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Labor Supply Elasticities: Can Micro Be Misleading for Macro?

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Abstract - In this paper we compare “micro” and “macro” labor supply elasticities in a MaCurdy-type equation. Using PSID data, we obtain the micro elasticity from standard panel techniques, and the macro elasticity from the time series generated by aggregating individuals every year. This procedure relies on the exact aggregation of first-order conditions in a life-cycle model with home production. We find an individual elasticity of about 0.1, a low value in line with mainstream microeconometric studies, and an aggregate elasticity of about 1, a much larger value often assumed in calibration studies. This discrepancy is not due to aggregation bias: it is due to the fact that individual and total hours are different variables, with the extensive margin that empirically dominates. A broader implication of our result is that micro evidence is not always appropriate for calibrating an aggregate model economy.

JEL Classification Codes: E13, E32, J22

Keywords: elasticity of labor supply, aggregation, calibration.

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1 Introduction

The intertemporal substitution of leisure is crucial for explaining business cycles in modern macroeconomics. When stating how the benchmark RBC model should be calibrated, Prescott (1986) suggested to restrict the stochastic growth model on the basis of the available micro-econometric evidence:

> "A fundamental thesis of this line of inquiry is that the measures obtained from aggregate series and those from individual panel data must be consistent. After all, the former are just the aggregates of the latter." (p. 14).

To reproduce all business cycles facts, however, the benchmark RBC model requires a much larger elasticity than typically estimated in micro studies. Microeconomic studies based on both cross-sectional and panel data estimate the intensive margin of labor adjustment and typically report a small real-wage elasticity [e.g. Pencavel (1986), Killingsworth and Heckman (1986), MaCurdy (1981), Altonji (1986)], ranging from about 0 to about 0.2 for men and from about 0 to about 1 for married women [Blundell and MaCurdy (1999)]. This difficulty is noticed by several authors such as Heckman (1993), Browning, Hansen and Heckman (1999) and by Prescott himself (2006).

In the meantime, the macroeconomic evidence is far less numerous, is generally mixed and typically includes the extensive margin of labor adjustment which empirically dominates the intensive one [Cho and Cooley (1994)]. In their seminal paper, Lucas and Rapping (1969) find that, for the US economy (1930-1965), total hours are strongly real-wage elastic in the short-run (1.4). Among the others, Hall (1980) finds an intertemporal elasticity of substitution which is about 0.5, while Mankiw, Rothemberg and Summers (1985) reject the intertemporal substitution hypothesis by estimating the intensive margin only, rather than the most appropriate aggregate hours changes [Heckman, (1993)]. The importance of including changes in employment is stressed by Alogoskoufis (1987), who shows that when applied to aggregate employment the intertemporal substitution hypothesis is not rejected by the US data.

The necessity of reconciling the relatively high aggregate elasticity assumed in calibration studies with the low elasticity estimated in microeconometric studies brought about a number of different orientations. In some cases [e.g. Summers (1986), Mankiw (1989)] the whole relevance of the RBC model was denied. A more constructive orientation explored several variants of the benchmark RBC model [i.e. Prescott (1986)] in order to better accommodate the data. A precursor is the seminal work of Kydland and Prescott (1982) based on non-separability of leisure at different points in time. This was followed by the lottery [Rogerson (1988)] and the indivisible labor model [Hansen (1985)] where people either work a fixed or a zero amount of hours. Among the other relevant extensions, the introduction of preference shocks [Bencivenga (1992)], government consumption [Christiano and Eichenbaum (1992)], home production [Benhabib, Rogerson and Wright (1991)], and taxation in general equilibrium [Baxter and King (1993); McGrattan (1994)] are all noteworthy efforts to add realism and policy focus to the benchmark RBC model.

More recent studies that generate a wedge between individual and aggregate labor elasticities have focused on heterogeneous reservation wages [Chang and Kim (2005) and (2006)], omission of variables such as wealth [Ziliak and Kniesner (1999)] and liquidity constraints [Domeji and Floden (2006)], human capital accumulation [Imai and Keane (2004)]. Needless to say, this list is incomplete.

The research question we address in this paper is whether and to what extent a small individual elasticity of labor supply is consistent with a large aggregate one. This question cannot be addressed as an aggregation bias issue [Theil (1954)] – i.e. as a situation where the aggregate parameters differ from the averages of the corresponding micro parameters – since micro and macro estimates of the elasticity of labor supply do not refer to the same variable. Typically, micro estimates deal with individual hours of work per unit of time (intensive margin), while macro estimates deal with total hours of work, i.e. the product of the intensive margin and the employment rate (extensive margin). In fact we exploit this difference to compare in a fully consistent way "micro" and "macro" labor supply elasticities referring to the same units in the same dataset (PSID). This comparison obviously requires that there is no difference in the data sources as well in the specification and estimation method, including the choice of instruments.

This same issue has been addressed in a few calibration models. Chang and Kim (2005; 2006) combine the indivisible labor assumption and the heterogeneity of reservation wages in an incomplete markets model. Assuming an individual elasticity of 0.4 they find an aggregate elasticity of about 1, although this number is no longer related to the intertemporal substitution of leisure. Conversely, Prescott, Rogerson and Wallenius (2006) assume a nonlinear mapping between hours of work and labor services that generates virtually unrelated aggregate and micro labor elasticities. Finally, Rogerson and Wallenius (2008) extend Prescott, Rogerson and Wallenius (2006) by including age, which also affects productivity and labor supply decisions over time.

