Labor Supply Elasticities: Can Micro be Misleading for Macro?

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Labor Supply Elasticities: Can Micro be Misleading for Macro?1

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INTRODUCTION

The intertemporal substitution between work and leisure is crucial for the explanation of aggregate fluctuations 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.” (op. cit., p. 14).

When labor supply is involved things are more complicated: it is well known that the RBC model requires a much larger elasticity than those estimated in micro studies (ranging from about 0 to about 0.2 for men and from about 0 to about 1 for married women; see Blundell and MacCurdy, 1999) in order to reproduce virtually all business cycles facts. This difficulty is pointed out by several authors [e.g. Heckman (1993); Browning, Hansen and Heckman (1999)], and by Prescott (2006) himself.

The research question we address in this paper is whether and, more importantly, to what extent a small individual labor supply elasticity is consistent with a large aggregate labor supply elasticity. This question, for several reasons, cannot be answered by looking at existing empirical studies. Microeconomic studies based on both cross-sectional and panel data generally report a small real-wage elasticity [e.g. Pencavel (1986), Killingsworth and Heckman (1986), MaCurdy (1981), and Altonji (1986)], which applies to the intensive margin only, i.e. hours per worker. On the other hand, the macroeconomic evidence is far less numerous and is generally mixed. 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 necessity of reconciling the relatively high aggregate elasticity used 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, and Mankiw, 1989) the whole relevance of the RBC model was denied. A more constructive orientation explored several variants of the benchmark RBC model (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 government consumption (Christiano and Eichenbaum 1992), the home production model (Benhabib, Rogerson and Wright, 1991) and the introduction of taxation in general equilibrium models (Baxter and King, 1993, and McGrattan, 1994) are all noteworthy extension of the benchmark RBC model (i.e.
More recent studies that generate a wedge between individual and aggregate labor elasticities range from heterogeneous reservation wages (Chang and Kim, 2005, 2006) to the omission of such different variables as wealth (Ziliak and Kniesner, 1999), liquidity constraints (Domeji and Floden, 2006) and human capital accumulation (Imai and Keane, 2004), to nonlinearities in the relation between labor services and hours of work (Rogerson and Wallenius, 2007). A large aggregate elasticity is also required to explain the difference in patterns of work in Europe and the US on the basis of different tax rates (Prescott, 2004). Needless to say, this list is incomplete.

In this paper we take a different, empirical, route based on a testable aggregation principle. We use all of the annual waves (1968-1997) of the Panel Study of Income Dynamics (PSID) to estimate the Frisch labor supply elasticity via a long-enough panel to be compared with the corresponding time series, resulting from the exact aggregation of individual units each year. Therefore, our exercise is the following: we take the PSID as some population of interest and ask two questions. What is the micro elasticity in this population? What is the macro elasticity in this same population? We address in detail possible problems associated with this procedure, and we show how far this population is from the US population. Thanks to this procedure, the elasticities obtained from the aggregate series and from the individual data are fully consistent. Therefore, we can assess their relative magnitude regardless of the value of our micro estimate. In this sense, our goal is mainly methodological. To our knowledge, what we do is something new despite its conceptual simplicity.

We show that microeconomic estimates are not a good source for calibrating total worked hours: our panel results, which by and large follow the pioneering study of MaCurdy (1981), deliver a Frisch elasticity of about 0.1, a small value that is consistent with benchmark estimates, while the aggregate time-series delivers a Frisch elasticity near 1, a much larger value. Moreover, we decompose the aggregate elasticity into the contribution of adjustment of hours (intensive margin) and of employment (extensive margin), finding that the latter accounts for about 3/4 of the difference between the two elasticities. This means that such a difference is mainly due to the positive covariance between number of workers and the wage rate. These results complement the findings of Rogerson and Wallenius (2008) as well as Chang and Kim (2005, 2006), which are based instead on calibrating the aggregate economy. Furthermore, they are consistent with the well-known fact that most of the changes in aggregate hours stem from the extensive rather than from the intensive margin (Kydland, 1995). Probably this explains why the intertemporal substitution hypothesis is not rejected when applied to aggregate employment, as the work of Alogoskoufis (1987) shows for the US.

