Income Effects on Services Expenditures

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Engel curves suffer from the fact that habit or addiction effects are not taken into account on cross sections. Also, income effects may differ between social groups, and cross-section parameters may be biased relatively to time-series estimations. We propose to estimate dynamic Engel curves on individual cross-section data using a new instrumentation of past expenditures based on cohort effects and compare the influence of income changes according to static and dynamic estimates. Finally, a domestic production model allows to calculate the opportunity cost of domestic activities and to explain the difference between the U.S. and European expenditures on services. The article uses the 1979, 1984, 1989 and 1995 Insee Family budget surveys.
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**Introduction**

In Kalwij et al (2003), estimation of Engel curves allows to distribute the change of service expenditures between three components: the income effect (measured by the income elasticities applied to income changes and the change of the income distribution, measured by a Theil index), the influence of demographic variables and the Baumol effect (the variation of budget shares for services due to the relative variation of prices for services, compared to other expenditures). Increases in household total expenditures and income inequality explain only 30% of the difference between budget shares for services in France and in the U.S. (Kalwij et al., Table 10). All income effects are measured in this study using cross-section estimates of income elasticities, which are likely to be biased compared to time-series estimates ones (see Gardes et al., 2001). Indeed, consumption laws for household expenditure on services may be discussed comparing short-term and long-term parameters, cross-section and time-series, monetary and complete income parameters. Differences between these different types of parameters may exist because of different dynamic behaviours, endogeneity biases in both CS and TS dimensions, non-monetary, such as time constraints.

In order to appreciate these results, we compute the same decomposition using estimation for France on pseudo-panels, dynamic specifications and propose a new model to calculate the effect of distributional changes.
I ENGEL CURVES ESTIMATION IN THE COMPARATIVE PERSPECTIVE

1.1 SPECIFICATION

The following reduced form Engel curve is estimated:

\[ w_i t = a_i + b_i \ln(Y_t / a(p_t)) + Z_i c_i + \varepsilon_i \]  

(1)

where \( w_i t \) is the budget share of good \( i \), \( Y_t / a(p_t) \) household’s real total expenditure with \( a(p_t) \) a Stone price index, \( Z_i \) household characteristics, \( \varepsilon_i \), a stochastic term which captures measurement errors and unobserved preferences. We estimate equation (1) taking into account possible measurement errors in total expenditures using its predicted value obtained from instrumentation equation with disposable household income and a few socio-demographic characteristics.

We include the following explanatory variables:

- \( \ln(\text{Expenditures}) \)
- \( \ln(\text{Household size}) \)
- Number of persons under 6 years of age divided by household size
- Number of person over 5 and under 18 years of age divided by household size
- Number of person over 17 and under 31 years of age divided by household size
- Number of person over 30 and under 65 years of age divided by household size
- Number of person over 64 years of age divided by household size
- Age and Age squared of the head of household
- Number of employed persons in the household
- A dummy variable equal to 1 if all adults are employed, 0 otherwise
- A dummy variable equal to 1 if all adults are employed and a person under 6 years of age is present in the household, 0 otherwise
1.2 GENERAL ESTIMATION RESULTS

1) Generally, the main variable of interest (instrumented logarithm of the total expenditure) is highly significant across all equations. It will enable to obtain a good estimate of total expenditure (or budget) elasticity. Other variables significance depends highly on the kind of item. For instance for food, the logarithm of household’s size parameter is highly significant and positive. Employment status does not influence significantly the food share and age generation variables have significant but heterogeneous impact. The income effect is negative. The similar pattern is observed for alcohol and tobacco.

2) Estimating the total services expenditure gives a positive income effect (positive relationship with the total expenditure), a negative impact of demographic variables and a positive impact of employment status variables.

Thus, the service expenditures are driven essentially by the income effect associated with high family work participation strengthened by the presence of young children. The last tendency is particularly visible in the case of the dummy variable indicating that both parents work and the presence of a young child. Its parameter estimate for total service expenditure is relatively high, positive and significant (0.028 (.0007)). It means that for these families there is a significant increase in the share of budget spent on services with respect to others, which amounts to almost 10% of average services’ budget share. This result of strong dependency of service expenditure and family labour force participation is tested more precisely in Gardes-Starzec (2003) by matching FBS with Labour Force Survey and Time Use Survey.