We are not aware of any other evidence addressing this issue by *estimating* micro and a macro labor supply from the same data, chosen to allow for a consistent and fully comparable aggregation procedure. Our empirical route relies on aggregation of individual first-order conditions – which allows us to solve the aggregation problem [Blundell and Stoker (2005)] – in a life-cycle model where the extensive margin matters because people, in equilibrium, are engaged in either market or home production. Our model generates a mapping between individual and aggregate elasticities: the latter is larger because people move into and out the pool of workers in response to productivity shocks.

We use all of the annual waves (1968-1997) of the Panel Study of Income Dynamics (PSID) to estimate the individual Frisch elasticity via a long-enough panel to be compared with the corresponding time series resulting from the exact aggregation of individual units each year. This procedure – which is not costless for reasons we will discuss later – offers an important advantage: the macro dataset is based on exactly the same units of observation that compose the micro dataset, and allows us to employ "isomorphic" micro and macro regressions.

Our panel estimate, which by and large follows the pioneering study of MaCurdy (1981), delivers a Frisch elasticity of about 0.1, a small value that is consistent with benchmark micro estimates. The aggregate time series delivers instead a Frisch elasticity of about 1. We decompose such aggregate elasticity into the contribution of adjustment of hours per worker and of employment, finding that the latter accounts for most of the difference between the two elasticities we estimate.

The paper is organized as follows. In Section 2 we discuss the relevance of disentangling between the intensive and the extensive labor margin. Section 3 illustrates the theoretical model. Section 4 presents our results and Sections 5 concludes. A detailed Data Appendix is available in Section 6.

2 Intensive vs. Extensive Margin

The indivisible labor case [Rogerson (1988); Hansen (1985)], where individuals either work a fixed amount of hours or do not work at all, accommodates in an extreme way the well-known evidence that labor adjustment on the extensive margin dwarfs adjustment on the intensive margin [Cho and Cooley (1994)]. Like in Hansen (1985), if we denote by n_t the employment stock and by $\overline{h}_t$ the average supply of hours, then aggregate labor is $H_t \equiv n_t\overline{h}_t$. By taking logs, the variance of labor input can be decomposed as follows:

$$\text{var}\left(\ln H_t\right) = \text{var}\left(\ln n_t\right) + \text{var}\left(\ln \overline{h}_t\right) + 2\text{cov}\left(\ln n_t, \ln \overline{h}_t\right). \tag{1}$$

The share of the total variation that is due to n_t provides a measure of the importance of the extensive margin. For quarterly US data ranging from 1955 to 1984, Hansen (1985) finds that employment changes account for 55% of the total hours deviations from the HP trend, while the hours per worker deviations account for only 20%. This pattern is observed in several countries: in HP-filtered, quarterly manufacturing data (1960-1989), Fiorito and Kollintzas (1994) found that the volatility of employment deviations from the smooth trend always exceeds the corresponding volatility in hours per worker: by a factor of about eight in the US, about four in Canada and West Germany and between two and three in the UK and in Japan, respectively.

The wedge between individual and aggregate elasticities, as conventionally estimated, reflects such a primacy of the extensive margin. This is easy to see in a regression framework. Henceforth, we use lower case for individual variables and upper case for the corresponding aggregate quantity. Denote by ε and $\mathcal{E}$ the micro and macro Frisch elasticities of labor supply, respectively, by w_t and W_t the individual and aggregate wage rates at time t, respectively, and by h_t the individual hours worked. Consider the following MaCurdy (1981) regressions, which provide the benchmark for estimating a Frisch elasticity:

$$\text{individual} \quad : \quad \Delta \ln h_t = \textit{const.} + \epsilon \Delta \ln w_t + e_t, \tag{2}$$

$$\text{aggregate} \quad : \quad \Delta \ln H_t = \textit{Const.} + \mathcal{E} \Delta \ln W_t + E_t. \tag{3}$$

The population elasticities are:

$$\epsilon = \frac{\text{cov}\left(\Delta \ln h_t, \Delta \ln w_t\right)}{\text{var}\left(\Delta \ln w_t\right)}, \tag{4}$$

$$\begin{aligned} \mathcal{E} &= \frac{\text{cov}\left(\Delta \ln H_t, \Delta \ln W_t\right)}{\text{var}\left(\Delta \ln W_t\right)} \\ &= \frac{\text{cov}\left(\Delta \ln \overline{h}_t, \Delta \ln W_t\right)}{\text{var}\left(\Delta \ln W_t\right)} + \frac{\text{cov}\left(\Delta \ln n_t, \Delta \ln W_t\right)}{\text{var}\left(\Delta \ln W_t\right)}. \end{aligned} \tag{5}$$

That is, the micro elasticity (4) consists of a single term that captures adjustment on the intensive margin only. The macro elasticity (5) instead, is the sum of two terms representing the intensive and the extensive margins, respectively. Notice that the second term is the covariance between employment and the aggregate wage rate which is positive if we move along a labor supply curve. Intuitively, since we do not expect the "aggregate intensive margin", i.e. the first term in (5), to be less than the individual elasticity, the aggregate elasticity is larger than the individual one. This decomposition illustrates that individual and aggregate elasticities are conceptually different objects [Prescott (2006)], at least because the aggregate variable is the product of two components reflecting different decisions and having, empirically, quite different relevance. To compare these different elasticities in a meaningful way we need a model. This is what we turn to next.