We are well aware of the fact that our procedure is not costless, for reasons we discuss later in the paper. However, it releases a number of benefits. In particular, the macro dataset is based on exactly the same units of observation that compose the micro dataset, and allows us to employ exactly the same estimation method, including the same instruments, in two

---

2 PSID data was collected every year from 1968 to 1997 and every two years from 1997 onward. As we explain below, this poses problems which we prefer to avoid by truncating the series in 1997.

3 In principle, there are problems in estimating the aggregate elasticity using annual micro data, because intertemporal substitution is a property of both preferences and the technology (see Prescott, 2006). We do not address this issue here.

4 In particular, Chang and Kim assume an individual elasticity of 0.4 and find an aggregate elasticity of about 1.
“isomorphic” micro and macro regressions.

The remainder of 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 the dataset and Section 5 the results. Sections 6 concludes.

INTENSIVE VS. EXTENSIVE MARGIN

The indivisible labor case (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. Like in Hansen (1985), if we denote by \( n \) the employment stock and by \( \bar{h} \) average supply of hours, then aggregate labor is \( H_t \equiv n_t \bar{h}_t \). By taking logs, the variance of labor

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

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), for which variance is scale-free, Fiorito and Kollintzas (1994) found that the variance of employment deviations from the smooth trend always exceed the corresponding variance in the 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 labor supply elasticities reflects movements along 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 \( \varepsilon \) and \( \xi \) the micro and macro Frisch elasticities of labor supply, respectively, and by \( w_t \) and \( W_t \) the individual and aggregate wage rates at time \( t \), respectively.

Consider the following regression models, which we derive below in detail:

**Individual** \[
\Delta \ln h_t = \text{const.} + \varepsilon \Delta \ln w_t + \epsilon_t, \tag{2}
\]

**Aggregate** \[
\Delta \ln H_t = \text{Const.} + \xi \Delta \ln W_t + \xi_t. \tag{3}
\]
The population elasticities are:
\[
\epsilon = \frac{\text{cov}(\Delta \ln h_i, \Delta \ln w_i)}{\text{var}(\Delta \ln w_i)},
\]
\[
\tilde{\epsilon} = \frac{\text{cov}(\Delta \ln H_i, \Delta \ln W_i)}{\text{var}(\Delta \ln W_i)} = \frac{\text{cov}(\Delta \ln \bar{h}_i, \Delta \ln \bar{W}_i)}{\text{var}(\Delta \ln \bar{W}_i)} + \frac{\text{cov}(\Delta \ln n_i, \Delta \ln W_i)}{\text{var}(\Delta \ln W_i)}.
\]

That is, the micro elasticity (5) consists of a single term — the covariance between individual hours and individual wage rate — which 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, the covariance between employment and the aggregate wage rate, 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). Our goal is to estimate them in a consistent way. In the next Section we develop a model that allows to address the ensuing aggregation problem in a legitimate way.

**THE MODEL**

Consider an economy populated by \(N\) individuals, indexed by \(i = 1, \ldots, N\), and 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 and the same endowment of labor services but differ in productivity.

Denote by \(\theta_{it}\) and \(\theta_{it}^H\) individual \(i\)'s productivity on the market and at home, respectively, at time \(t\), and by \(h_{it}\) and \(h_{it}^H\) fraction of hours spent producing on the market and at home. Market and home productivities evolve stochastically. The total amount of the consumption good in the economy at time \(t\) is
\[
c_t = c_t^M + c_t^H,
\]
where
\[
c_t^M = \sum_{i=1}^{N} \theta_{it} h_{it},
\]
\[
c_t^H = \sum_{i=1}^{N} \theta_{it}^H h_{it}^H.
\]

In other words, work on the market and at home are perfect substitutes in production, and
the respective outputs are perfect substitutes in consumption. Labor services can be sold on the market at an individual-specific and time-varying wage rate \( \omega_{it} \). Profit maximization implies that \( \omega_{it} = \theta_{it} \) 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. We also allow heterogeneity in the endowment of assets and other exogenous sources of income.