On the other hand potential needs for services resulting only from particular demographic family situation (number of persons, age, presence of children), does not seem to have a significant effect (ceteris paribus) on services purchase.

1.3 TOTAL EXPENDITURE ELASTICITIES (BUDGET ELASTICITIES)

Total expenditure elasticities are computed using the beta coefficients and sample average budget shares for all corresponding items (table 7). The values of elasticities obtained by instrumenting the total expenditure by income and socio-demographic variables gives are almost systematically superior to those obtained when no instrumentation is used, but their hierarchy is maintained (there is a problem concerning additivity). As expected elasticities
for goods are lower than those for services. The highest elasticities are obtained for home and holiday services (above 2 when instrumenting), the lowest - below 1 - for food, alcohol, home energy, and communication services. The total service expenditure has a budget elasticity 1.08, above the average for goods but somewhat smaller than the elasticities computed in other countries.

Generally services have highest budget elasticities than goods and belong more often to the category of luxuries. However many goods and services have similar and high level of elasticities (food away, entertainment goods) or similar and low level of elasticities (communication services and home energy). The chance that a luxury is a service rather than a good is probably higher and can increase in the future but this hypothesis should be tested more precisely. We conclude that the income elasticity is perhaps under-estimated by simple Engel curve estimation on cross-sections. In sections 2 and 3, we evaluate the rather low value obtained for this elasticity by differencing the Engel curve according to the household levels of being. Second, we examine a dynamic specification of the Engel curve, then we finally compare cross-section and time-series estimates.

Table 1. Beta (log exp) coefficients and budget elasticities: Total Expenditures instrumented (IV) or non–instrumented (non IV)

<table>
<thead>
<tr>
<th>Expenditures</th>
<th>Beta coeff (non IV)</th>
<th>Budget elasticity (non IV)</th>
<th>Beta coeff (IV)</th>
<th>Budget Elasticity (IV)</th>
</tr>
</thead>
<tbody>
<tr>
<td>Food and non-alcoholic beverages</td>
<td>-0.12</td>
<td>0.55</td>
<td>-0.10</td>
<td>0.61</td>
</tr>
<tr>
<td>Alcoholic beverages and tobacco</td>
<td>-0.01</td>
<td>0.66</td>
<td>-0.02</td>
<td>0.64</td>
</tr>
<tr>
<td>Clothing and Footwear</td>
<td>0.02</td>
<td>1.24</td>
<td>0.01</td>
<td>1.14</td>
</tr>
<tr>
<td>Private Transport Goods</td>
<td>0.09</td>
<td>1.82</td>
<td>0.03</td>
<td>1.23</td>
</tr>
<tr>
<td>Furnishing and Appliances</td>
<td>-0.01</td>
<td>0.87</td>
<td>0.02</td>
<td>1.25</td>
</tr>
<tr>
<td>Entertainment Goods</td>
<td>0.03</td>
<td>1.33</td>
<td>0.03</td>
<td>1.32</td>
</tr>
<tr>
<td>Personal Goods</td>
<td>0.01</td>
<td>1.27</td>
<td>0.01</td>
<td>1.23</td>
</tr>
<tr>
<td>Home Energy</td>
<td>-0.05</td>
<td>0.39</td>
<td>-0.03</td>
<td>0.65</td>
</tr>
<tr>
<td>Food and bev. away from home</td>
<td>0.01</td>
<td>1.26</td>
<td>0.02</td>
<td>1.38</td>
</tr>
<tr>
<td>Holiday Services</td>
<td>0.02</td>
<td>1.89</td>
<td>0.03</td>
<td>2.08</td>
</tr>
<tr>
<td>Household Services</td>
<td>0.01</td>
<td>1.77</td>
<td>0.02</td>
<td>2.03</td>
</tr>
<tr>
<td>Personal Services</td>
<td>0.00</td>
<td>1.19</td>
<td>0.00</td>
<td>1.30</td>
</tr>
<tr>
<td>Public Transport Services</td>
<td>0.00</td>
<td>1.14</td>
<td>0.00</td>
<td>1.05</td>
</tr>
<tr>
<td>Private Transport Services</td>
<td>0.02</td>
<td>1.34</td>
<td>0.01</td>
<td>1.13</td>
</tr>
<tr>
<td>Communication Services</td>
<td>-0.02</td>
<td>0.54</td>
<td>0.00</td>
<td>0.86</td>
</tr>
<tr>
<td>Entertainment Services</td>
<td>0.01</td>
<td>1.91</td>
<td>0.00</td>
<td>1.82</td>
</tr>
<tr>
<td>Miscellaneous goods and services</td>
<td>-0.03</td>
<td>0.72</td>
<td>0.00</td>
<td>1.00</td>
</tr>
<tr>
<td>Total services</td>
<td>0.02</td>
<td>1.08</td>
<td>0.06</td>
<td>1.21</td>
</tr>
</tbody>
</table>