3 The Model

Consider an economy populated by N individuals, indexed by $i = 1, ...N$. There is a composite consumption good, which includes services and which can be produced on the market (c^M) or at home (c^H). In both cases production is characterized by a constant returns to scale technology where labor is the only input. There is no intermediate consumption. Individuals have identical preferences but differ in productivity and endowments.

Denote by θ_{it} and θ_{it}^H two random variables representing individual i's productivity on the market and at home, respectively, at time t, and by h_{it} and h_{it}^H the fraction of hours spent producing on the market or at home. Market and home productivities evolve stochastically following idiosyncratic i.i.d. shocks. Total consumption in the economy is

$$c_t = c_t^M + c_t^H, \tag{6}$$

where

$$c_t^M = \sum_{i=1}^{N} \theta_{it} h_{it}, \tag{7}$$

$$c_t^H = \sum_{i=1}^{N} \theta_{it}^H h_{it}^H. \tag{8}$$

In other words, work on the market and at home are perfect substitutes in production, and the respective outputs are perfect substitutes in consumption.[1] Labor services can be sold on the market at a wage rate w_{it}. Profit maximization implies that $w_{it} = \theta_{it}$ is the market wage of individual i at time t. Individuals are assumed to be forward-looking and markets clear. Due to data limitations, we assume that the tax rate on labor is constant, so that it is immaterial whether the wage rate is pre- or after-tax. Preferences are defined over consumption (c) and leisure (l), and are represented by $\mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t u(c_{it}, l_{it})$, where $u(c_{it}, l_{it})$ is a strictly increasing, twice differentiable, strictly quasi-concave function, and β is the discount factor.

The individual problem is to choose sequences of consumption, $\{c_{it}\}_{t=0}^{\infty}$, labor supply to market, $\{h_{it}\}_{t=0}^{\infty}$, and home production, $\{h_{it}^H\}_{t=0}^{\infty}$, as well as asset holdings, $\{a_{it+1}\}_{t=0}^{\infty}$, that maximize preferences, given the budget and time constraints:

$$\begin{aligned} & \max_{\{c_{it}, h_{it}, h_{it}^H, a_{it+1}\}} \mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t u(c_{it}, l_{it}) \\ \text{subject to} \quad &: \\ c_{it} + a_{it+1} &\leq w_{it} h_{it} + \theta_{it}^H h_{it}^H + (1+r) a_{it} + z_{it}, \\ h_{it} + h_{it}^H + l_{it} &= 1, \end{aligned} \tag{9}$$

where r is the real return on assets – which we assume to be constant in time and across individuals – z_{it} summarizes other exogenous sources of income. The no-Ponzi game condition, $\lim_{T \to \infty} \beta^T \frac{\partial u(c_{iT}, l_{iT})}{\partial c_{iT}} a_{iT+1} = 0$, is also required to hold.

In order to derive a structural equation, we assume that utility is separable in both time and consumption-leisure and is of the CRRA class:

$$u(c_{it}, l_{it}) = \frac{c_{it}^{1-\gamma}}{1-\gamma} - \alpha \frac{(1-l_{it})^{1+\eta}}{1+\eta}, \tag{10}$$

where $\alpha > 0$ reflects the relative preference for leisure.

Denoting by λ_{it} marginal utility of wealth, and by ν_{it}^M and ν_{it}^H the multipliers of the non-negativity constraints on hours spent producing on the market and at home, respectively, the following intratemporal and intertemporal conditions hold at an optimum:

[1] This a special case of the general CES composition of Benhabib, Rogerson and Wright (1991).

$$
\begin{aligned}
c_{it} &: \quad c_{it}^{-\gamma} = \lambda_{it}, & (11)\\
h_{it} &: \quad \alpha\left(h_{it} + h_{it}^{H}\right)^{\eta} = \lambda_{it} w_{it} + \nu_{it}^{M}, & (12)\\
h_{it}^{H} &: \quad \alpha\left(h_{it} + h_{it}^{H}\right)^{\eta} = \lambda_{it}\theta_{it}^{H} + \nu_{it}^{H}, & (13)\\
a_{it+1} &: \quad \lambda_{it} = \beta\left(1+r\right) E_t\left[\lambda_{it+1}\right], & (14)\\
\lambda_{it} &: \quad c_{it} + a_{it+1} = h_{it} w_{it} + \theta_{it}^{H} h_{it}^{H} + \left(1+r_t\right) a_{it} + z_{it}. & (15)
\end{aligned}
$$

Individuals will either supply a positive number of hours to the market or spend a positive number of hours at home but never both, their choice depending at time t on the difference between market productivity θ_{it} and the individual reservation wage ($\widetilde{w}_{it}$), which is equal to θ_{it}^{H}:

$$
h_{it} = \begin{cases} \left(\lambda_{it} w_{it}/\alpha\right)^{1/\eta} & \text{if } \theta_{it} \geq \theta_{it}^{H} \\ 0 & \text{otherwise,} \end{cases} \tag{16}
$$

$$
h_{it}^{H} = \begin{cases} \left(\lambda_{it}\theta_{it}^{H}/\alpha\right)^{1/\eta} & \text{if } \theta_{it} < \theta_{it}^{H} \\ 0 & \text{otherwise.} \end{cases} \tag{17}
$$