Preferences are defined over consumption \( (c) \) and leisure \( (l) \), and are represented by the utility function \( u(c_i, l_i) \) a strictly increasing, twice differentiable, strictly quasi-concave function. 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 the expected discounted present value of the utility stream, given the budget and time constraints:

\[
\max_{\{c_{it}, h_{it}, h_{it}^H, a_{it+1}\}} \mathbb{E}_0 \left[ \sum_{t=0}^{\infty} \beta^t u(c_{it}, l_{it}) \right]
\]

subject to:

\[
c_{it} + a_{it+1} \leq \omega_{it} h_{it} + \theta_{it} h_{it}^H + (1 + r) a_{it} + z_{it},
\]

\[
h_{it} + h_{it}^H + l_{it} = 1,
\]

and the no-Ponzi game condition: \( \lim_{T \to \infty} \beta^T \frac{\partial u(c_{iT}, l_{iT})}{\partial a_{iT}} a_{iT+1} = 0 \). In this problem, \( \beta \) is the discount factor and \( r \) the real return on assets. Both are assumed, due to data limitations, to be invariant in time and across individuals. Finally, \( z_{it} \) summarizes other exogenous sources of income. 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:

\[
u_M(h_{it}), \nu_H(h_{it}^H)
\]

Denoting by \( \lambda_{it} \) the Lagrange multiplier, i.e. marginal utility of wealth, and by \( \nu_{it}^M \) and \( \nu_{it}^H \) 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:

\[
c_{it} : c_{it}^{-\gamma} = \lambda_{it},
\]

\[
h_{it} : \alpha (h_{it} + h_{it}^H)^\gamma = \lambda_{it} \omega_{it} + \nu_{it}^M,
\]

\[
h_{it}^H : \alpha (h_{it} + h_{it}^H)^\gamma = \lambda_{it} \theta_{it}^H + \nu_{it}^H,
\]

\[
a_{it+1} : \lambda_{it} = \beta (1 + r) \mathbb{E}_t [\lambda_{it+1}],
\]

\[
\lambda_{it} : c_{it} + a_{it+1} = h_{it} \omega_{it} + \theta_{it}^H h_{it}^H + (1 + r_t) a_{it} + z_{it}.
\]

5 This is a special case of the general CES composition of Benhabib, Rogerson and Wright (1991).
It is straightforward to see that individuals will either supply a positive number of hours to the market or spend a positive number of hours at home, but never both. Specifically, in equilibrium:

\[
\begin{align*}
    h_{it} &= \begin{cases} 
    (\lambda_{it} w_{it} / \alpha)^{1/\eta} & \text{if } \theta_{it}^H \geq \theta_{it}^H \\
    0 & \text{otherwise},
    \end{cases} \\
    h_{it}^{H} &= \begin{cases} 
    (\lambda_{it} \theta_{it}^H / \alpha)^{1/\eta} & \text{if } \theta_{it} < \theta_{it}^H \\
    0 & \text{otherwise}.
    \end{cases}
\end{align*}
\]

It follows that \( \theta_{it}^H \) is exactly individual \( i \)'s reservation wage at time \( t \), which we denote \( \tilde{\omega}_{it} \). Therefore, individual \( i \) works on the market at time \( t \) if the reservation wage is below market productivity.

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

\[
\begin{align*}
    \ln c_{it} &= -\frac{1}{\gamma} \ln \lambda_{it}, \\
    \ln h_{it} &= k + \frac{1}{\eta} \ln \lambda_{it} + \frac{1}{\eta} \ln \omega_{it}, \\
    \ln \lambda_{it} &= \ln \beta (1 + \gamma) + \ln E \left[ \lambda_{it+1} \right],
\end{align*}
\]

where \( k \equiv -\eta^{-1} \ln \alpha \) a constant. Equation (19) cannot be estimated, since we do not observe \( \lambda_{it} \). Blundell and MaCurdy (1999, p. 1597) show that if the expected (exponential) forecast error is approximately constant, then (20) implies the following stochastic process for \( \lambda_{it} \):

\[
\ln \lambda_{it} = \lambda_0 + \ln \lambda_{it-1} + \nu_{it},
\]

where \( \lambda_0 \) is a constant. Denoting for any variable \( X \), \( \Delta X_t \equiv X_t - X_{t-1} \) rewrite (19) in first differences:

\[
\Delta \ln h_{it} = \frac{1}{\eta} \Delta \ln \lambda_{it} + \frac{1}{\eta} \Delta \ln \omega_{it}.
\]

Replacing (21) into this equation and defining \( e_{it} \equiv \eta^{-1} \nu_{it} \) we obtain to the following estimable equation:

\[
\Delta \ln h_{it} = \text{const.} + \frac{1}{\eta} \Delta \ln \omega_{it} + e_{it}.
\]