Households expenditures may depend on their relative position (social interactions) or their absolute poverty (which impose various constraints). So, it is important to evaluate the influence of changes in the income distribution. We define sub-populations characterized by their well-being situations and compare their consumption functions.

The Synthetic Index of Poverty and Richness (SIPR) defined in Gardes et al. (2000) allows classifying households according to three different criteria:

1) the poor are defined as families which have a budget share of food greater than the average in their reference population by one third. Disposable income per unit of consumption define then poor as those which have less than half the national average (conversely, smaller by one half for the rich).

2) are poor those families the total expenditure of which are smaller by one third to the average in their reference population (conversely, greater by one half for the rich).

3) are poor those families which are in the first (conversely the fourth for the rich) quartile of disposable income (in the whole population). The first criterium refers to subsistance constraints, the second to the non-satisfaction of basic needs, the third to the capability of the household to have decent living conditions through their income.

The poor are defined as those families which are poor according to all criteria (conversely the rich as those which are rich according to all criteria), the rest of the population being classified as quasi-poor (quasi-rich) if they are poor according to two criteria and not rich for the third, or in the middle class. About 5% of the population is classified as poor (or rich), 15% as quasi-poor (or quasi-rich) and 60% in the middle class (the proportion are quite the same in different countries, see Cardoso-Gardes, 1996, because of the relative definition of the criteria).

Note that the SIPR does not aim to count the poor, but rather to define homogenous populations in order to compare their behaviour. The classification as poor and as rich, as defined by the SIPR, is restrictive, so that the difference between the five sub-populations can be clearly evaluated.
Table 2. Total Expenditure-Elasticities by poverty groups (SIPR index)

<table>
<thead>
<tr>
<th></th>
<th>poor</th>
<th>q-poor</th>
<th>medium</th>
<th>q-rich</th>
<th>rich</th>
</tr>
</thead>
<tbody>
<tr>
<td>Food at home</td>
<td>0.9087</td>
<td>1.228</td>
<td>0.9328</td>
<td>0.849</td>
<td>0.7057</td>
</tr>
<tr>
<td>Services</td>
<td>1.0772</td>
<td>1.0043*</td>
<td>1.1902*</td>
<td>1.4461</td>
<td>1.2722</td>
</tr>
<tr>
<td>Goods</td>
<td>0.9228</td>
<td>0.9958</td>
<td>0.8098</td>
<td>0.554</td>
<td>0.7277</td>
</tr>
</tbody>
</table>


* not significant at 95% level

The results of total expenditures elasticities estimations by poverty groups (table 2) show relatively important differences between poor and rich categories, the income elasticity for services being one third larger for the two rich classes compared to the poor and quasi-poor. On the other hand the elasticity differences between different expenditures are much stronger for the rich and quasi rich when compared to the poor. The demand response for services can be really observed only for relatively high-income households indicating that a large part of services is a special type of luxury with a relatively high minimum expenditure needing a high minimal level of income. Thus further analysis should pay much more attention to this specific, rich household services consumption response (see below section 4).