For individuals who work on the market, we can rewrite (11), (12) and (14) in logs:

$$
\begin{aligned}
\ln c_{it} &= -\frac{1}{\gamma}\ln\lambda_{it}, & (18)\\
\ln h_{it} &= k + \frac{1}{\eta}\ln\lambda_{it} + \frac{1}{\eta}\ln w_{it}, & (19)\\
\ln\lambda_{it} &= \ln\beta\left(1+r\right) + \ln E_t\left[\lambda_{it+1}\right], & (20)
\end{aligned}
$$

where $k \equiv -\eta^{-1}\ln\alpha$ is a constant. Equation (19) cannot be estimated, since we do not observe λ_{it}. If it were estimated as it is, i.e. including $\eta^{-1}\ln\lambda_{it}$ in the error term, one would be estimating the Marshall, rather than Frisch, elasticity. Following Blundell and MaCurdy (1999) we deal with this issue as follows. Define a one-step-ahead forecasting error in marginal utility of wealth as:

$$
\varepsilon_{it} = \ln\lambda_{it} - E_{t-1}\left[\ln\lambda_{it}\right]. \tag{21}
$$

Equations (20) and (21) allow to characterize the implicit stochastic process for λ_i:[2]

$$\ln \lambda_{it} = -\ln \beta \left(1+r\right) + \ln \lambda_{it-1} + \upsilon_{it}, \tag{22}$$

where $\upsilon_{it} \equiv \varepsilon_{it} - \ln E_{t-1}\left[\exp\left(\varepsilon_{it}\right)\right]$. Next, denoting by $\Delta X_t \equiv X_t - X_{t-1}$ the first difference of any variable X_t, we can rewrite (19) accordingly:

$$\Delta \ln h_{it} = \frac{1}{\eta}\Delta \ln \lambda_{it} + \frac{1}{\eta}\Delta \ln w_{it}. \tag{23}$$

Substituting (22) into this equation and defining $e_{it} \equiv \eta^{-1}\upsilon_{it}$, we obtain:

$$\Delta \ln h_{it} = const. + \frac{1}{\eta}\Delta \ln w_{it} + e_{it}, \tag{24}$$

i.e. a standard MaCurdy regression which allows to estimate η^{-1}, the intertemporal (Frisch, or λ-constant) elasticity of labor supply. We interpret e_{it} as containing an individual fixed-effect, and label (24) the "micro" regression.

Aggregating hours supplied to the market, i.e. equation (16), across individuals who are employed – i.e. across the $n_t \leq N$ workers – yields aggregate labor supply at time t, denoted H_t:

$$H_t = \sum_{i=1}^{n_t} \left(\frac{\lambda_{it} w_{it}}{\alpha}\right)^{\frac{1}{\eta}}. \tag{25}$$

Likewise, the average wage of workers, denoted W_t, is equal to:

$$W_t \equiv n_t^{-1} \sum_{i=1}^{n_t} w_{it}, \tag{26}$$

where we are assuming that the first n_t individuals are market workers. We treat the working-age population, N, as constant. This implies that the market/home production choice reflects only relative productivities, which are not related to age or other demographic factors. We denote by $\overline{W}_t$ the "imputed" average wage, i.e. the weighted average of the observed market wages for actual workers (w_{it}) and the unobserved reservation wages $(\widetilde{w}_{it})$ claimed by non-workers:

[2] See Blundell and MaCurdy (1999, p. 1597, footnote 13) for details.

$$\overline{W}_t \equiv N^{-1}\left(\sum_{i=1}^{n_t} w_{it} + \sum_{i=n_t+1}^{N} \widetilde{w}_{it}\right), \tag{27}$$

which can also be written as:

$$\overline{W}_t \equiv N^{-1}\left[n_t W_t + (N-n_t)\,\widetilde{\theta}_t^H\right]. \tag{28}$$

Here, $\widetilde{\theta}_t^H$ denotes the average productivity at home of non-workers whereas the "imputed" wage variable $\overline{W}_t$ is the same as the imputed average productivity of both market and home workers. We denote by $\overline{\Theta}_t$ such an average. Actual and "imputed" mean wages are balanced in each period by a parameter δ_t:

$$\overline{W}_t = W_t^{\delta_t}. \tag{29}$$

Substituting the identity $w_{it} \equiv \left(\theta_{it}/\overline{\Theta}_t\right)\overline{W}_t \equiv \left(\theta_{it}/\overline{\Theta}_t\right)W_t^{\delta_t}$ into (25) and taking logs yields:

$$\ln H_t = k + \frac{\delta_t}{\eta}\ln W_t + \frac{1}{\eta}\ln\sum_{i=1}^{N}\lambda_{it}\theta_{it} + v_t, \tag{30}$$

where $v_t \equiv -\eta^{-1}\ln\overline{\Theta}_t$. If we take the average market wage of workers – which we observe in the data – to be the appropriate aggregate wage rate, then the Frisch macro labor elasticity is δ_t times the micro elasticity. This reflects the individual trade-off between market and home production, i.e. decisions on the extensive margin. The aggregate elasticity is larger than the individual one when $\delta_t > 1$, which is equivalent to $\overline{W}_t > W_t$. From equation (28) it is immediate to see that this is the case when:

$$\widetilde{\theta}_t^H > W_t, \tag{31}$$

i.e. when the average reservation wage of non-workers is larger than the average market wage of market workers.