Equation (11) allows to estimate \( \eta^{-1} \), the intertemporal (Frisch, or \( \lambda \)-constant) elasticity of labor supply. We label (23) the micro regression. We derive the analogous macro regression by aggregating the individual units. Aggregating hours supplied to the market, i.e. equation (16), across individuals who are employed — i.e. workers, these are \( n_t \leq N_t \) — yields aggregate labor supply at time \( t \), denoted \( H_t \):

\[
H_t = \sum_{i=1}^{n_t} \left( \frac{\lambda_{it} w_{it}}{\alpha} \right)^{1/\eta}.
\]

The average wage of workers, denoted \( W_t \), is equal to:
where we are adopting the convention that the first \( n_t \) individuals are workers. On the other hand, the “imputable” average wage — i.e. the weighted average of observed market wages (actual) for workers and unobserved reservation wages (claimed) for non-workers — is denoted and is eq \( \bar{W}_t \):

\[
\bar{W}_t = \frac{1}{N} \left( \sum_{i=1}^{n_t} w_{it} + \sum_{i=n_t+1}^{N} \tilde{w}_{it} \right).
\]  

(26)

This can also be written as:

\[
\bar{W}_t = \frac{1}{N} \left( n_t W_t + (N - n_t) \tilde{\theta}_t^H \right),
\]

(27)

where \( \tilde{\theta}_t^H \) is the average productivity at home of non-workers. \( \bar{W}_t \) is of course identically equal to the average of market productivities of workers and home productivities of non-workers. We denote with \( \tilde{\theta}_t \) such an average. Actual and “imputed” mean wages are balanced by a parameter \( \delta_t \):

\[
\bar{W}_t = W_t^{\delta_t}.
\]

(28)

By construction, this balancing parameter is related the extensive margin: except for the hairline case in which the average home productivity of non-workers is equal to the average market productivity of workers, \( \delta_t \) is equal to 1 only if everybody is working on the market, in which case movements along the extensive margin are not possible. Using the identity

\[
\omega_{it} \equiv \left( \frac{\theta_{it}}{\bar{\theta}_t} \right) \bar{W}_t = \left( \frac{\theta_{it}}{\bar{\theta}_t} \right) W_t^{\delta_t}
\]

replacing this into (24) and taking logs yields

\[
\ln H_t = k + \frac{\delta_t}{\eta} \ln W_t + \frac{1}{\eta} \ln \left( \sum_{i=1}^{N} \lambda_{it} \theta_{it} + v_t \right).
\]

(29)

where \( v_t \equiv -\eta^{-1} \ln \bar{\theta}_t \). If we take the average market wage of workers — which we observe in the data — to be the appropriate aggregate wage rate, then the macro (Frisch, or \( \lambda \)-constant) labor elasticity is \( \delta_t \) times the micro elasticity. This reflects the individual trade-off between market and home production, which accounts for the extensive margin. Therefore, in our model the macro and the micro elasticities are conceptually different. The former is larger than the latter when \( \delta_t > 1 \), which is equivalent to \( \bar{W}_t > W_t \). From equation (27) it is immediate that this is the case if and only if

\[
\tilde{\theta}_t^H > W_t.
\]

(30)

In other words, the average reservation wage of non-workers must be larger than the average market wage of workers.\(^6\)

Condition (30) is testable, provided equation (29) can be identified. The problem here is the same encountered in the micro case: neither \( \lambda_{it} \) nor \( \theta_{it} \) are observed for all individuals. However, we can rewrite equation (21) as

---

\(^6\) Therefore, in our model the individual and aggregate elasticities are different because, like in Chang and Kim (2005), the reservation wage distribution is nondegenerate and so there are some individuals who are adjusting on the extensive margin in response to wage shocks.
\begin{equation}
\lambda_{it}\theta_{it} = \lambda_{i,t-1}\theta_{it-1}\exp(\lambda_0 + \nu_{it}) ,
\end{equation}

aggregate across all individuals and take logs. This yields:

\begin{equation}
\ln \sum_{i=1}^{N} \lambda_{it}\theta_{it} = \lambda_0 + \ln \sum_{i=1}^{N} \lambda_{i,t-1}\theta_{it-1} \exp(\nu_{it}) .
\end{equation}