Table 3. Quartile regression by poverty groups
Food at home and total services

<table>
<thead>
<tr>
<th>Quartiles of Services budget share</th>
<th>Food at home</th>
<th>Total services</th>
</tr>
</thead>
<tbody>
<tr>
<td>Poverty</td>
<td>0.25</td>
<td>0.25</td>
</tr>
<tr>
<td></td>
<td>0.5</td>
<td>0.5</td>
</tr>
<tr>
<td></td>
<td>0.75</td>
<td>0.75</td>
</tr>
</tbody>
</table>

<table>
<thead>
<tr>
<th>Poverty class</th>
<th>0.25</th>
<th>0.5</th>
<th>0.75</th>
</tr>
</thead>
<tbody>
<tr>
<td>Poor</td>
<td>0.903</td>
<td>0.83</td>
<td>0.808</td>
</tr>
<tr>
<td>Quasi poor</td>
<td>1.287</td>
<td>1.325</td>
<td>1.211</td>
</tr>
<tr>
<td>Medium</td>
<td>0.9648</td>
<td>0.8798</td>
<td>0.8612</td>
</tr>
<tr>
<td>Quasi rich</td>
<td>0.607</td>
<td>0.7838</td>
<td>0.9676</td>
</tr>
<tr>
<td>Rich</td>
<td>0.8093</td>
<td>0.686</td>
<td>0.7496</td>
</tr>
<tr>
<td>Quasi rich + rich</td>
<td>0.6424</td>
<td>0.6631</td>
<td>0.7532</td>
</tr>
<tr>
<td>Quasi pauvres + pauvres</td>
<td>1.0479</td>
<td>0.9318</td>
<td>0.832</td>
</tr>
<tr>
<td><strong>Total</strong></td>
<td>0.4825</td>
<td>0.4885</td>
<td>0.522</td>
</tr>
</tbody>
</table>


Note: Quartile regression is made on the dependent variable- budget share. Note that for food, the 0.25 estimation corresponds to households having the smaller food budget share among the sub-populations, for instance the poor, thus corresponding to the richer households among the poor (and the opposite for services expenditure).
Quartile regressions within the Poverty-Richness groups indicates a further heterogeneity of the income effect. Looking at elasticities computed by quartile regression (table 3) for a basic good like food compared to services we can observe quite different patterns: Food elasticities are relatively stable when shifting from small to higher levels of budget coefficients within the total population (last line), while they are diminishing significantly within the poor, quasi-poor and medium and increasing within rich and quasi rich groups. On the contrary Services elasticities are decreasing within the total population as well is in all poverty classes.

Note that that this differentiation between quartiles corresponds, for services, to higher level of well being (according to the Engel law) for higher quartiles. Similarly the opposite is observed for food expenditure so that elasticities are decreasing for both items according to households’ level of well-being. We conclude that the heterogeneity in the income effect is quite significant and should be taken into account when decomposing the influence of income, prices and demographic changes on the households expenditure on services.

So, it seems important to analyze the income effects screening the whole population to estimate on homogeneous sub-populations, since some structural (due to changes in the proportions of the sub-populations) or aggregation difficulties may bias the estimations on the whole population. Second, the change in the Food at Home and services income elasticity are less important for the rich and quasi-rich, perhaps because they have already reached their satiation level in the first quartile of the distribution for these social classes. These estimations show in general the interest of analysis by sub-populations which reveals more demand response differences, and more specifically the special case of rich families having in the French case much higher services consumption potential than other households. Finally, we conclude that the heterogeneity in the income effect is quite significant and would be taken into account when decomposing the various influences of income, prices and demographic changes on the households expenditure on services.
3 Dynamic Specification of Engel Curves

The static nature of Engel curves may be disputed, as habit or addiction effects have been proved to exist for all types of goods. We propose to estimate, first a simple auto-regressive model, adding the lagged expenditure to equation (1), then a dynamic reduced equation based on a partial adjustment of permanent income.

