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Abstract

This note provides an extensive survey of studies estimating steady-state labor supply elasticities for Western Europe and the US. Differences are driven by the heterogeneity in work preferences across countries and by methodological difference across studies (data, selection or model estimation and specification). While the former exists but is shown to be relatively small (Bargain et al., 2013), we focus here on modeling choices: Large elasticities are mainly found in studies estimated in the 1980s and relying on the Hausman approach. More recent estimates based on discrete-choice models with tax-benefit simulations show smaller and more similar estimates across countries. While we confirm that elasticities decline over time in the US, there is some evidence that both time effect and modeling choices affect estimates for Europe.

Key Words: household labor supply, elasticity, taxation, Europe, US.
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1 Introduction

Static models of labor supply are very useful to predict ex ante the effect of tax-benefit policy reforms or more generally to provide an order of magnitude of the short-term response to financial incentives. Responsiveness is often summarized by a measure of what Chetty et al. (2011) refer to as "steady-state elasticities", i.e. wage or income elasticities of labor market participation or worked hours stemming from a static framework. These estimates are useful in many contexts, for instance to calibrate an optimal tax model, and in principle, they allow comparing labor supply responses across countries. However, many factors may affect the differences in the size of elasticities across studies. The variation in magnitude of labor supply elasticities found in the literature is huge (see Evers et al., 2008), and there is little agreement among economists on the elasticity size that should be used in economic policy analyses (Fuchs et al., 1998).

Several excellent surveys report evidence on elasticities for different countries and different periods. Those written in the 1980s mainly focus on estimations using the continuous labor supply model of Hausman (1981) and provide evidence essentially for individuals in couples (Hausman, 1985b, Pencavel, 1986, for married men, Killingsworth and Heckman, 1986, for married women). More recent surveys incorporate some evidence from recent methods (see Blundell and MaCurdy, 1999; Meghir and Phillips, 2008) or focus on life-cycle models (Keane, 2011; Keane and Rogerson, 2012; McClelland and Mok, 2012). Most of these surveys mainly summarize the available evidence for the US and the UK. Evers et al. (2008) suggest a meta-analysis based on estimates for different Western countries, focusing essentially on those obtained with the traditional Hausman approach.

This note attempts to compare (more) broadly the international evidence on steady-state labor supply elasticities. We collect old and recent estimates for Europe and the US, covering the studies based on the Hausman method, more recent ones based on discrete-choice structural models and, when available, estimates drawn from natural experiments.[1] We acknowledge that differences across studies can be driven by differences in work preferences across countries and by methodological differences (data selection and period of investigation,

[1] We focus on labor supply decisions (hours and participation). Hence, we ignore the other margins that are captured in the literature on the elasticity of taxable income (see Meghir and Phillips, 2008, and Saez et al., 2012, for surveys). Arguably, these other margins partly relate to responses not directly pertaining to productive behavior, like tax evasion and tax optimization. In this regard, hours of work still constitute an interesting benchmark. Another margin is work effort that may affect wage rates. In the short run, however, hours and participation are the only variables of adjustment for a large majority of workers. We also leave aside the macroeconomic literature, in which elasticities are often obtained by calibration of general equilibrium models. These elasticities are much larger than in microeconomic studies (e.g., Prescott, 2004). Several reasons have been suggested for this: the use of representative agents and difficulties around aggregation theory when heterogeneity matters (see Blanchard, 2006), the existence of a social multiplier whereby the utility from not working is increasing in the number of people who do not work (see Alesina et al., 2005), and factors related to the timing and the nature of labor supply adjustments (Chetty et al., 2011).

estimation method and model estimation and specification, etc.). Measuring international differences in consumption-leisure requires using a uniform approach for many countries. This exercise, undertaken by Bargain et al. (2013) for the EU and the US, shows that cross-country variation exists but is relatively small – at least smaller than what is expected from comparing estimates in the literature. In the present companion paper, we therefore focus on the other sources of difference across studies, namely modeling choices. To do so, we compare 214 elasticity estimates resulting from 110 estimations, i.e. 86 estimations in 57 studies on married men and women and 24 studies on single individuals (with or without children).

Our survey substantially completes previous reviews on static labor supply models, which, as stated above, concentrate mainly on evidence from the Hausman model, for the 1980s and 1990s and for Anglo-Saxon countries. Our results go as follows. First, we broadly confirm the modest consensus reached in the literature, establishing that own-wage elasticities are largest for married women, smaller for men. Recent studies confirm these findings but not negative elasticities for men as sometimes found in older studies. Estimates for men are generally positive and small, with some exceptions (for instance Ireland and some German studies). Some of the studies for the US and the UK, but not all, point to substantial elasticities for single parents while estimates for childless singles are usually missing. Second, for each demographic group, we observe a very large variance in estimates across all available studies. This is partly due to the use of the Hausman approach, which seem to overstate elasticities compared to what is found with more recent approaches and notably the use of discrete choice models. The other main methodological factor behind different estimates is the time period. For the US, we corroborate the findings of Heim (2007) and Blau and Kahn (2007), who show, using a uniform approach for different periods, that married women's wage-elasticities decline over time. Given that the use of the Hausman method coincides with older studies, it is nonetheless difficult to disentangle the two factors. Restricting our meta-analysis to years of common support, we find evidence in favor of both explanations for Europe. This means that the results of Heim and Blau and Kahn might be generalized to EU countries. For the US, there is no clear evidence that estimation methods matter – in fact, estimates from discrete choice modeling are missing for the long period and should be the subject of future research.

The paper is organized as follows. In Section 2, we describe the various empirical approaches. Section 3 reports the survey results. Section 4 concludes.

2 Methods: A Critical Review

The principal object of examination in this study is the size of wage and income elasticities stemming from static labor supply models. Responsiveness to financial incentives in these models has been identified in various ways. There is no generally agreed-upon standard estimation approach and we provide here a brief critical review. A more technical and comprehensive presentation of these methods and their identification strategies are provided in Blundell and MaCurdy (1999) and Blundell et al. (2007).

Traditional estimation techniques rely on some functional specification of a labor supply function and the underlying consumption-leisure preferences. Estimation is then made through local linearization of the budget constraint, accounting for the fact that after-tax wages depend on the labor supply choice (Hall, 1973) or using more comprehensive techniques (Hausman, 1981,1985a, 1985b). The approach relies on cross-section variation in working hours and in the two main covariates, i.e. the after-tax wage and the virtual income (i.e. the intercept of the linearized budget constraint). As a result, the main identification issue is the endogeneity of wages and unearned income, which can be seen as an omitted variable problem. Indeed, wages may be endogenous because unobservables affecting preferences for work, e.g., being a hard-working person, may well be correlated with unobservables affecting productivity and hence wages. Unearned income may be endogenous for similar reasons, i.e. individuals who work harder because of unobserved preferences for work are also likely to have accumulated more assets; if unearned income also represents income from the spouse, positive assortative mating could imply that hard working individuals will tend to marry similar persons, another reason for the endogeneity issue. Hence, estimates obtained from cross-sectional variation in wages and nonlabor income across individuals are potentially biased. Instrumental variables methods have been suggested and the validity of the Hausman approach hinges on whether the exclusion assumptions of the economic model hold. Also, estimates are potentially contaminated by measurement errors from the division bias (cf. Ziliak and Kniesner, 1999). In addition, a series of practical difficulties limit the application of the method. First, relying on tangency conditions, the Hausman model is mainly restricted to the case of piecewise linear and convex budget sets, i.e., a partial representation of the effect of tax-benefit policies on household budget constraints. This limitation applies equally to generalizations of the technique to non-parametric estimations (Blomquist and Newey, 2002). To account for nonconvexities, as in Hausman (1985b) and Hausman and Ruud (1984), labor supply must be specified parametrically together with the corresponding direct utility function, which implies rather restrictive forms for preferences (see the discussion in Van Soest and Das, 2001).[2] Second, quasi-concavity of the utility function is implicitly imposed *a priori*. As discussed by MaCurdy et al. (1990) and MaCurdy (1992), the Hausman method thus requires global satisfaction of the Slutsky condition by the labor supply function for internal consistency of the model, an unnecessary behavioral restriction

[2] Another approach is the reconvexification of the budget set. For instance, to estimate the labor supply of married women on 1985 French data, Bourguignon and Magnac (1990) use the Hausman technique and eliminate minor nonconvexities by replacing the budget set by its convex envelope. This approach is not possible for later years as the implementation of a minimum income scheme in 1988 has introduced high nonconvexity in the budget constraint. Similar nonconvexities arise in all countries with substantial means-tested transfers.

that may bias estimates (see a modern statement in Heim and Meyer, 2003, and Meghir and Phillips, 2008). Third, the model makes it difficult to handle joint labor supply decisions within a couple or participation decisions. Instead of non-participation following simply from the corner solution of the model, fixed costs of work can be introduced, yet this additional source of nonconvexity has to be dealt with and results seem to be very sensitive to the model specification (see the discussion in Bourguignon and Magnac, 1990).