If we define a new error term \( E_t = \Delta v_t + \ln \sum_{i=1}^{N} \lambda_{i,t-1}\theta_{it-1} \exp(\nu_{it}) - \ln \sum_{i=1}^{N} \lambda_{i,t-1}\theta_{it-1} \), replace in (32) and back into (29) after rewriting this in first-differences, we obtain the following estimable equation:

\begin{equation}
\Delta \ln H_t = \frac{\delta_t}{\eta} \Delta \ln W_t + E_t.
\end{equation}

For estimation purposes we treat \( \delta_t \) as a constant.\(^7\) To summarize, referring to the notation introduced earlier in the paper, we will estimate models (23) and (33) where a constant is added to account for a possible trend:

\begin{align*}
\text{Individual} & \quad \Delta \ln h_{it} = \text{const.} + \varepsilon \Delta \ln w_{it} + e_{it}, \\
\text{Aggregate} & \quad \Delta \ln H_t = \text{Const.} + \varepsilon \Delta \ln W_t + E_t.
\end{align*}

In sum, our simple theoretical framework provides a precise link between the individual and the aggregate level, i.e. solves the aggregation problem, so that we can compare individual and aggregate elasticities in a meaningful way (see Blundell and Stoker, 2005).

Of course the first difference of the relevant wage rate is endogenous because it is determined in equilibrium. In a rational expectations framework, we use lags as instruments.\(^8\) Furthermore, it is well known that the estimate of the individual elasticity may suffer from ignoring the zero-hours wages. We show below that our results are robust to performing selection-correction using the “Heckit” estimator. We use a model in first differences and the fixed effects estimator, which considerably reduces the need for controls thus mitigating the limitations of our data. This also allows us to preserve the correspondence between micro and macro regressions. However, we check for the role of non-labor income (like income from assets, intrafamily and public transfers), which are of obvious importance. This way we will estimate first-difference versions of what Heckman (1993) labels the labor supply of workers and the aggregate labor supply curve, respectively.

**DATA**

Our data come from the Panel Study of Income Dynamics (PSID), a panel of about 8,000

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7 This assumption is not too strong because in our PSID sample, the employment rate ranges between 91% and 94% with a coefficient of variation of 1%. The employment rate is high compared to the US population because, as we comment below in detail, our sample over-represents employed individuals.

8 This is also justified by the autoregressive structure of individual productivity, which we don’t write explicitly. Levels can be preferred to changes for the efficiency reasons mentioned by Arellano (1989).
households. This choice has an important disadvantage: PSID data do not report for all waves such important variables as wealth, tax rates, and the real interest rate — whose nominal component might differ among individuals and groups. However, it has an important advantage: it covers 35 years. For estimation purposes, 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. To avoid arbitrary interpolation of the microdata, we prefer using the annually released data only. We aggregate each wave, creating an “artificial” time series. We call this artificial because the PSID is a panel of households and we use labor market data for the household head only. When the household includes a couple, the husband is conventionally defined as the head. Therefore, women and young people are under-represented in our sample. This presumably reduces the estimated labor supply elasticity. More generally, our population is different from the US population. However, what matters is that we do have a well-defined population with well-defined individual and aggregate elasticities of labor supply. As in MaCurdy (1981), we exclude from the sample permanently disabled or retired individuals, i.e. include only those units displaying nonzero wage and labor supply data in any particular year. We report in Table 1 the means of key demographic and labor supply variables (age, sex, marital status, self-employment, hours per worker). This Table gives a rough picture of the population we are working with, as well as its dynamics.

We compared our artificial time series with real aggregate US labor data, using historic data provided by the Bureau of Labor Statistics (BLS). Figure 1 illustrates the case for average hours worked by employed individuals. The real aggregate series is smoother, while our artificial series jumps around, although the two move together very closely, up to a difference of about 100 hours, until 1984. Even after that time, the two tend to move together. The difference in levels can be easily explained by the over-representation of men, who typically work more than women. The different volatility can be explained by the changes occurring in the panel due to attrition and — most importantly — the inclusion of new households in the survey. Part of the latter is nonrandom: the PSID automatically includes as new households children that leave a family already included in the survey. In Figure 2 we summarize this same information taking logs and using first differences. Figure 3 compares variations in the log real wage rate in the PSID and in the US. All nominal values are converted into real terms using the CPI deflator provided by the BLS. Again the two series move quite closely together, apart from a scale factor reflecting the fact that the BLS series refers to wages of production and non-supervisory workers only, until 1989. We return on this in a moment. This graph suggests that 1992 (i.e. 1993 wave) is an anomalous year in the PSID, and we have no explanation for this. The best we can do is to use a dummy to control for this anomaly, and further truncate the series as a robustness check — more on this below. Finally, Figure 4 shows the co-movements of the aggregate wage rate and hours per workers in the PSID: these are roughly consistent with NBER recessions, which are located by dashed rectangles.