We define the partial adjustment on permanent income $Y'$ by the equation:

$$\ln\left(\frac{Y'_{ht}}{a(p_t)}\right) = \left[\ln\left(\frac{Y'_{h,t-1}}{a(p_{t-1})}\right) + g\right] + \beta \left[\ln\left(\frac{Y'_{ht}}{a(p_t)}\right) - \left(\ln\left(\frac{Y'_{h,t-1}}{a(p_{t-1})}\right) + g\right)\right]$$  \hspace{1cm} (2)

With $E_{t-1}(\ln(Y'_{h,t}/a(p_t))) = \ln(Y'_{h,t-1}/a(p_{t-1})) + g$, the tendancial expected income for period $t$ made one period before (g=expected rate of income change) and $\beta$ the adjustment parameter. In the linear version of equation (1), the tendancial expected income and its difference with current income (interpreted as a logarithmic conjonctural income $y''$) are substituted to income with coefficients $b_{11}$ and $b_{12}$, giving rise with equation (1')

$$w_{it} = a_i + b_{11}\ln\left(\frac{Y'_{t}}{a(p_t)}\right) + b_{12}\ln\left(\frac{Y''_{t}}{a(p_t)}\right) + Z_t . c + \varepsilon_{it}$$  \hspace{1cm} (1)

to the reduced form:

$$w_{it} = a'_{i} + (1-\beta).w_{i,t-1} + b_{i1}\beta.\ln(Y'_{t}/a(p_t)) + b_{i2}(1-\beta) d\ln Y + Z_t . c - (1-\beta) Z_{t-1} . c + \eta$$  \hspace{1cm} (3)

with $d\ln Y = \ln(Y_{t}/a(p_t)) - \ln(Y_{t-1}/a(p_{t-1}))$, $\eta$ a MA(1) error term and $Z$ the vector of price (coefficients $c$ are constrained by the additivity, homogeneity and symmetry of price effects).

As the error term is autocorrelated, an endogeneity bias may appear in the estimation of $\beta$, which can be taken into account instrumenting the lagged budget share.

---

1 We do not discuss here the dynamic relationship between income and saving, thus the partial adjustment on income applies also to total expenditure.
Estimation of dynamic models, such as the partial adjustment equation, needs panel data (and special estimation technics). We propose to estimate them on cross-sectional data using a cohort instrumentation procedure defined in Gardes (2003 a, b).2

The method consists to define, for each agent \( h \) in a cohort \( C_h \), an agent \( S(h) \) in the same survey and aged one year less and with similar observable permanent characteristics \( Z' \). Then, we correct for the cohort effect associated with these characteristics by computing for each variable of interest \( x \) its estimated value for an agent in the same cohort \( C_h \), i.e. having characteristics \( Z_h \) in the previous year. Suppose saving \( x \) depends on variables \( Z \), so that, for the first order approximation:

(i) Between two periods for individual \( h \): 
\[
(x(Z_{h,t}) - x(Z_{h,t-1})) = (Z_{h,t} - Z_{h,t-1}) \cdot \beta_{ts} + \epsilon_{h,t} - \epsilon_{h,t-1}
\]

(ii) Between \( S(h) \) and \( h \) in period \( t \): 
\[
(x(Z_{h,t}) - x(S(Z_{h,t}))) = (Z_{h,t} - Z_{S(h),t}) \cdot \beta_{cs} + \epsilon_{h,t} - \epsilon_{S(h),t}
\]

Suppose now that \( Z_{h,t-1} \) is equal to \( Z_{S(h),t} \). Saving by the similar individual \( S(h) \) in \( t \) must be corrected to be compared to saving by \( h \) in \( t+1 \) by the formula, residuals being set to zero:

\[
E_x(Z_{h,t-1}) = x(Z_{S(h),t}) + (Z_{S(h),t} - Z_{h,t}) \cdot (\beta_{ts} - \beta_{cs}) \quad (4)
\]

The coefficients \( \beta_{ts} \) can be estimated on aggregate time-series or on a panel or pseudo-panel containing at least two periods3. \( Z_{S(h),t} \) can be computed as the average on households having the same permanent characteristics as household \( h \).