Condition (31) is testable, provided equation (30) is identified. The problem here is the same encountered in the micro case: neither λ_{it} nor θ_{it} are observed for all individuals. However, we can rewrite equation (22) as

$$\lambda_{it}\theta_{it} = -\lambda_{it-1}\theta_{it}\beta\,(1+r)\exp\left(v_{it}\right). \tag{32}$$

Aggregating across all individuals and taking logs yields:

$$\ln \sum_{i=1}^{N} \lambda_{it}\theta_{it} = -\ln \beta \left(1 + r\right) + \ln \sum_{i=1}^{N} \lambda_{it-1}\theta_{it} \exp \left(\upsilon_{it}\right). \tag{33}$$

Define $E_t = \upsilon_t + \ln \sum_{i=1}^{N} \lambda_{it-1}\theta_{it} \exp \left(\upsilon_{it}\right) - \ln \sum_{i=1}^{N} \lambda_{it-1}\theta_{it-1}$, replace in (33) and back into (30). After rewriting the latter in first-differences, we obtain the following equation:

$$\Delta \ln H_t = Const. + \frac{\delta}{\eta} \Delta \ln W_t + E_t, \tag{34}$$

where we replace for estimation purposes the time-varying parameter δ_t with its constant counterpart δ. This simplifying assumption is not too strong because in our PSID sample the employment rate ranges between 91% and 94% with a coefficient of variation which is rather small, about 1%.[3] To summarize, we will estimate models (24) and (34) as follows:

$$\text{individual} \quad : \quad \Delta \ln h_{it} = \textit{const.} + \epsilon \Delta \ln w_{it} + e_{it}, \tag{35}$$

$$\text{aggregate} \quad : \quad \Delta \ln H_t = \textit{Const.} + \mathcal{E} \Delta \ln W_t + E_t. \tag{36}$$

Comparing equations (34) and (36) it should be noted that the aggregate elasticity parameter $\mathcal{E}$ does not reflect the extensive margin only but accounts also for the intensive elasticity parameter η. The latter plays no role when the reservation wage heterogeneity occurs instead in an indivisible labor model. Hence, although our results are empirically similar to those of Chang and Kim (2005; 2006), our aggregate elasticity is still related to the willingness to substitute leisure over time.

4 Results

Our estimates use data obtained from the core sample of the PSID. This is not a costless choice because in this panel one does not find, for all waves, such important labor supply variables as wealth, tax rates, and the real interest rate.

[3]The PSID employment rate is high compared to the US population since it over–represents employed individuals (see Data Appendix).

However, it has an important advantage: it covers 35 years, thus allowing for the construction of a relatively long time series for this type of data. For estimation purposes, however, our series is shorter than it might otherwise be. The reason is that PSID data were collected annually from 1968 to 1997, and every two years afterwards: in order to avoid arbitrary interpolation of the microdata, we preferred using only the annually released portion of the panel.

We aggregated each wave to create a macro series to be used for comparing in a fully consistent way the micro and the macro labor supply elasticity. We are aware that our sample is not representative of the US population and, therefore, we do not claim to provide the right estimate of the aggregate elasticity of labor supply in the US. This is not the goal of the paper which instead is estimating the micro and macro elasticities on the basis of a consistent aggregation procedure. However, when we compare in the Data Appendix the properties of our series with aggregate US data, it turns out that they are not too dissimilar.

As in MaCurdy (1981), we exclude from the sample permanently disabled or retired individuals, i.e. we include only those units displaying nonzero wage and labor supply data in any particular year. We also use two dummy variables to account for important modifications underwent by the PSID in 1993 and 1996 (see the Data Appendix for details).

It is well known that wage reported in the PSID may be affected by substantial measurement errors (Pischke 1995). Such errors are likely to be washed-out by aggregation but remain a concern in the individual regression. As a check, we will later exclude self-employed individuals – wages in this category are more likely to be affected by relevant measurement errors.

In estimating the first-difference equation (35) we use, as in MaCurdy (1981), the fixed effects estimator to mitigate the problems arising from the limitations of the data. While this prevents us from estimating the long-run labor supply response to productivity, it should also avoid mixing substitution and income effect, since the latter is likely to prevail in the long-run.

In a rational expectations framework, we use lags as instruments to account for the endogeneity of the real wage. The perfect correspondence between our micro and macro estimates is ensured – among the other things – by the fact that we use exactly the same instruments in both cases. Therefore, the autoregressive terms enter the micro and the macro equations with exactly the same lags, although aggregation may change the dynamics pertaining to each individual component [Granger and Newbold (1986)]. This possibility is another reason for regressing first-differenced data which tend to reduce (or eliminate) differences in persistence, otherwise to be dealt with via a longer series of instruments in the macro estimate.