Instead of estimating a labor supply function, the discrete choice approach is based on the concept of random utility maximization (see van Soest, 1995, or Hoynes, 1996, among others). Thus, it requires the explicit parameterization of consumption-leisure preferences, for utility to be evaluated at each discrete alternative. Tangency conditions need not be imposed and the model is in principle very general. Labor supply decisions are reduced to choosing among a discrete set of possibilities, e.g., inactivity, part-time and full-time. This solves several problems encountered with the Hausman method. In particular, discrete choice modeling includes non-participation as one of the options so that both extensive and intensive margins are directly estimated. The complete effect of the tax-benefit system is easily accounted for, even in the presence of nonconvexities in budget sets. Work costs, which also create nonconvexities, are dealt with relatively easily. Estimated as model parameters as in Callan et al. (2009) or Blundell et al. (2000), they usually improve the fit of these models as they account for the fact that very few observations exist with a small positive number of worked hours. Very few restrictions on preferences need to be imposed in discrete choice models, notably because fixed costs of work cannot be disentangled from preference parameters, so that it makes no sense to impose the convexity of preferences (see van Soest et al., 2002, Heim and Meyer, 2003, Bargain, 2009). The only restriction to the model is the imposition of increasing monotonicity in consumption, which seems a minimum requirement for meaningful interpretation and policy analysis. Joint labor supply decision for couples is a straightforward extension of the basic model in the discrete choice setting. Yet, many applications still treat husbands' working hours fixed at observed levels and focus on the labor supply of women, i.e. a male chauvinist model (e.g., Bargain, 2009; such treatment is typical in Hausman models, e.g. Killingsworth and Heckman, 1986). The implication of such separable treatment of spouses' labor supply choices is relatively unknown.
In the discrete choice approach, identification is mainly provided by nonlinearities, nonconvexities and discontinuities in the budget constraint due to tax-benefit rules (see the discussion in Blundell et al., 2007, and Bargain et al., 2013). Precisely, individuals with the same gross wage usually receive different net wages. Indeed, as they are characterized by different circumstances (different marital status, age, family compositions, home-ownership status, disability status) or levels of nonlabor income, their effective tax schedules are different, i.e., different actual marginal tax rates or benefit withdrawal rates. Arguably, some of the conditioning characteristics (age, children) are also included as preference variables in the model so that identification is essentially parametric. In practice, some exclusion

restrictions come naturally. Indeed, tax-benefit rules depend on characteristics which are much more detailed than usual taste-shifters (e.g. benefit rules depending on detailed geographical information while preferences are assumed to depend only on urban versus rural areas or on whether the household lives in the capital city). Additional, more convincing sources of exogenous variation are also used in some studies. Closer to the natural experiment method, these consist in time or regional variation in tax-benefit rules. For instance, in the US, variation in income tax rules or in the parameters of the Earned Income Tax Credit (EITC) across states is used in Eissa and Hoynes (2004) or Hoynes (1996). Time variation in tax-benefit rules also provide a better identification when policy reforms occur over the period under consideration, as discussed, e.g., in Bargain et al. (2013)

A third approach consists in using policy reforms explicitly in order to identify labor supply responses, without attempting to estimate a structural model (e.g., Eissa and Liebman, 1996). Natural experiments based on important tax-benefit reforms in the US and the UK have been extensively used to identify behavioral parameters (see the survey of Hotz and Scholz, 2003, for the US). For example, Eissa and Liebman (1996) use a difference-and-difference approach to identify the impact of the EITC reforms on the labor supply of single mothers. They find compelling evidence that single mothers joined the labor market in response to increased financial incentives to work. There is less evidence for other countries and notably for continental Europe so that structural models are still much in use.[3] The timing of response to such policy reforms or policy discontinuity is unclear. Nonetheless, the implicit model that analysts have in mind when discussing the "next-morning" effect of the policy impact is often a static one (cf. Lemieux and Milligan, 2008, or Bargain and Doorley, 2011). This reduced-form approach is increasingly used because natural experiments probably offer one of the most credible sources of identification. These studies do not systematically report wage elasticities, however. They rather report labor supply elasticities to benefit or tax rate changes. Thus, for comparability purposes, we could include only a few of them in the present survey. Also, the fact that actual reforms – notably welfare reforms in the US and the UK – typically affect couples or single women with children makes that very little evidence is available for other demographic groups, in particular for childless single individuals. Regarding identification, the definition of control groups might be an issue in difference-in-difference approaches. For instance, responses to EITC expansions affecting single mothers were evaluated using childless women as control group, which may not be ideal given different long-term trends in labor supply in the two groups (see Hotz and

[3] Things are changing in the recent period. For France, for instance, some studies have recently used tax-benefit changes to evaluate the responsiveness of the labor force, including the introduction of a small tax credit (Stancanelli, 2008), time change in income tax schedule (Carbonnier, 2008), changes in the possibility to cumulate welfare payment for lone mothers and earnings (González, 2008), and age condition on children for a replacement income targeted at low-income mothers who opt for full-time childcare (Piketty, 1998). RD estimations using age conditions on the level of social assistance program are also used in Bargain and Doorley (2011), in a similar way as Lemieux and Milligan (2008) for Canada.

Scholz, 2003).[4] RD experiments are deemed better in this respect since the nature of individuals on both sides of the discontinuity is "as good as random" (cf. Lemieux and Milligan, 2008). Finally, a few studies rely on long-term changes in wages as well as on observation grouping in order to address endogeneity and the problem of measurement error in hourly wages discussed above (Devereux, 2003, 2004). Blundell et al. (1998) also use tax-benefit policy variation over the long period to identify labor supply responses in the UK. Long-term variation may pose the problem of assuming that preferences remain stable in the long run, an issue which is rarely discussed.

3 Static Labor Supply Elasticities: A Survey

We present here existing evidence on labor supply elasticities for European couples (Tables 1 and 2), European single individuals (Table 3) and all demographic groups in the US (Table 4). The reason for this classification is that US studies are more numerous (and hence deserve a particular focus) and sometimes consider several demographic groups simultaneously (e.g. Pencavel, 2002, Devereux, 2003). We distinguish wage-elasticities (total hour and participation responses) and income-elasticities. The survey essentially distinguishes between estimates based on models with a continuous labor supply function (using the Hausman approach), discrete choice models and grouped estimations / natural experiments. We put a certain emphasis on the studies based on discrete choice models with taxation, as this method is increasingly used around the world to analyze the effect of fiscal and social policy reforms. Yet we do not pretend to be exhaustive, simply to give a sense of the range of elasticities obtained in the literature for Europe and the US. Some studies do not report elasticities and unfortunately could not be included in our tables. This is the case with some studies using labor supply models (e.g., Hoynes, 1996, does not report wage-elasticities) and more generally the case with studies using policy reforms as natural experiments, as indicated above (for instance Bingley and Walker, 1997, for the UK, or Eissa and Liebman, 1996, for the US).