Table 4 presents the estimation of a simple auto-regressive model, just adding the instrumented lagged endogenous variable to the static equation. This term is significant and the difference between short term and long-term elasticities is quite large and significant: the long-term elasticity show that services are a highly luxury consumption, while the short term is not much different from the unity.

---

2 This I.V. method is used in Gardes-Starzec (2002) to estimate addiction effects and provides excellent estimates both of addiction and habit coefficients and of the inter-temporal substitution coefficient.

3 Note that the estimation of dynamic models on time-series needs at least four periods to instrument the lagged endogenous variable when some endogeneity is suspected. Whenever the coefficients \( \beta \) are used to define the endogenous variable, they can be calibrated on another data set or estimated by convergence using the model and equation (3).
Table 4. Simple Autoregressive model on cross-section (Total Services)

<table>
<thead>
<tr>
<th>Variable</th>
<th>Parameter</th>
<th>Standard Error</th>
<th>t Value</th>
<th>Pr &gt;</th>
<th>t</th>
<th></th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Estimate</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Constant</td>
<td>-0.48128</td>
<td>0.04703</td>
<td>-10.23</td>
<td>&lt;.0001</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Log Household Size</td>
<td>-0.04293</td>
<td>0.00599</td>
<td>-7.17</td>
<td>&lt;.0001</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Fraction Age 6-17</td>
<td>0.00148</td>
<td>0.01837</td>
<td>0.08</td>
<td>0.9357</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Fraction Age 18-30</td>
<td>-0.00766</td>
<td>0.01876</td>
<td>-0.41</td>
<td>0.6828</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Fraction Age 31-64</td>
<td>-0.03247</td>
<td>0.02028</td>
<td>-1.60</td>
<td>0.1094</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Fraction Age 65-99</td>
<td>-0.03761</td>
<td>0.02141</td>
<td>-1.76</td>
<td>0.0790</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Age head of household</td>
<td>-0.00009449</td>
<td>0.00076495</td>
<td>-0.12</td>
<td>0.9017</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Age squared head of household</td>
<td>0.00000935</td>
<td>0.00000717</td>
<td>1.30</td>
<td>0.1927</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Number of Employed</td>
<td>0.00485</td>
<td>0.00362</td>
<td>1.34</td>
<td>0.1800</td>
<td></td>
<td></td>
</tr>
<tr>
<td>All adults employed</td>
<td>0.00811</td>
<td>0.00550</td>
<td>1.47</td>
<td>0.1408</td>
<td></td>
<td></td>
</tr>
<tr>
<td>All adults employed &amp; kids&lt;6</td>
<td>0.02612</td>
<td>0.00753</td>
<td>3.47</td>
<td>0.0005</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Log Expenditures (instr)</td>
<td>0.05151</td>
<td>0.00408</td>
<td>12.63</td>
<td>&lt;.0001</td>
<td></td>
<td></td>
</tr>
<tr>
<td>W(t-1)</td>
<td>0.70760</td>
<td>0.04259</td>
<td>16.61</td>
<td>&lt;.0001</td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

Reference group: education (3), localization (3), family type (5) , age group-1,(6)
Average budget share for total services= 0.257
Elasticity (Short Term) 1.20 (σ = 0.016)
Elasticity (Long Term) 1.69

The estimation of equation (3) yields directly (i.e. not with respect with the adaptive coefficient β) the permanent and transitory income elasticities (table 4).
Table 5. Dynamic model on cross-section (Dempatem classification)

