is non-negative. To take care of measurement error in ΔU^*_{it}, the IV estimates use as instrument an indicator of the unemployment status of the male head, constructed on the basis of a question different from the one used to compute ΔU^*_{it} (see Section 3). F-tests for the inclusion of the IV in the first stage are reported (see Staiger and Stock 1997). The test for the equality of OLS and IV estimates is the one proposed by Arellano (1993) and we report the p-value. The regression also includes the changes in the number of children aged less than 14, in the number of children aged between 14 and 19, in the number of other household members, and in the fraction of females in the household, as well as the levels of the age and years of schooling of the male head, and a dummy for home ownership (lagged). Year dummies (quarter dummies for Spain) are included as well. Descriptive statistics for the household samples used in the regressions are given in Tables A.2 to A.5. Robust standard errors, adjusted for within-household serial correlation, are reported in parentheses. The reported effects of ΔU^*_{it}, based on the IV estimates, are computed for a male head with 6 months of unemployment and with 50% importance of lagged income.

Table 8: The impact of the male head's labor income and unemployment shocks on the female head's labor income[a]

	Italy	Spain	Britain	US
Change in male head's labor income	0.043	0.007	0.002	0.008
	(0.011)	(0.014)	(0.008)	(0.006)
Change in male head's labor income	-0.007	-0.001	-0.008	-0.003
$\times$ number of months of unemployment	(0.004)	(0.003)	(0.005)	(0.003)
Male head's months of unemployment	14.219	-1.796	-11.470	26.544
	(79.042)	(3.836)	(31.580)	(38.651)
No. of observations	4811	5779	8895	26088

[a] For each country the table reports OLS estimates of the coefficients in equation (7). The dependent variable ΔYL_{it}^{w} is the change in the female head's labor income. The regressors are the change in the male head's labor income, ΔYL_{it}^{h}, the change in the number of months of unemployment experienced by the male head, ΔU_{it}, and the product of the two variables. Income variables are in real terms and in national currency (for Italy in thousand lira, for Spain in thousand pesetas). The sample and other regression characteristics are as described in the footnote to Table 3. There are fewer observations in this table, since only observations with a wife present are included. Robust standard errors, adjusted for within-household serial correlation, are reported in parentheses.

Table 9: Food consumption change (in %) and importance-adjusted male head's months of unemployment, controling for female head's and children labor supply[a]

	Italy	Spain	Britain	US
IV	-1.7	-0.2	-2.9	-2.5
	(1.1)	(0.7)	(1.3)	(1.7)
OLS	-2.0	-0.7	-1.2	-2.3
	(1.0)	(0.4)	(0.6)	(0.5)
F-test for IV	109.0	81.2	50.1	57.8
p-value of test for OLS = IV	0.51	0.24	0.15	0.89
No. of observations	5201	6113	10762	31681

[a] The sample and regression characteristics are as described in the footnote to Table 7. The only difference is that, to control for the female head's labor supply, dummy variables are included for transitions from working to not working, not working to working, and working in both periods (female heads not working in either period form the reference group). Moreover, to control for children's labour supply we include also the change in the number of working children in the household. For IV estimates, F-tests for the inclusion of the IV in the first stage are reported (see Staiger and Stock 1997). Robust standard errors, adjusted for within-household serial correlation, are reported in parentheses.