[4] This issue is shared with the literature on the elasticity of taxable income, whereby results are sensitive to the type of reforms exploited for identification (Saez et al., 2012). Indeed, control groups definition follows from their income level, so that specific preferences are identified and results cannot be extrapolated. For instance, changes in tax rates (tax credits) identify the preferences of high (low) income groups, and may not be generalized to the whole population.

3.1 Overview

Figure 1 plots the distribution of wage-elasticity estimates by demographic group.[5] The vertical axis reports the frequency (number of estimates). The first observation is that married women is the group with the largest number of available estimates. The second lesson from these graphs is that in line with conventional wisdom, elasticities are largest among married women and single mothers, with mean values of .48 and .53 respectively. These groups also show much dispersion across available studies. Married and single men (mean value: .11) and childless single women (mean value: .15) show much less variation, with most estimates between 0 and .30. These conclusions do not change radically if we look separately at total hour or participation elasticities (detailed results available from the authors). We now discuss each group specifically.

Married Women. Considering Tables 1, 2 and 4, we observe much dispersion in estimates. This is confirmed in Figure 2 where we plot, for each demographic group, the distribution of wage-elasticity estimates for each country. The grey triangular indicates the mean value over all available estimates. Mean elasticities for the UK and the US hide a very broad dispersion across studies. Large elasticities for France may be driven by methodological reasons as discussed below. As far as genuine international differences are concerned, we suggest that larger wage-elasticities prevail in countries where women's participation is low: This seems to be the case in our survey estimates for Ireland and Italy, which is confirmed in the discussions in Callan et al. (2009) and Aaberge et al. (2002) for these two countries respectively. In contrast, women's participation is high in Nordic countries and elasticities tend to be fairly small there, notably in Finland and Sweden. An exception is Blomquist and Hansson-Brusewitz (1990) for Sweden, but the authors examine data from the 1980s, while more recent evidence by Flood et al. (2004) confirm small hour elasticities for this country. Comparing Italy and Norway/Sweden, Aaberge et al. (1999) show that lower participation rates among married women in Italy leads to a larger potential for reforms that increase financial incentives to work. Larger elasticities coincides with more intermittent labor force participation patterns in Southern countries and Ireland as opposed to more consistent participation and more constant hours in Scandinavian countries. Apart from these extreme cases, differences across EU countries, and notably countries of Continental Europe, may not be very large, as suggested by Evers et al. (2008). This is confirmed by Bargain et al. (2013): Using an harmonized framework for 17 EU countries and the US, they find estimates for married women ranging in a narrow interval $.2 - .6$. This is indeed where mean values lie in Figure 2 (top left quadrant), with few exceptions. As argued above, direct comparisons across studies are muddled by methodological differences (notably the period

[5] All figures reported in this study are based on the estimates of wage elasticities from Tables 1-4. For comparability purpose, we exclude estimates based on model without taxation, based on long term wage variation or with too specific selection (e.g. estimates on couples with children only, as in Choné et al., 2003). We lose 15 estimates, i.e. around 10% of our sample.

of investigation and the estimation approach). We investigate this point in more detail in the next sub-section.

Single Mothers. Most studies on single mothers are available for the UK (see Tables 3) and, to a lesser extent, for the US and Sweden. This demographic group has received much attention in the literature because of its importance for welfare analysis, given its higher risk of poverty, and because single parent families were primarily concerned by reforms like tax credit extensions in the US (cf. Hotz and Scholz, 2003) or the UK (Blundell et al., 2000). This group is found to be more responsive to financial incentives than the average, at least in the UK, the US and Sweden. This is confirmed in Tables 3 and 4, where relatively large elasticities are shown in several studies – but not all. Similar to the result for married women, much variance is found across studies, as illustrated in Figure 2 (bottom right quadrant). Moderate estimates are found in some studies for the UK (Blundell et al., 1992) and the US (Dickert et al., 1995) while other studies point to much larger elasticities (e.g., Keane and Moffitt, 1998, for the US or many of the British studies). Importantly, the size of this group has become much larger in the recent period in Anglo-Saxon countries, which implies possible changes in the selection effects. That is, this group may be less negatively selected in terms of labor market participation in the recent period. For the US, Bishop et al. (2009) study all single women over a long period (1979-2003), using a simple estimation of hours and participation on repeated cross-sections. They report a significant decline in hour wage-elasticities over the period and relatively small elasticities in the recent years (at least compared to typical estimates for married women).

Married Men and Childless Singles Individuals. There is a long history of estimating male labor supply (see surveys of Hausman, 1985b, and Pencavel, 1986, for married men). Estimates of wage-elasticities for this group are usually very small, often not significant and sometimes negative. There are few exceptions, with larger elasticities in Ireland and in some of the German studies, as seen in Tables 1 and 2 for married men. Evidence for childless single men and women, gathered in Table 3, is very limited, despite the growing proportion of this demographic group in the population. Limited evidence is essentially explained by methodological reasons. First, estimates are usually more precise for couples or single mothers than for childless single individuals. This can be due to the fact that there is less variation in labor market behavior among childless singles or that non-participation corresponds more often to demand-side constraints (rather than to voluntary choice) in their case. This argument equally applies to single men – yet the fit of labor supply model for married men should be overall better when male and female decisions are jointly estimated. Second, estimates stemming from natural experiments are also limited for this group, given the fact that most welfare reforms in Anglo-Saxon countries concerned individuals or households with children (see the discussion in Bargain and Doorley, 2011). The few available estimates point to very small elasticities.[6] For both men (married or single) and childless single women, estimates are not only small but very concentrated across studies with countries. This small variance is illustrated in Figure 2 (top right quadrant for men and bottom left quadrant for childless single women). Nonetheless, these mean values may hide much variaton in participation responses across different wage or income levels, with important implications for welfare analysis as suggested by see Eissa and Liebman (1996) and confirmed for single individuals in Bargain et al. (2013).

Income Elasticities. Most studies show negative income elasticities, i.e. positive income elasticities of non-market time. Yet, despite being at odds with theory, positive income elasticities are encountered in some studies, which include Kuismanen (1997) for Finland, Flood and MaCurdy (1992) for Sweden, van Soest (1995) for the Netherlands and Blau and Kahn (2007) and Cogan (1981) for the US. Also, despite being generally small, income elasticities vary across countries. Blundell and MaCurdy (1999) report that variation between studies regarding income elasticity appears to be greater than the corresponding variation with respect to wage elasticities. This is not confirmed in the extensive study by Bargain et al. (2013).

3.2 Year of Observation and Estimation Methods

In Tables 1-4 and Figures 1-2, we have observed lots of variation across studies in the size of wage-elasticities for married women and single mothers. This may correspond to genuine international differences in preferences. Using a uniform approach, Bargain et al. (2013) show that cross-country variation is small, however. Therefore, most of the heterogeneity across studies must be driven by various methodological choices and in particular the period of observation and the estimation method. We focus on these aspects hereafter, looking at the two groups showing most variation across studies, namely married women and single mothers.

Time Trend. In Figure 3, we plot estimates by data year. A very clear declining trend emerges, showing in particular a concentration of low elasticities in the recent periods (year 2000s). The left quadrant shows for all countries that this pattern is especially driven by married women, while the trend is not so strong among single mothers. Therefore, in the right quadrant, we focus on married women and now distinguish between EU and US estimates. The trend is similar in both regions, with a strong negative correlation between the period of observation and the elasticity level. These findings tend to corroborate the result of Heim (2007) and Blau and Kahn (2007), who show that the labor supply elasticity of married

[6] For instance, Euwals and Van Soest (1999) report wage elasticities for childless single individuals in the Netherlands of around $.10 - .11$. For Germany, a series of studies report estimates between .10 and .36 for childless single men and women.

has strongly declined over time in the US, and extend it to EU countries. We also find similar results when looking separately at hour wage-elasticities (correlation of $-.63$) and participation wage-elasticities (correlation of $-.54$). Yet, results in Heim (2007) and Blau and Kahn (2007) rely on a uniform approach for the different periods while our meta-analysis possibly mixes time effects (including selection effects) and changes in estimation methods over time.