<table>
<thead>
<tr>
<th>Model</th>
<th>Short Term Elasticity</th>
<th>Long Term Elasticity</th>
<th>β</th>
<th>R²</th>
</tr>
</thead>
<tbody>
<tr>
<td><strong>Static</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Food</td>
<td>0.55</td>
<td>-</td>
<td>-</td>
<td>0.25</td>
</tr>
<tr>
<td>Goods (no food)</td>
<td>0.93</td>
<td>-</td>
<td>-</td>
<td>0.04</td>
</tr>
<tr>
<td>Total Services</td>
<td>1.21</td>
<td>-</td>
<td>-</td>
<td>0.04</td>
</tr>
<tr>
<td><strong>Simple autoregressive</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Food at home</td>
<td>0.596</td>
<td>0.355</td>
<td>-</td>
<td>0.264</td>
</tr>
<tr>
<td>Goods (no food)</td>
<td>1.07</td>
<td>1.017</td>
<td>-</td>
<td>0.084</td>
</tr>
<tr>
<td>Total Services</td>
<td>1.20</td>
<td>1.69</td>
<td>-</td>
<td>0.099</td>
</tr>
<tr>
<td><strong>Dynamic Partial adjust. 1</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Food at home</td>
<td>0.771 (.069)</td>
<td>0.427 (.088)</td>
<td>0.408 (.055)</td>
<td>0.27</td>
</tr>
<tr>
<td>Goods (no food)</td>
<td>1.017 (.049)</td>
<td>0.985 (.085)</td>
<td>0.340 (.068)</td>
<td>0.09</td>
</tr>
<tr>
<td>Total Services</td>
<td>1.081 (.058)</td>
<td>1.533 (.070)</td>
<td>0.185 (.050)</td>
<td>0.10</td>
</tr>
<tr>
<td><strong>Dynamic Partial adjust. 2</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Food at home</td>
<td>0.828 (.053)</td>
<td>0.402 (.071)</td>
<td>0.299 (.0325)</td>
<td>0.27</td>
</tr>
<tr>
<td>Goods (no food)</td>
<td>1.014 (.049)</td>
<td>0.981 (.079)</td>
<td>0.299 (.0325)</td>
<td>0.09</td>
</tr>
<tr>
<td>Total Services</td>
<td>1.117 (.052)</td>
<td>1.507 (.080)</td>
<td>0.299 (.0325)</td>
<td>0.10</td>
</tr>
</tbody>
</table>


Dynamic 1: without adjustment for socio-economic variables; corrected from cohort effects; separate estimation for each expenditure

Dynamic 2: without adjustment for socio-economic variables; corrected from cohort effects; estimation by system (SUR) for the three expenditures

Standard error in parenthesis.

For total expenditures on services, the long-term elasticity is significantly greater than the short term but to less extent than in a simple autoregressive model. The adaptive coefficient is 0.30 for the estimation by Seemingly Unrelated Regression on the three expenditures, which is quite plausible. Finally, both the simple auto-regressive and the partial adjustment model yield much greater long-term elasticities than those obtained from static specification estimates. The difference between short term and long-term elasticities is even stronger for a detailed Dempatem classification (see Table A1 in Appendix A). Thus, we conclude that using static estimates of the income elasticities strongly underestimates the income effect when compared with the full dynamic specification.

Consequently, substituting static total expenditure parameter estimates (β) by the dynamic ones, in the decomposition of total effect (table 6) gives much stronger influence of income
and residual effects largely dominating demographic and employment effects. The role of the price effects becomes relatively weaker (less than a half of budget effect).

Table 6. Empirical Results on the Explanations for the change in the expenditure shares on Non-Durable Goods and Services (dynamic specification).

<table>
<thead>
<tr>
<th>Cells: %-points</th>
<th>Total Change</th>
<th>Demographics</th>
<th>Employment</th>
<th>Budget Level</th>
<th>Budget Inequality</th>
<th>Price Effects</th>
<th>Substitution &amp; Preferences</th>
</tr>
</thead>
<tbody>
<tr>
<td>FR. 1980-1995</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Non-Durable Goods (1-8)</td>
<td>-8.0</td>
<td>-1.1</td>
<td>0.2</td>
<td>-4.98 (-2.6)*</td>
<td>0.0</td>
<td>-5.2</td>
<td>3.08 (0.7)*</td>
</tr>
<tr>
<td>Services (9-20)</td>
<td>8.0</td>
<td>1.1</td>
<td>-0.2</td>
<td>4.98 (2.6)*</td>
<td>0.0</td>
<td>5.2</td>
<td>-3.08 (-0.7)*</td>
</tr>
</tbody>
</table>

Static specification results in parenthesis
4 PSEUDO-PANEL ESTIMATES OF THE INCOME ELASTICITIES:

4.1 SPECIFICATION AND ECONOMETRICS OF THE CONSUMPTION MODEL:

First differencing and within operators are common procedures employed to eliminate biases caused by persistent omitted variables (which are likely correlated with the specific effects).