Our main result is summarized in Table 1. Column 1 reports instrumental variables fixed-effects estimates of the individual elasticity while Column 2 reports instrumental variables estimate of the aggregate, time-series, elasticity. The LHS variables are the variation in log individual and aggregate hours, respectively. The RHS variables are the variation in log individual and aggregate wage rates deflated by the consumer price index. Instruments are in both cases the 2nd, 3rd and 4th lags of the individual and aggregate wage rates, respectively[4].

Table 1. Individual and aggregate Frisch elasticities.

	Individual $\Delta \ln(h_{it})$ 1	Aggregate $\Delta \ln(H_t)$ 2
$\Delta \ln(\text{wage})$	0.11**	1.06**
	(0.05)	(0.49)
Constant	-0.00	0.01**
	(0.00)	(0.00)
J-stat	0.28	3.36
p-value	0.87	0.19
Observations	42,280	26
Individuals	4,276	-

* Significant at 10%; ** Significant at 5% or better

The macro elasticity (1.06) is much larger than the estimated individual elasticity (0.11): a result which is very robust, being always confirmed by side estimates (not shown) where we check for the role of nonlabor income and for the use of pre-1993 data only, to avoid the inclusion of the 1993 and 1996 dummies. The effect of nonlabor income coefficient is negative and negligible in the (unchanged) micro equation while negative and insignificant in the macro regression where the wage elasticity increases to 1.4. When using pre-1993 data only, the individual elasticity remains 0.1, while the aggregate elasticity increases to 1.5.

[4] Instruments are in log levels rather than in log differences for the efficiency reasons outlined by Arellano (1989).

Finally, when using the full - rather than the core - PSID sample, the individual elasticity is still 0.1 while the aggregate is only slightly reduced (0.9).[5]

We find noteworthy that the aggregate elasticity is not only an order of magnitude larger than the individual one but that is also consistent with the unit value often assumed in calibration studies. The J-stat shows that instrument over-identification is not a concern in the micro estimate. This test is less satisfactory in the aggregate estimate, probably because the aggregate wage is less related to aggregate productivity.[6]

Table 2. Individual Frisch elasticity: Heckit and employees only

	$\Delta \ln (h_{it})$		
	1	2	3
$\Delta \ln (\text{wage})$	0.11**	0.17**	0.17**
	(0.05)	(0.07)	(0.07)
Constant	-0.00	-0.00	-0.00
	(0.00)	(0.00)	(0.00)
Heckit	yes	no	yes
Self-employed	yes	no	no
J-stat	0.28	1.49	1.48
p-value	0.87	0.47	0.48
Observations	42,280	38,169	38,169
Individuals	4,276	4,076	4,076

* Significant at 10%; ** Significant at 5% or better

Since the estimate of the individual elasticity may suffer from ignoring the zero-hours wages, we present in Table 2 the results obtained with an "Heckit" estimator.[7] The individual elasticity (see Column 1) does not differ from the benchmark elasticity reported in Table 1. To control for a possible source of measurement error in the wage rate and still using the same instruments as in the

[5]All of these regressions are available upon request.

[6]This may be due to the fact that bargaining and other forms of interaction introduce some dependence among the single units.

[7]The predicted probability of participation (positive hours) is produced by a fixed-effects probit regression of the participation indicator on age, sex, family income and size. These are then used to construct the inverse Mills ratio, whose first difference is used as an additional regressor.

benchmark regression, we excluded-self employed individuals. The corresponding elasticity is reported Column 2. The same result is obtained in the corresponding "Heckit" regression (Column 3): in both cases, the micro elasticity is slightly increased (0.17) with respect to the benchmark case (0.11). An implication of Table 2 is that correcting for selection does not introduce into our individual estimates the same shift which occurs when the role of the extensive margin is made explicit, i.e. when the decision of entering or leaving the labor market is made observable through the aggregation channel.[8]

In Table 3 we split the Macurdy regression (36) into the two terms of equation (5), separately estimating, using the same instruments, the aggregate elasticity of average hours per worker and the employment level. The aggregate elasticity of hours per worker is 0.16 – a value that not surprisingly is close to the individual elasticity of Tables 1 and 2 – while the elasticity of employment, i.e. the extensive margin, is 0.9. Therefore, the latter accounts for virtually all of the difference between micro and macro elasticities, confirming the unconditional stylized facts evidence on the relative importance of the two margins of adjustment. In both the individual and the aggregate regressions the constant has a natural interpretation in terms of a time trend in hours worked. Therefore, in our sample, aggregate labor supply increases by about 1% per year, with no trend in average hours worked by individuals. This is in fact possible if the extensive margin is the principal margin of adjustment.

Table 3. Decomposition of the aggregate elasticity

	$\Delta \ln \left(\overline{h}_t\right)$	$\Delta \ln \left(n_t\right)$
	1	2
$\Delta \ln \left(\text{wage}\right)$	0.16	0.90
	(0.30)	(0.47)
Constant	-0.00	0.01
	(0.00)	(0.00)
Observations	26	26

* Significant at 10%; ** Significant at 5% or better

[8]It is well known that in practice Heckman-correction may not make a big difference [Moffitt (1999)].