Estimation Methods. To investigate this point further, let us get back to survey tables 1-4. A first observation is that early evidence using the Hausman technique points to relatively large own-wage elasticities for married women, sometimes close to 1, or even larger, for instance in early studies for France, Germany, Italy or the UK. In contrast, recent evidence based on discrete-choice models shows more modest elasticities for this demographic group, in a range between .1 and .5, with some exceptions. In Table 4, we observe a similar pattern for the US, with very large estimates in early studies, including Hausman (1981), and more modest and comparable elasticities in the recent studies (hour elasticities ranging between .2 and .4). Hence, we can conjecture that the estimation method explains time differences. With the Hausman approach, the combination of restrictive functional forms (linear labor supply) and estimation methods that impose theoretical consistency of the labor supply model everywhere in the sample (global satisfaction of Slutsky conditions) can lead to biased estimates and possibly an overstatement of work incentives, as discussed above. Mroz (1987) also shows how the wage effects of married women's labor supply varies dramatically depending on whether and how one controls for nonrandom selection into work as well as to alternative exclusion restrictions in the instrument set for wages. Bourguignon and Magnac (1990) discuss the sensitivity of their results to the model specification and show that the Hausman approach can lead to implausibly high elasticity values, as they find in some of their specifications. Drawn from our Table, we can see for instance that married women's wage elasticity obtained with the Hausman approach vary from .28 (Triest, 1990) to .97 (Hausman, 1981) in the US, even when similar periods are considered (1983 and 1975 in these two studies respectively). For France, estimates for married women are also very high with the basic Hausman model, but almost zero when introducing fixed costs (cf. Bourguignon and Magnac, 1990). Estimates obtained with discrete choice models are somewhat more comparable from one study to the next. Yet there are still differences, which are more likely driven by selection criteria (for France, high elasticities are found for families with children in Choné et al., 2003).

Meta-Analysis. To clarify whether the conjecture is true, we plot time trends according to two broad modeling choice in Figure 3 (upper panels), namely estimates obtained with continuous models (which rely mainly on the Hausman approach for identification) and those from discrete-choice models (as recently used in many policy papers). We first consider trends obtained with the continuous model. For our group of interest, and whether single mothers are included or not, the time shrinking elasticity hypothesis is verified over all estimates relying on the Hausman approach. When differentiating between regions (Figure 3, lower panels), the meta-analysis corroborates the find in Heim (2007) and Blau and Kahn (2007) for the US (both studies relying on a Hausman-type approach). A similar pattern is found for EU estimates but it is noticeable that there are very few estimates based on the Hausman model for the period after 1990, so the result is more fragile than for the US.[7] Then, we consider estimates from discrete choice model estimations. We first observe the lack of clear pattern in this case. There is nonetheless a negative linear correlation ($-.36$ for married women) between years and estimates due to the high density of low estimates after the end of the 1990s. The correlation becomes positive if we focus on the years before 1998 ($+.31$). This does not mean that the two main modeling strategies have opposite conclusions. In fact, there is a clear lack of common support (years) between the two, essentially due to the fact that very few estimates based on discrete models are available for the early period. If the shrinking elasticity trend is driven by a change in preferences (and correlated with an increase female labor market participation) precisely between the 1970s/1980s and the 1990s/2000s, then it cannot be captured by the available estimates based discrete choice modeling.[8] Finally, we restrict our sample to years equal or above the data year corresponding to the first estimate obtained with a discrete model (estimates on CPS 1985 in Eissa and Hoynes, 2004, and on the Dutch Labor Mobility Survey in van Soest et al., 1990). Thus we can conduct a meta regression over the years for which estimates with both discrete and continuous models are available. We regress estimates for married women on a set of model characteristics. Results are reported in Table 5. The main conclusion is that both years and modeling choice matter. An additional year decreases female elasticities by around .02 while using the discrete approach reduces female elasticities by around .27. The overestimation due to the Hausman model is particularly strong for participation elasticities and for EU estimates.[9] We do not report a similar estimation for the US only, given the small number of observations in this case. Yet it transpires from such regression and the graphs discussed above that only the time effect matter in the US.[10] Notice that using desired rather than observed hours inflate hour elasticites, which may reflect constraints on working time. So does modeling joint decision in couples rather than estimating married women's

[7] Figure 3 (bottom left quadrant) is actually obtained when taking out the outlier estimate of participation elasticities by García and Suárez (2003) for Spain. When including it, the pattern is U-shaped rather than monotonically decreasing (using quadratic trend curves instead of standard trend lines).

[8] This calls for further research comparing methods over the long run, i.e. a replication of Heim (2007) and Blau and Kahn (2007) using the discrete model.

[9] This is illustrated by example of France given above. Note that we have checked that results are not driven by particular outliers like the implausibly high values for some studies, as mentionned above.

[10] Note that recent period estimates are very similar whether they stem from grouped estimations (Devereux, 2004), natural experiments (Eissa and Hoynes, 2004) or structural models (Heim, 2009). As already discussed in footnote 8, additional estimates using discrete choice models, and for many years including the older period, should be produced.

labor supply separately (a male chauvinistic model). None of these two effects is statistically significant, however.

4 Conclusion

In this paper, we provide an extensive survey of studies estimating labor supply elasticities for Western Europe and the US. Beyond confirming conventional wisdom, we derive original results concerning the variation in labor supply estimates across studies. While Bargain et al. (2013) show that international heterogeneity in work preference matters but is small, we investigate here the role of two major methodological differences across studies, i.e time period and estimation methods, that explain differences. Large elasticities are mainly driven by studies estimated in the 1980s and relying on the Hausman approach. More recent estimates based on structural discrete-choice models with tax-benefit simulations show smaller estimates and relatively more similarity across studies. More points of observations are however needed to disentangle the two factors, i.e. possibly larger elasticities in the 1970s/1980s related to lower female participation on the one hand and overestimations due to the Hausman model on the other. Our meta-analysis nonetheless allows us to conclude the following: While we confirm that elasticities decline over time in the US, there is some evidence that both time effects and modeling choices affect estimates in Europe. Time effects, and notably declining elasticities over time in the US, are to some extent consistent with cross-country comparisons, i.e. the fact that countries with more firmly established female participation, elasticities are smaller (Bargain et al., 2013).