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9 Omitted wealth is likely to bias downward estimated elasticities, as showed by Domeij and Flodén (2006).
10 It is well known that women have a higher labor elasticity than men. See, for instance, Killingsworth and Heckman (1986).
11 The importance of this point for obtaining a reliable aggregation was stressed in private conversations by Edward Prescott and Richard Rogerson.
12 We consider the year the were collected, i.e. one year before the official date of the wave.
As we suggested above, the volatility of our series relative to the real ones is probably due to variations in the composition of the PSID relative to the US population. 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. Table 2 reports, for each year of annual release, the number of households present in the sample, as well as the variation over the previous year. After minor variations from the beginning (except 1968) until 1988, 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 1994 wave (which collected our 1993 data) also was characterized by a sizeable, although smaller, increase in sample size. Then in two successive steps (1996 and 1997 waves, i.e. our 1995 and 1996 years) sample size was drastically reduced, in an attempt to reduce the cost of the survey. The effect of these changes on the comparability with real US data is clear in Figure 5, which reports the variations in the log of employment in the PSID and in the US (Source: BLS). Employment in the PSID appears to be less volatile than in the US. The two series move together until 1988, which makes are confident the problem of mixing variations in employment with variations in sample size has no serious consequences on our estimates up to 1988. However, after this years there are clear outliers in the PSID series. We address this problem in two alternative ways. First, we construct dummy variables to control for anomalous behavior of the series in 1989, 1993, 1995 and 1996. This makes a total of 5 time dummies — recall we also use a dummy for year 1992. Second, we estimate the macro regression with a shorter series that avoids all of these troublesome years, i.e. 1967-1988. As we show below, results do not change. In fact using dummies in the longer series leads to more conservative estimates. After removing the troublesome years, variations in employment in our sample and in the US show a coefficient of correlation of 0.52. A univariate regression of variation in employment in the PSID over variation in the US yields a significant slope of 0.46 and a negligible intercept.

Finally, an obvious problem concerns individual wages. It is well known that wage reports in the PSID are affected by measurement errors (Pischke 1995). Such errors are of course washed out by aggregation, but remain a concern in the individual regression. After presenting the main results, we will perform a check by excluding self-employed individuals, who may be more likely to misreport their hourly wage — employees’ wage is more stable and is written in their pay sheet. Our results are quite robust to this check.

RESULTS

Our results are summarized in tables 3 to 6. Table 2 contains our baseline estimates of the individual and aggregate labor supply elasticities. Columns 1 and 2 report fixed-effects

13. The dummies are used in both the micro and the macro regressions, in order to preserve the exact correspondence between the two. However, the micro regression is also estimated without dummies, as we believe they are unnecessary here.
estimates of the individual elasticity not controlling and controlling for non-labor income, respectively. The first estimate yields a Frisch elasticity of 0.11. Controlling for non-labor income yields an insignificant estimate, but the \(J\) test indicates that this is not a reliable estimate. Because of the presumption that the year dummies are not important in the micro regression we re-estimated the two models without them. Columns 3 and 4 show that in fact the results are very similar. Additional results reported in Table 5 shows that our estimate of the individual elasticity is about the same when the “Heckit” estimator is used to control for self-selection into the pool of workers. This table also shows that excluding self-employed workers to minimize measurement errors in the wage rate produces a larger elasticity. However, this is still within the conventional range and — more importantly — still significantly smaller than the aggregate one. Columns 5 and 6 report estimates of the aggregate elasticity. Not controlling for mean non-labor income yields an aggregate Frisch elasticity of 0.86, which increases to about 1 but becomes insignificant when controlling for mean non-labor income. This second estimate, however, is meaningless because the rank condition is not satisfied with the instruments we use for non-labor income. Notice that the constant has a natural interpretation in terms of a time trend in hours worked. The difference in sign between the individual and the aggregate trend may be explained in terms of a slightly shorter workweek but higher employment, or simply by aging f the sample (see Table 1). Table 6 shows that discarding the part of the series that behaves anomalously (i.e. starting in 1989) instead of using dummies does not change much, and actually increases, the estimated aggregate elasticity.