5 Conclusions

Using PSID data for about 30 years (1967-96), we consistently compare the individual and the aggregate Frisch elasticities of labor supply in a MaCurdy-type equation. The panel estimate of the individual labor supply elasticity differs by an order of magnitude from the time-series estimate obtained by aggregating each year the hours worked in the sample. For the micro elasticity we find a low value (0.11) in line with mainstream empirical results. For the macro elasticity, we find a relatively large value (1.06) which is not too far from the pioneering estimate of Lucas and Rapping (1969). This value also suggests that a large body of business cycle analysis based on a log-log specification of preferences might have a sound empirical base.

This difference between micro and macro elasticities is not new in the literature and is often invoked as a reason for rejecting the RBC model. Our result, however, shows that there is no contradiction between the two, because they pertain to two different variables and concepts: the intensive margin in one case and its product with the much more volatile extensive margin, in the other. The issue is not aggregation bias. The underlying utility maximization model aims at explaining the dominance of the extensive margin on the basis of the intertemporal and intra-temporal choice between leisure and labor to be allocated to market- or home-production.

We recognize a number of limitations in our data and we do not interpret our results as necessarily relevant for the US and even less for other countries for which sufficiently long data are unavailable. In particular, we cannot employ a number of important controls such as marginal tax rates, individual wealth, and the individual interest rate.

The main contribution of this paper is showing that aggregation alone leads to a much larger elasticity than one finds in micro estimates because micro estimates are apparently unable to reflect participation decisions, even when using some selectivity correction mechanism. We regard our work as a simple empirical exercise, but we are not aware of other empirical studies indicating the relevance of the extensive margin via the exact aggregation of the individual units. Our result also suggests a methodological point: parameter estimates from micro data are not always appropriate for calibrating an aggregate model economy.

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6 Data Appendix

In this appendix we provide additional details on our dataset. In particular, we compare the time series derived from aggregating individuals in the PSID with official aggregate US data (sources: BLS and OECD labor statistics).

Figure 1 reports series of the means of key demographic and labor variables (age, sex, employment rate). This gives a rough picture of the population we are working with, as well as its dynamics. It is clear that our sample is not representative of the US population:[9] men and workers are substantially over-represented. The reason is that we use family rather than individual data – full longitudinal information on labor supply and wages is available only in the 'family' portion of the PSID. As a consequence, our units of observation are household heads: these, by convention, are males when a family includes a married couple.

Figure 1. Sample statistics.

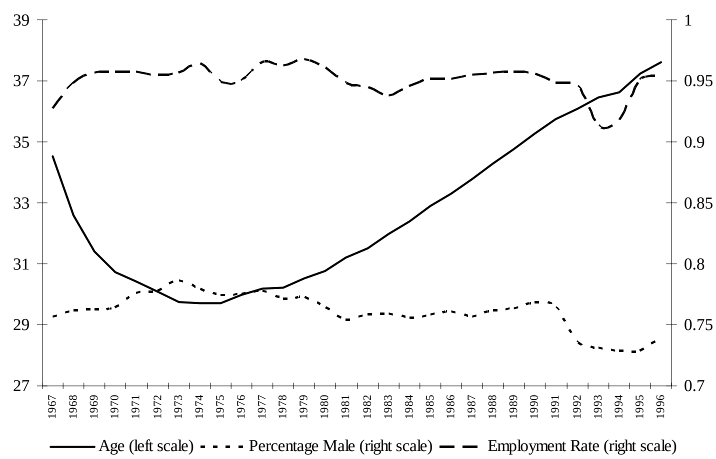

Figure 2 shows the series of average hours worked by employed individuals. Both series refer to total employment. The US aggregate series is smoother than the PSID series. Although they tend to move together (the coefficient of correlation is 0.51), there is a noticeable difference in levels: on average, individuals in our PSID sample work every year 100-200 hours more than US workers. This is at least partly explained by the over-representation of men, who typically work more than women. The different volatility may

[9]These are unweighted statistics. Using sampling weights (as provided by the PSID) does not change things significantly. Weighted means are available from the authors upon request.

be due to changes occurring in the panel due to attrition and—most importantly—to the inclusion of new households in the survey.

Figure 2. Average hours worked.

Table 4. Variation in the composition of the PSID (annual waves)

Year Collected	Sample Size	Variation	Year Collected	Sample Size	Variation
1967	4802		1982	6852	1.63%
1968	4460	-7.12%	1983	6918	0.96%
1969	4654	4.35%	1984	7032	1.65%
1970	4840	4.00%	1985	7018	-0.20%
1971	5060	4.55%	1986	7061	0.61%
1972	5285	4.45%	1987	7114	0.75%
1973	5517	4.39%	1988	7114	0.00%
1974	5725	3.77%	1989	9371	31.73%
1975	5862	2.39%	1990	9363	-0.09%
1976	6007	2.47%	1991	9829	4.98%
1977	6154	2.45%	1992	9977	1.51%
1978	6373	3.56%	1993	10764	7.89%
1979	6533	2.51%	1994	10401	-3.37%
1980	6620	1.33%	1995	8511	-18.17%
1981	6742	1.84%	1996	6747	-20.73%