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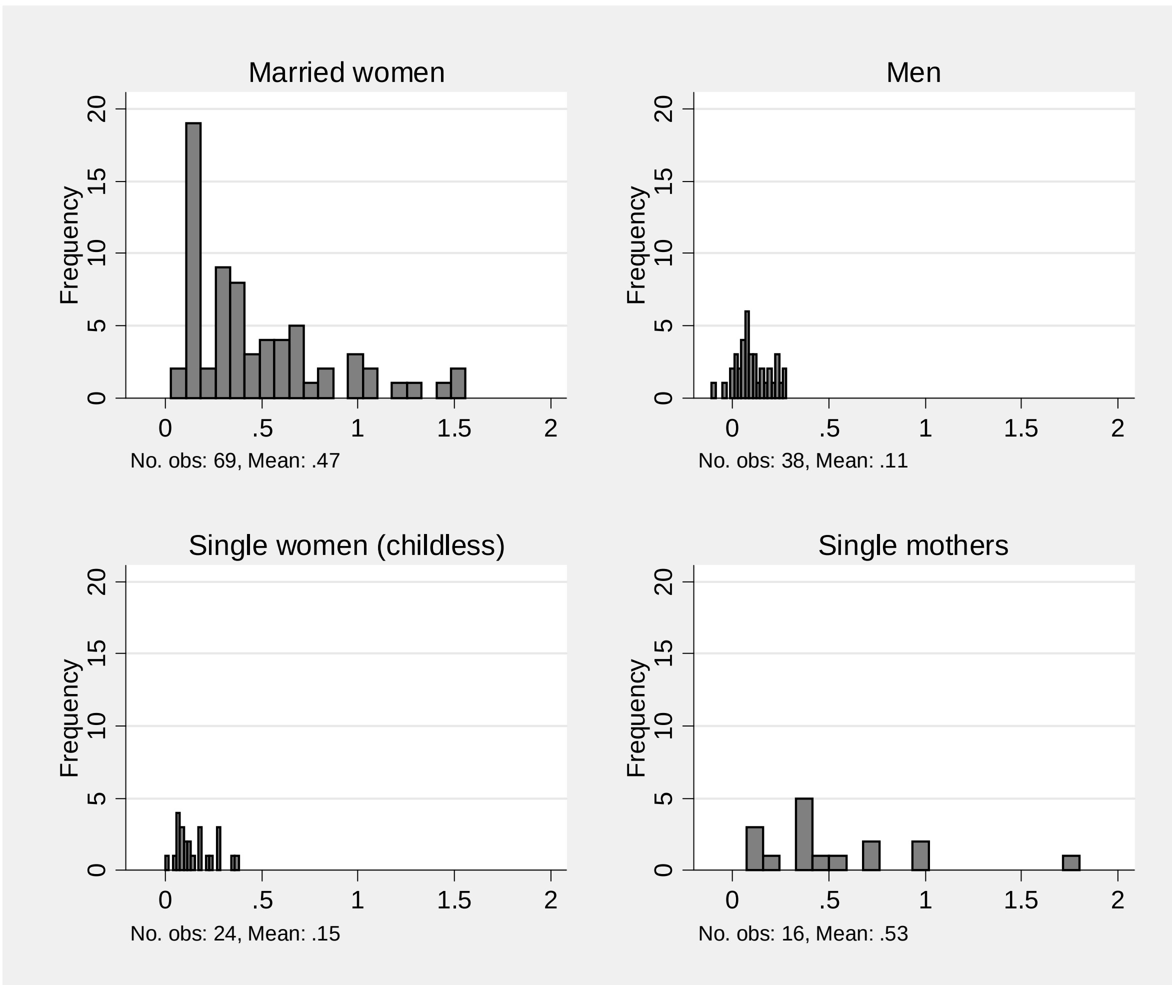