These results show that aggregation alone magnifies the aggregate elasticity by several times. This is due to variation along the extensive margin, as illustrated in Table 4. This table evaluates empirically the two terms in equation (5), i.e. it splits the aggregate elasticities into the contribution due to the intensive and extensive margins, respectively. Specifically, the dependent variable in column 7 is the first difference of log mean hours worked in the economy (intensive margin), while in column 8 is the first difference of log employment (extensive margin). Of course these two terms sum to the baseline estimate of the aggregate elasticity. The estimate forcefully shows that the extensive margin explains most of the difference between micro and macro elasticities, as well as that the elasticity of the labor stock is positive as required by movements along a labor supply curve. This confirms the variance decomposition result that real wages affect much more employment/participation decisions than hours.

CONCLUSIONS

In this paper we have estimated the individual and aggregate Frisch elasticities of labor

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14 The log variation in the wage rate and non-labor income are instrumented using, respectively, 3 lags (the 2nd to the 4th) of the log wage rate and log non-labor income in levels.

15 Specifically, we estimated a random effects Probit model where the dummy variable for worker status (i.e. an individual with positive hours and wage) is regressed on age, family size, number of children, total family income, sex and self-employment status. The predicted probabilities are then used to construct the inverse Mill ratio, whose first differences is used as an additional regressor.

16 Instruments are the corresponding aggregate variables, including the lag length, like in the micro regression. While this does not necessarily imply the most efficient aggregate estimate, it preserves exact correspondence between the micro and macro regressions.
supply, using exact aggregation of the microeconomic units on which the individual estimate is based. We found that the micro elasticity is about 0.1 and the macro elasticity is much larger, i.e. near 1. As expected, the difference is explained by adjustments at the extensive margin.

We do not claim that our estimates are the right ones, although our elasticities are close to what one finds in well-established literature. Given our aggregation procedure and the limitations of PSID data, we are aware of the fact that our specification does not account for such important variables as marginal tax rates, individual wealth, and after-tax return on assets though these limitations equally apply to the individual and the aggregate estimates.

Moreover, what we estimate is a short-run elasticity only, because the scarcity of data points (in the aggregate dataset) prevent from exploiting a richer dynamics helping to disentangle between the short and the long-run response. We are also aware of the fact that our results apply to the PSID data only and do not necessarily hold for the US.

Despite these limitations, the main achievement of the present paper is in showing through a very clear empirical exercise that aggregation alone leads to a much larger elasticity because of the implied inclusion of the extensive margin. In our model this is captured by allocation of work between market and home production. Despite being a simple empirical result, we are not aware of other econometric studies indicating the relevance of adjustments on the extensive margin based on exact aggregation of the individual units. Finally, our results show that micro estimates are not always appropriate for calibrating the national economy. This does not imply resurrecting old style macroeconomics; rather, it implies that calibrating a macroeconomic model is not simply a matter of importing microeconometric evidence into an aggregate framework.
Tab. 1  
Means of demographic and labor variables

<table>
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<tr>
<th>Year</th>
<th>Age</th>
<th>Male</th>
<th>Single</th>
<th>Self-empl.</th>
<th>Hours</th>
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<td>0.13</td>
<td>0.070</td>
<td>2040.3</td>
</tr>
<tr>
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<td>0.76</td>
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<td>0.14</td>
<td>0.054</td>
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<td>0.15</td>
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<tr>
<td>1972</td>
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<td>0.77</td>
<td>0.16</td>
<td>0.055</td>
<td>2025.1</td>
</tr>
<tr>
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<td>0.77</td>
<td>0.17</td>
<td>0.054</td>
<td>2043.2</td>
</tr>
<tr>
<td>1974</td>
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<td>0.77</td>
<td>0.18</td>
<td>0.056</td>
<td>1962.2</td>
</tr>
<tr>
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<td>0.063</td>
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<tr>
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<tr>
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<td>0.77</td>
<td>0.19</td>
<td>0.063</td>
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</tr>
<tr>
<td>1978</td>
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<td>0.76</td>
<td>0.19</td>
<td>0.066</td>
<td>2022.6</td>
</tr>
<tr>
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<td>0.76</td>
<td>0.20</td>
<td>0.067</td>
<td>1994.5</td>
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<tr>
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<td>0.069</td>
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<td>0.20</td>
<td>0.078</td>
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<tr>
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<td>0.19</td>
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<td>0.101</td>
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<tr>
<td>1989</td>
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<td>0.18</td>
<td>0.100</td>
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</tr>
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<tr>
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<td>0.19</td>
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<td>0.74</td>
<td>0.18</td>
<td>0.104</td>
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</table>
Fig. 3 Variation of log average real wage.