This issue is illustrated in Table 4, where we report the percentage variations in the PSID sample during the period 1967-1996. While changes in the composition of the panel is not a concern when estimating the individual elasticity—the only consequence is an unbalanced panel—this is not the case for the aggregate elasticity. The reason is that variations in employment will reflect exogenous modifications of the sample. The troublesome years are those when the PSID underwent substantial modifications. After minor variations until 1988 (except 1968, which does not affect our estimated since we use 3 lags as instruments), we notice a few major changes. First, the 1990 wave (containing data collected in 1989) was almost a third larger than the previous one. The reason is that 2,200 new households, the so-called Latino sample, were added to take into account the substantial demographic changes occurred in the US since the inception of the PSID. The remaining part of the Latino sample was subsequently dropped in 1995. This does not affect our estimates because we work with the so-called core sample only. The 1994 wave (i.e. our 1993 data) was also characterized by a sizeable increase in sample size because of the inclusion of a large number of 'recontacts', i.e. persons who had been lost during the 10 previous years. This is a significant exogenous increase in sample size we need to control for with a dummy variable. Finally, the 1997 wave (i.e. our 1996 data) was released after a major reduction in sample size due to the attempt to reduce the cost of the survey. This is another major exogenous variation we control for with a dummy variable. Results from estimation using the shorter time series 1967-1992 show that in fact these two dummies are appropriate and actually provide a conservative estimate.

Figure 3. Variation of log mean hours per worker

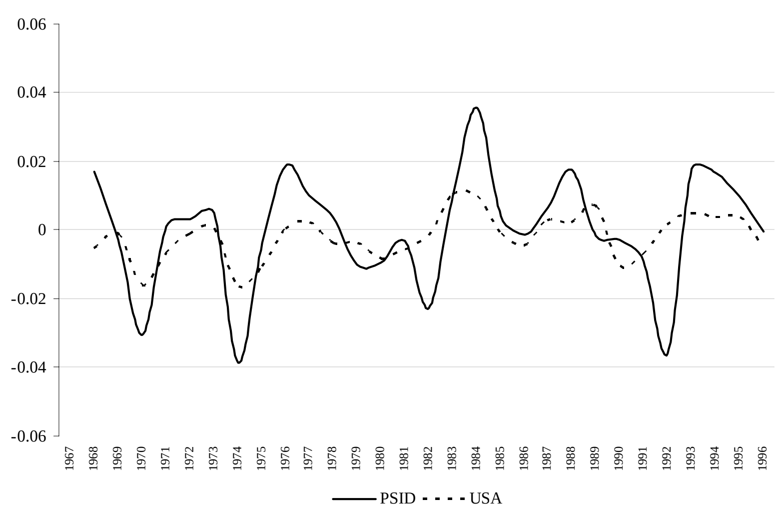

Figure 3 summarizes the same information reported in Figure 1 taking logs and using first differences. The figure confirms that hours per workers in our sample and in the US population, despite the differences noticed above, tend to co-vary (the coefficient of correlation of these transformed series in 0.64).

Figure 4 compares variations is the log real wage rate in the PSID and in the US. The two series are not fully comparable because the US series is released by the BLS only for private nonfarm workers involved in production and non-supervisory tasks. All nominal values are converted into real terms using the CPI. The two series move quite closely together, at least until the end of the 1980s.

Figure 4. Variation of log average real wage.

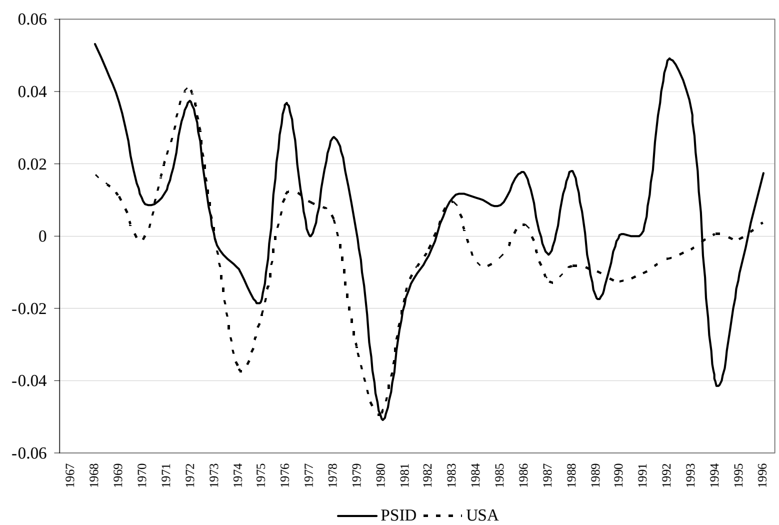

Figure 5 shows the variation of log employment—the extensive margin–in the PSID and the US. Again, while the PSID series refers to total employment in the sample, the US series represents private nonfarm employment. The large variations in the PSID series at the extremes (1968 and 1996) is explained by the large variations in the composition of the sample in those years—see Table 5. Since, as we noticed above, year 1968 is not used for estimation and year 1996 is alternately controlled for by a dummy or a shorter series, such large variations do not affect our estimates. Between the two extremes and ignoring year 1993 which is also controlled for by a dummy, the two series display some degree of co-movement, especially in the first part (coefficient of correlation 0.49 until mid-1985, and 0.24 overall), making us confident that we are not mixing, to any significant extent, variations in employment and in sample composition.

Figure 5. Variation of log employment.