Figure 1: Distribution of Wage-Elasticities by Demographic Group

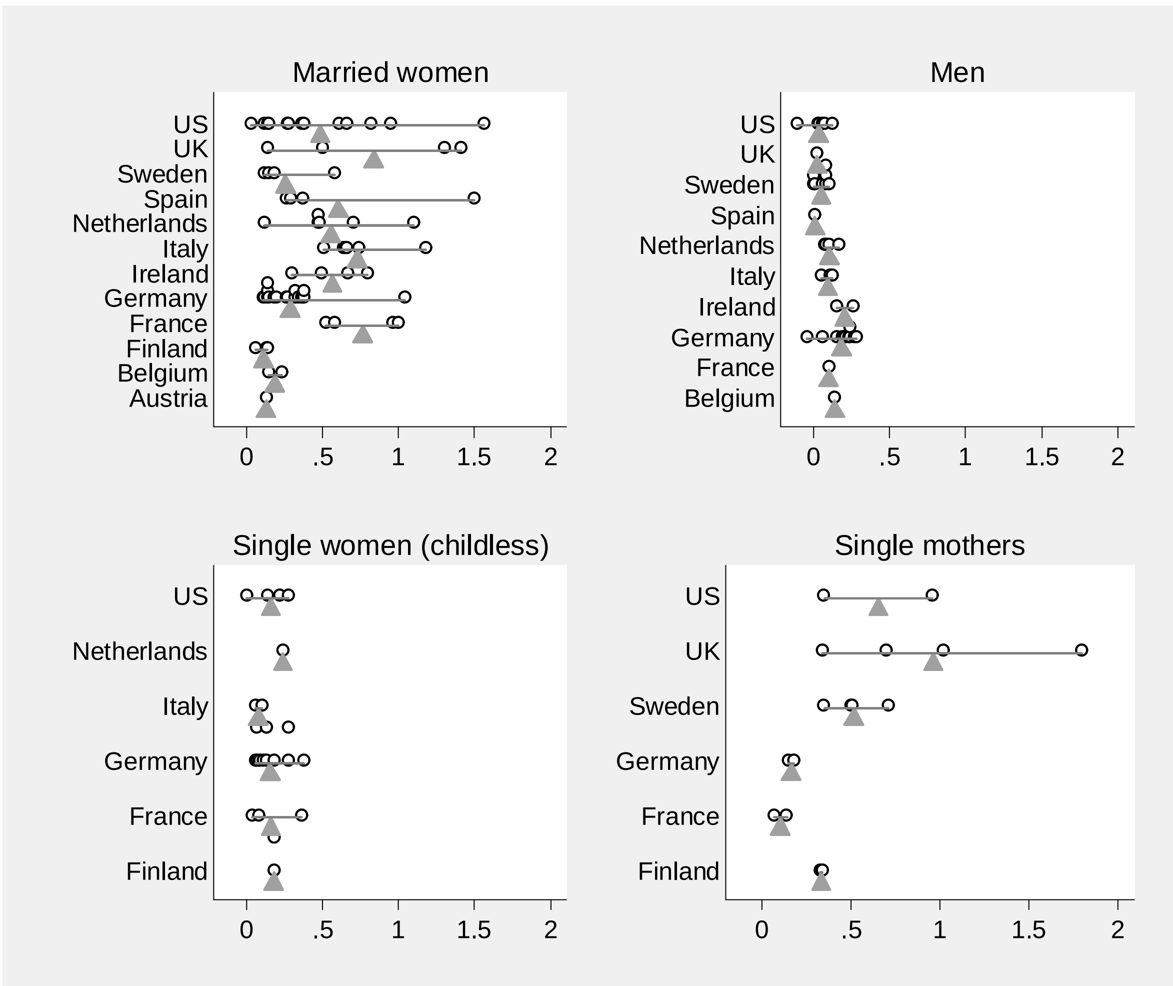

Figure 2: Distribution of Wage-Elasticities by Demographic Group and Country

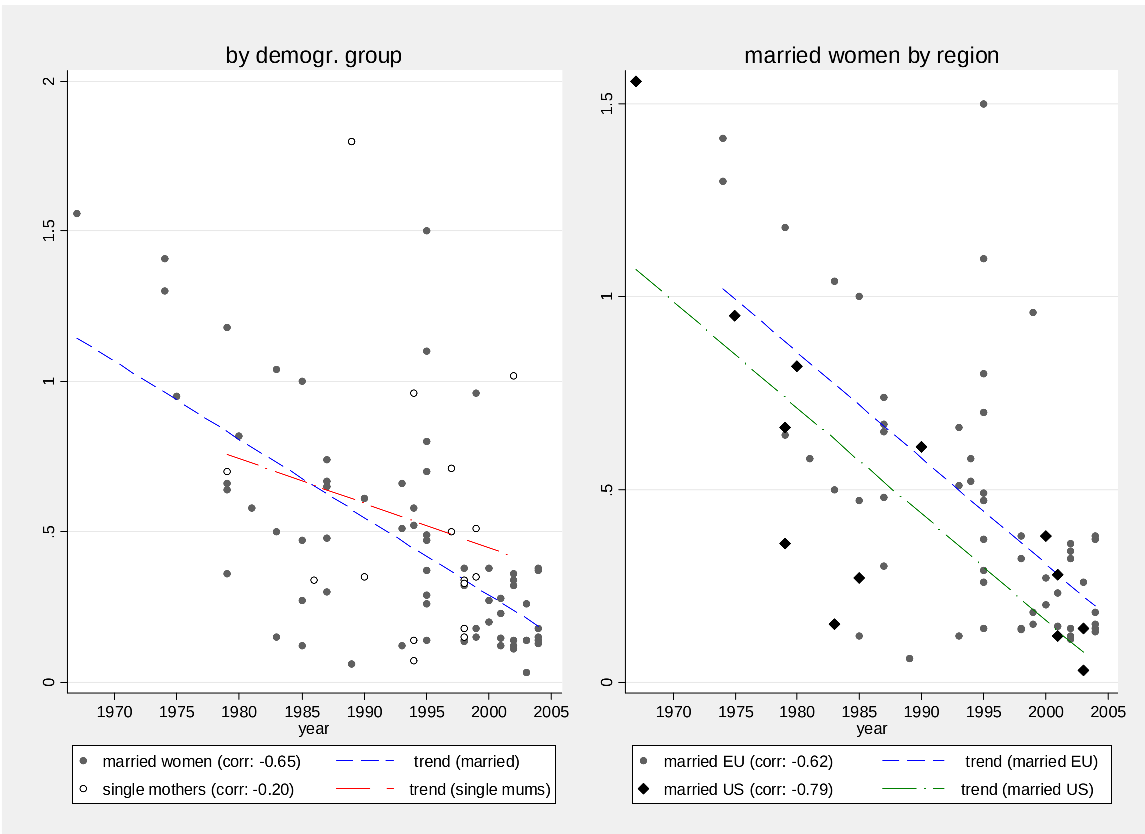

Figure 3: Time Trend in Wage-Elasticities of Married Women and Single Mothers

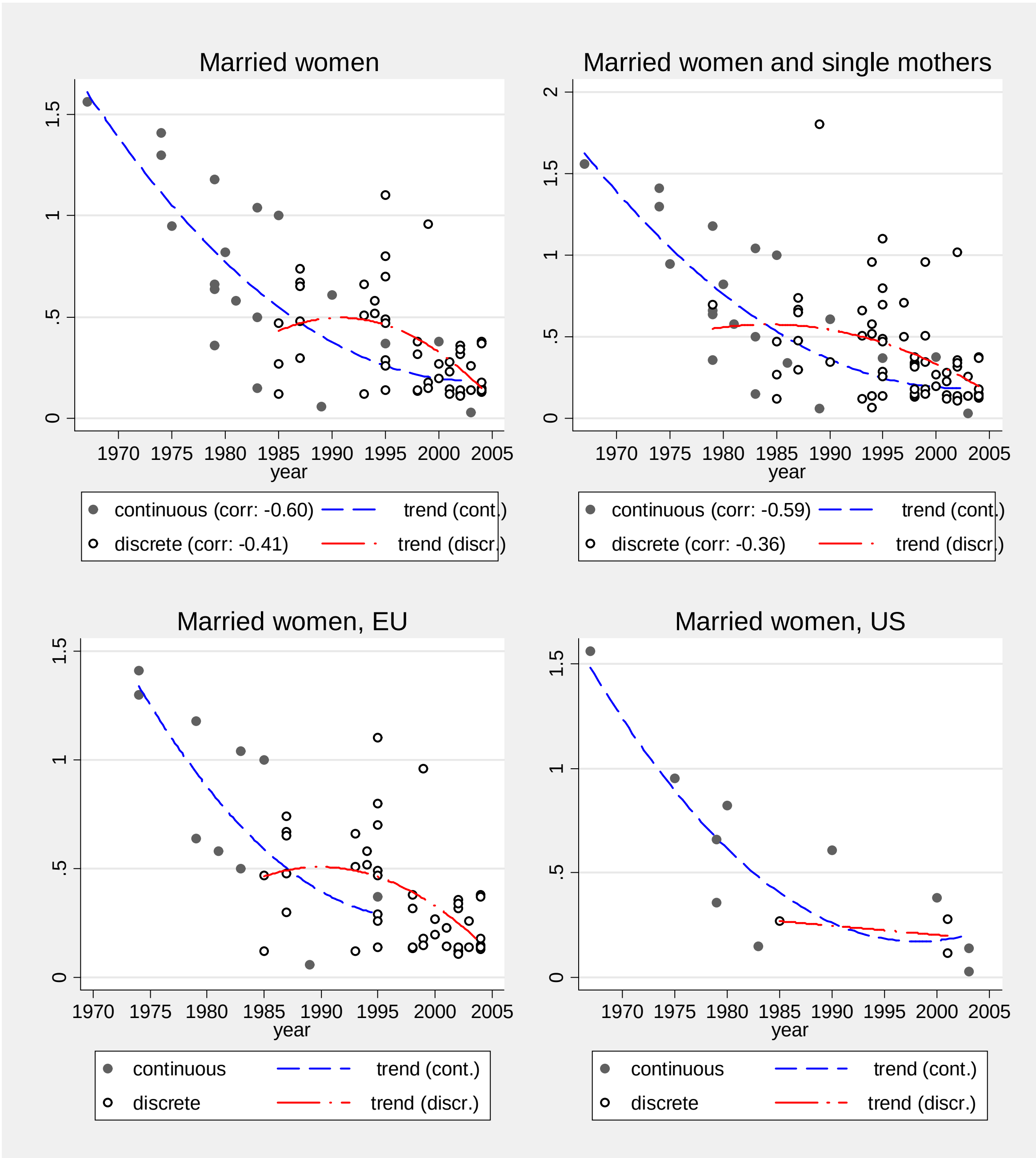

Figure 4: Time Trend in Wage-Elasticities: by Broad Estimation Methods

Table 1: Labor Supply Elasticities in Europe: Couples

Country	Authors	Data selection	Model	Specification	Tax-benefit	Female wage elast.		Male wage elast.		Incom
						hours	particip.	hours	particip.	female
Austria	Dearing et al. (2007)	SILC (2004), at least 1 child aged <10	D	QU; M	ITABENA		[.07, .19] @			
Belgium	Orsini (2007, 2012)	Panel Survey of Belgian Households (2001), working age	D	QU and GU + PTD; J	MODETE	[.16, .31]	[.10, .19]	[.10, .18]	[.08, .15]	
Finland	Kuismainen (1997)	LFS (1989), survey & tax register; 25-60	C	SL, R	PL	[0, .06]				[.11, .27]
	Bargain & Orsini (2006)	IDS (1998), working age, men all employed	D	QU + FC; M	EUROMOD	[.10, .18]	[.10, .17]*			
France	Bourguignon & Magnac (1990)	LFS (1985), couples aged 18-60	C/T	LL + R; M or J	PL, D	1 (.05 with FC)		.10		-.03 (-.02 with FC)
	Laroque & Salanie (2002)	matched LFS-Tax returns (1999), women aged 25-49	D	joint particip. & wage; unempl. & min. wage	own calc.		(.96)			/ -.11*
	Choné, Le Blanc & Robert-Bobée (2003)	matched LFS-Tax returns (1997), working age, children aged <6	D	QU, joint wage & CC; min. wage	own calc.	1.05	[.8, .9] @			-.19 / -.18*
	Bargain & Orsini (2006)	HBS (1994/5), working age women, men all employed	D	QU + FC; M	EUROMOD	[.52, .65]	[.46, .58]*			
	Donni & Moreau (2007)	HBS (2001), aged 20-60, all employed, no children aged<3	C	QL; s-conditional collective LS	no taxation	[.24, .59]				[-.35, -.06]
Germany	Kaiser et al. (1992)	SOEP (1983), working age	C	LL	C, NC, D	1.04		-.04		-.18
	Bonin, Kempe & Schneider (2002)	SOEP (2000), working age, W & E	D	TL + PTD; J	IZAmod	.27	.20	.21	.19	.15 / .09
	Steiner & Wrohlich (2004)	SOEP (2002), working age, W & E	D	TU + PTD; J	STSM	[.16, .55] @	[.07, .21] @	[.11, .38] @	[.07, .23] @	
	Haan & Steiner (2004)	SOEP (2002), working age, W & E, one- or two-earner couples	D	TU + PTD; J	STSM	[.08, .56]	[.04, .20]	[.08, .46]	[.07, .26]	
	Bargain & Orsini (2006)	SOEP (1998), working age, men all employed, W & E	D	QU + FC; M	EUROMOD	[.31, .45]	[.27, .38]*			
	Clauss & Schnabel (2006)	SOEP (2004/5), couples aged 20-65	D	TU; J	STSM	.37	.14	.24	.16	
	Wrohlich (2006)	SOEP (2002), working age, W & E	D	TU; J; CC	STSM	[.14, .53] @	[.06, .16] @			
	Dearing et al. (2007)	SOEP (2004), at least 1 child aged <10, W	D	QU; M	STSM		[.13, .24] @			
	Bargain et al. (2009)	SOEP (2003), working age, potential one- or two-earner	D/H	QU + PTD, R; J	STSM	[.19, .34]	[.08, .20]	[.05, .08]	[.04, .13]	
	Fuest et al. (2008)	SOEP (2004), working age, W & E, potential one- or two-earner	D	TU+PTD;J	FiFoSiM	0.38	0.15	0.20	0.14	

Data: Income Distribution Survey (IDS), Household Budget Survey (HBS), Socio Economic Panel (SOEP), Family Expenditure Survey (FES), Labor Force Survey (LFS), EU Statistics on Income and Livi Conditions (SILC). For Germany: West (W), East (E).