Fig. 4 Average wage (right scale) and average hours (left scale) per worker, PSID
### Tab. 2  Variation in the composition of the PSID (annual waves)

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<tr>
<th>Year Collected</th>
<th>Sample Size</th>
<th>Variation</th>
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<tr>
<td>1967</td>
<td>4802</td>
<td></td>
</tr>
<tr>
<td>1968</td>
<td>4460</td>
<td>-7.12%</td>
</tr>
<tr>
<td>1969</td>
<td>4634</td>
<td>4.35%</td>
</tr>
<tr>
<td>1970</td>
<td>4840</td>
<td>4.00%</td>
</tr>
<tr>
<td>1971</td>
<td>5060</td>
<td>4.55%</td>
</tr>
<tr>
<td>1972</td>
<td>5235</td>
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<td>1973</td>
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<td>4.39%</td>
</tr>
<tr>
<td>1974</td>
<td>5725</td>
<td>3.77%</td>
</tr>
<tr>
<td>1975</td>
<td>5862</td>
<td>2.39%</td>
</tr>
<tr>
<td>1976</td>
<td>6007</td>
<td>2.47%</td>
</tr>
<tr>
<td>1977</td>
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</tr>
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<td>1.33%</td>
</tr>
<tr>
<td>1981</td>
<td>6742</td>
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<tr>
<td>1982</td>
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<td>1984</td>
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<tr>
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<tr>
<td><strong>1989</strong></td>
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<tr>
<td><strong>1996</strong></td>
<td><strong>6747</strong></td>
<td><strong>-20.73%</strong></td>
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Fig. 5 Variation of log employment, 1969-1988

- • Variation ln(employment) PSID
- ■ Variation ln(employment) USA
Tab. 3  Individual and aggregate Frisch elasticities.

<table>
<thead>
<tr>
<th></th>
<th>Individual</th>
<th>Aggregate</th>
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<tr>
<td></td>
<td>( \Delta \ln (\ell_{it}) )</td>
<td>( \Delta \ln (\hat{\ell}_t) )</td>
</tr>
<tr>
<td></td>
<td>1</td>
<td>2</td>
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<tr>
<td>( \Delta \ln (\text{wage}) )</td>
<td>0.11**</td>
<td>0.07</td>
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<tr>
<td></td>
<td>(0.04)</td>
<td>(0.05)</td>
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<tr>
<td>( \Delta \ln (\text{non-labor income}) )</td>
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</tr>
<tr>
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<td>-</td>
<td>(0.01)</td>
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<th>Rank condition (IV)</th>
<th>J-stat</th>
<th>p-value</th>
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<th>Individuals</th>
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* Significant at 10%; ** Significant at 5% or better

Tab. 4  Decomposition of the aggregate elasticity

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<th>( \Delta \ln (\ell_t) )</th>
<th>( \Delta \ln (\eta_t) )</th>
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</thead>
<tbody>
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<td></td>
<td>5</td>
<td>6</td>
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<td>( \Delta \ln (\text{wage}) )</td>
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<td>0.57</td>
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<td>(0.27)</td>
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<td>(0.00)</td>
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<td>Observations</td>
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### Individual Frisch elasticity: robustness checks

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<td>0.09**</td>
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<td>(0.04)</td>
<td>(0.06)</td>
<td>(0.04)</td>
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<td>no</td>
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* Significant at 10%; ** Significant at 5% or better

### Aggregate Frisch elasticity: full and shorter sample

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</table>

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