Model: C = continuous labor supply (Hausman 1981 type); T = tobit model; D = discrete-choice model (van Soest 1995 type); A = estimation of joint distributions of wage and hours (sets of hour-wage o vary across individuals); H = double hurdle model (labor supply and risk of unemployment).

Specification: for Hausman model, labor supply is either linear (LL), quadratic (QL) or semi-log (SL); in discrete-choice models, utility is either quadratic (QU), translog (TU) or generalized Stone-Geary (C preferences are sometimes accounted for (R) as well additional flexibility, either through fixed costs (FC) or part-time dummies (PTD). Models are male-chauvinistic (M) or account for joint decision in cou Welfare programme participation (W). Childcare costs (CC).

Tax-benefit: Hausman model often accounts for piecewise-linear budget set (PL) or more generally convex set (C); nonconvexities are sometimes accounted for (NC); differentiability of the budget functio used (D); with discrete choice models, complete tax-benefit systems are simulated and we indicate the name of the microsimulation model when it is known.

Elasticities: brackets indicate the range of values for all specifications (or the confidence interval when available). '@ indicates that the range also includes values for different age and number of children. F participation elasticities, corresponding to the increase in employment rate in % points, except when indicated by * (in that case, % increase in employment rate).

Table 2: Labor Supply Elasticities in Europe: Couples (cont.)

Country	Author	Data selection	Model	Specification	Tax-benefit	Female wage elast.		Male wage elast.		Income elast.	
						hours	particip.	hours	particip.	female	male
Ireland	Callan & van Soest (1996)	IDS (1987), desired hours	D/H	TU + FC, R; J	SWITCH	[.50, .85]	.31 /.20*	[.10, .20]			
	Callan, van Soest & Walsh (2009)	Living in Ireland Survey (1995), desired hours	D	TU + FC, R; J	SWITCH	[.71, .90]	.49	[.21, .31]	.20 /.21*		
Italy	Colombino & Del Boca (1990)	Turin Survey of Couples (1979), working age	C	LL	PL	1.18	.64			.52	
	Aaberge et al. (1999)	Survey of Income and Wealth (1987), aged 20-70	A	non-linear hours, exog. wage and unearned inc.	own calc.	.74	.65	.053	.046	/ -.014	/ -.003
	Aaberge et al. (2002, 04)	Survey of Income and Wealth (1993)	A	GU; J	own calc.	.66	.51	.12	.02		
Netherlands	van Soest et al. (1990)	Labor mobility survey (1985), working age	C/D	LL, R; discrete wage-hours combinations	PL	[.35, .59]	.12	[.15, .19]		-.23	-.01
	van Soest (1995)	SOEP (1987)	D	TU + PTD, R; J	own calc.	[.42, .54]	-	[.05, .09]	-	.008	-.03
	van Soest & Das (2001)	SOEP (1995), aged 16-64, desired hours	D	TU + FC, R; J	own calc.	[.67, .74]	-	[.07, .10]	-		
	van Soest et al. (2002)	Dutch SOEP (1995), aged 16-64, desired hours	D	QU (+ more flexible) + FC, R; simult. wage estimation, J	own calc.	[.83, 1.36]	[.35, .58]*				
Spain	García and Suárez (2003)	ECHP (1994-95), aged 16-65, obs. and desired hours	C	LL	taxes	.37	1.51*			-.06	
	Fernández-Val (2003)	ECHP (1994-99), aged<65 and in work	C	unitary/collective model	no taxation	.31					
	Crespo (2006)	ECHP (1994-99), aged<65 and in work	C	QL, unitary/collective	no taxation	.14		.01			
	Labeaga, Oliver & Spadaro (2008)	ECHP (1995), working age	D	QU + FC; J	GLAD-HISPANIA	.29	.26	.01	.11		
Sweden	Blomquist (1983)	Level of Living Survey (1974), all employed, aged 25-55	C	LL, R	PL			.008			-.03
	Flood & MaCurdy (1992)	Household Market-Nonmarket Survey (1983), all employed, 25-65	C	LL and SL, R	PL, D			[.-.25, .21]			[.-.01, .04]
	Blomquist & Hansson-Brusewitz (1990)	Level of Living Survey (1981), all employed, aged 25-55	C	LL and QL, R	PL, C and NC	[.38, .77]		[.08, .13]		[-.24, -.03]	
	Blomquist & Newey (2002)	Level of Living Survey (1973, 80, 90), all employed, aged 18-60	C	non-parametric labor supply	PL			[.04, .12}			-.02
	Flood, Hansen & Wahlberg (2004)	Household Income Survey (1993), aged 18-64	D	TU, R; stigma of W	own calc.	.12		0		-0.017	-0.003
	Brink et al. (2007)	Longitudinal Individual Data, Income Distribution Survey, 1999	D	TU, R	FASIT	.18	.15	.06	0		
UK	Arellano & Meghir (1992)	British FES and LFS (1983), aged 20-59, with pre-school children (upper bound for all children)	C	SL + FC, search costs, endogenous wage and unearned income (IV)	PL	[.29, .71]	-			[-.13, -.40]	
	Arrufat & Zabalza (1986)	British General Household Survey (1974), aged <60	C	CES utility based labor supply, R	PL	[.62 - 2.03]	1.41			-.2 / -.14	
	Blundell & Walker (1986)	FES (1980), all employed, aged 18-59	C	Gorman polar form and translog hours, R	PL			.024			-.287
	Blundell, Ham & Meghir (1987)	FES (1981), aged 16-60	T/H	non-linear labor supply, unemployment risk	own calc.		[.04, .08]				
	Blundell, Duncan & Meghir (1998)	FES (1978-1992), 20-50, young children (lower bound if no child)	C	generalized LES, R	PL	[.13, .37] @	-			[-.19, 0] @	
	Blundell et al. (2000)	Family Resources Survey (1994-96)	D	QU + FC, R, W	TAXBEN	[.11 - .17]					

Note: see previous table. For Spain, several additional references are cited in García and Suárez (2003) which point to similar elasticities as in the basic model in this study.

Table 3: Labor Supply Elasticities in Europe: Single Individuals

Country	Author	Data selection	Model	Specification	Tax-benefit	wage elasticites		income elast.
						hours	particip.	
Finland	Bargain & Orsini (2006)	IDS (1998), SW, SP	D	QU + FC	EUROMOD	[.18, .34]	[.18, .33]	
France	Bargain & Orsini (2006)	HBS (1994/5), aged 25-49, SW, SP	D	QU + FC	EUROMOD	[.08, .14]	[.04, .07]	
	Laroque & Salanie (2002)	LFS-Tax return matched dataset (1999), women aged 25-49, no civil servants, SW	D	participation (and full/part-time) model, simultaneous wage and labor supply estimation, probability of unemployment, min. wage	own calc.		{.36}	
Germany	Bargain & Orsini (2006)	SOEP (1998), SW, SP	D	QU + FC	EUROMOD	[.09, .18]	[.08, .15]	
	Steiner & Wrohlich (2004)	SOEP (2003), SW	D	TU + PTD	STSM	[.20, .36]	[.05, .09]	
	Haan & Steiner (2004)	SOEP (2002), SW	D	TU + PTD	STSM	[.02, .24]	[.01, .10]	
		SM				[.08, .31]	[.04, .28]	
	Clauss & Schnabel (2006)	SOEP (2004/5), aged 20-65, SW	D	TU + PTD	STSM	.38	.18	
		SM				.23	.17	
	Bargain et al. (2009)	SOEP (2003), working age, SW	D/H	QU + PTD; involuntary unemployment	STSM	[.06, .16]	[.04, .10]	
		SM				[.10, .20]	[.05, .12]	
	Fuest et al. (2008)	SOEP (2004), working age, SW	D	TU + PTD	FiFoSiM	0.28	0.13	
		SM				0.28	0.17	
Italy	Aaberge et al. (2002)	Survey on Household Income and Wealth (1993), SW	A	GU	own calc.	.10	.06	
		SM				.11	.08	
Netherlands	Euwals & Van Soest (1999)	Dutch SOEP (1988), actual and desired hours, SW	D	TU + FC, R	own calc.	[.03, .45]		
		SM				[.03, .18]		
Sweden	Andren (2003)	HINK (1997-98), SP	D	QU + FC; simulat. with W and CC	own calc.	[.55, .87]	.50	-0.1
	Brink et al. (2007)	Longitudinal Individual Data, IDS, 1999, SP	D	TU, R	FASIT	.51	.35	
UK	Walker (1990)	FES (1979-84), SP	D	participation model	benefits only		.70	
	Ermisch & Wright (1991)	General household survey (1973-82), SP	D	participation model, demand-side controls	simplified system		1.7	
	Jenkins (1992)	Lone parents survey (1989), SP	D+H	two positive hour choices, unemployment risk, FC	benefits only		1.8	
	Blundell, Duncan & Meghir (1992)	FES (1981-1986), SP	C	marginal rate of substitution function, endogenous wage and income	taxation only		.34	
	Brewer et al. (2006)	FES (1995-2002), aged <60, SP	D	QU + FC, joint with W and CC, R	TAXBEN		1.02	

Data & Selection: Income Distribution Survey (IDS), Household Budget Survey (HBS), Socio Economic Panel (SOEP), Family Expenditure Survey (FES), Labor Force Survey (LFS); Selection: single women (SW), single men (SM), single parents/mothers (SP)

Model: C = continuous LS (Hausman 1981 type); T = tobit model; D = discrete model (van Soest, 1995 type); A = estimation of joint distributions of wage and hours (sets of hour-wage opportunities vary across individuals); H = double hurdle model (labor supply and risk of unemployment).

Specification: for Hausman model, labor supply is either linear (LL), quadratic (QL) or semi-log (SL); in discrete-choice models, utility is either quadratic (QU), translog (TU) or generalized Stone-Geary (GU); random preferences (R); fixed costs (FC); welfare participation (W); childcare costs (CC)

Tax-benefit: Hausman model often accounts for piecewise-linear budget set (PL) or more generally convex set (C); nonconvexities are sometimes accounted for (NC); differentiability of the budget function can be used (D); with discrete choice models, complete tax-benefit systems are simulated and we indicate the name of the microsimulation model when it is known.

Elasticities: brackets indicate the range obtained in function of the specification at use, or the confidence interval when available. Particip. = participation elasticities, corresponding to the increase in employment rate in percentage points.

Table 4: Labor Supply Elasticities for the US

Authors	Data selection	Model	Specification	Female wage elast.		Male wage elast.		Income elast.	
				hours	particip.	hours	particip.	female	male
Cogan (1981)	US National Longitudinal Study of Mature Women 1967, married women aged 30-35	C	SL; reservation hours to account for FC; no tax-benefit	[.86 , 2.40]				[.16 , .66]	
Hausman (1981)	PSID 1975, married women	C	LL, PL (C and NC: FC)	[.90 , 1.00]				[-.13 , -.12]	
Triest (1990)	PSID 1983, married women, aged 25-55	C	LL; C and PL; taxes and benefits	[.03 , .28]				[-.15 , -.19]	
MaCurdy, Green & Paarsch (1990)	PSID 1975: married men, aged 25-55	C	LL; PL and D (reconvexified) budget set; taxes			[-.24, .03]		-.01	
Dickert, Houser and Scholz (1995)	SIPP 1990, single mothers, no assets	D	joint program and labor force participation		.35				
Pencavel (1998)	CPS 1975-94, women aged 25-60	C	Log-L; no tax-benefit		[.77,.1.80]				
Hoynes (1996)	SIPP panel, 1984, married men and women with children	D	Stone-Geary; stigma from AFDC; tax-benefit system; FC					- .46	- .12
Keane and Moffitt (1998)	1994 SIPP, single mothers, no assets	D	joint labor supply and welfare program participation; benefits but no tax		.96				
Pencavel (2002)	CPS 1999, married and single men	C	LL; no tax-benefit			[.12,.25]			
Devereux (2003)	Census and PSID, all men	C	Log-L, no tax-benefit			[–.022, .017]	[–.061, .001]		
Devereux (2004)	PUMS 1980,1990, married couples (participating men)	C	Log-L, no tax-benefit	[.17,.38]		[.00,.07]			
Eissa & Hoynes (2004)	CPS 1985 to 1997, less educated married couples with children	D	Participation Probit, joint estimation		0.27		.03	-.039	-.007
Blau & Kahn (2007)	CPS 1980, married men and women age 25-54	C	Log-L	[.77,.88]		[.01,.07]		.004	.001
	CPS 1990	C	Log-L	[.58,.64]		[.10,.14]		.002	.002
	CPS 2000	C	Log-L	[.36,.41]		[.04,.10]		.001	.002
Heim (2009)	PSID 2001, couples		quadratic utility with continuous labor supply, J, FC, R	[.24,.33]	[.07,.18]	[.04,.07]	[.00,.003]	[-.007, -.006]	[-.0007,-.0004}
Bishop et al. (2009)	CPS, 1979-2003, sing. women		SL, participation, some account for tax	.14 (1979) to -.03 (2003)	.28 (1979) to .22 (2003)			-.014 (1979) to -.019 (2003)	
Heim (2007)	CPS, 1979-2003, married women		SL, participation, some account for tax	.36 (1979) to .14 (2003)	.66 (1979) to .03 (2003)			-.05 (1979) to -.015 (2003)	

Data: Current Population Survey (CPS), Panel Study on Income Dynamics (PSID), Public Use Microdata Sample (PUMS), Survey of Income and Program Participation (SIPP)

Model: C= continuous labor supply (Hausman 1981 type); D= discrete-choice model (often a simple participation probit)

Specification: Hausman labor supply is either linear (LL), log-linear (Log-L) or semi-log (SL); random preferences are sometimes accounted for (R) as well as fixed costs (FC). Models sometimes account for piecewise-linear budget set (PL) or more generally convex set (C) or nonconvexities (NC), and differentiable budget constraint (D).

Elasticities: brackets indicate ranges of values over different specifications, or reported confidence intervals. Participation elasticities ("particip"): increase in employ. rate in % points.

Table 5: Meta Regression of Married Women's Wage-Elasticities

Model	All elasticities	Participation elasticities	Hour elasticities	Without the US
year	-0.016 **	-0.021 **	-0.023 ***	-0.013
	(.007)	(.010)	(.008)	(.008)
discrete model	-0.274 **	-0.675 ***	-0.096	-0.347 **
	(.129)	(.234)	(.138)	(.158)
desired hours	0.151	-0.014	0.178	0.183
	(.116)	(.180)	(.129)	(.127)
joint decision	0.022	-0.006	0.139	-0.009
	(.092)	(.133)	(.111)	(.102)
fixed cost #	0.032	-0.058	0.107	0.042
	(.082)	(.120)	(.094)	(.086)
US	-0.202 *	-0.454 **	0.056	
	(.126)	(.194)	(.147)	
constant	0.771 ***	1.289 ***	0.601 ***	0.809 ***
	(.128)	(.265)	(.121)	(.145)
Nb of observations	56	26	30	49
R2	0.20	0.31	0.33	0.19

for discrete models