Pseudo-panel estimate:

The grouping of data for pseudo-panels is made according to five age cohorts, three education levels and location (Paris/Province) for the Insee Family Expenditures Surveys (1979, 1984, 1989, 1995). The grouping of households (h,t) in the cells (H,t) gives rise to the exact aggregated model:

\[ \sum_{h \in H} \gamma_{ht} w_h^i = W^i_H = \left( \sum_h \gamma_{ht} X_{ht} \right) A^i + \gamma_{H} + \sum_h \gamma_{ht} \varepsilon_{ht}^i \]

with \( \gamma_{ht} = \frac{Y_{ht}}{\sum_{h \in H} Y_{ht}} \) under the hypothesis \( \alpha_h^i = \alpha_H^i \) for \( h \in H \) (a natural hypothesis, according to the grouping of households into a same H cell). A heteroskedasticity factor \( \delta_{ht} = \sum_{h \in H} \gamma_{ht}^2 \) arises for the residual \( \varepsilon_{ht}^i \), which is due to the change of cells sizes (as \( \gamma \equiv \frac{1}{|H|} \) if the two grouping criteria homogenize the household’s total expenditures).

Pseudo-panel data on households provide opportunities to reduce these biases, since they contain information on changes in expenditures and income for the same households. Differencing successive surveys nets out the biasing effects of unmeasured persistent characteristics. But while reducing bias due to omitted variables, differencing income data is likely to magnify another source of bias: measurement error. Altonji and Siow (1987) demonstrate the likely importance of measurement error in the context of first-difference consumption models by showing that estimates of income elasticities are several times higher when income change is instrumented than when it is not.

Deaton (1986) presents the case for using “pseudo-panel” data to estimate demand systems. He assumes that the researcher has independent cross sections with the required
expenditure and demographic information and shows how cross sections in successive years can be grouped into comparable demographic categories and then differenced to produce many of the advantages gained from differencing individual panel data.

True panel and pseudo-panel methods each offer advantages and disadvantages for handling the estimation problems inherent in expenditure models.

A first set of concerns center on measurement error. Survey reports of household income are measured with error; differencing reports of household income across waves undoubtedly increases the extent of error. Instrumental variables can be used to address the biases caused by measurement error (Altonji and Siow, 1987). Like instrumentation, aggregation in pseudo-panel data helps to reduce the biasing effects of measurement error, so we expect that the income elasticity parameters estimated with pseudo-panel data to be similar to those estimated on instrumented income using true panel data. Since measurement error is not likely to be serious in the case of variables like location, age, social category, and family composition, we confine our instrumental variables adjustments to our income and total expenditure predictors. Measurement errors in our dependent expenditure variable are included in model residuals and, unless correlated with the levels of our independent variables, should not bias the coefficient estimates.

Second, the aggregation inherent in pseudo-panel data produces a systematic heteroskedasticity. This can be corrected exactly by decomposing the data into between and within dimensions and computing the exact heteroskedasticity on both dimensions. Indeed the heteroskedastic factor depends on time. So, correcting it by GLS makes individual specific effects vary with time, thus cancelling the spectral decomposition in between and within dimensions. This can result in serious estimation errors (Gurgand, Gardes and Bolduc, 1997). The approximate correction of heteroskedasticity that we use consists in weighting each observation by a heteroskedasticity factor that is a function of but not exactly equal to cell size. Thus the LS coefficients computed on the grouped data may differ slightly on those estimated on individual data. This approximate and easily implemented correction thus consists of using GLS on the within and between dimensions with a common variance-covariance matrix computed as the between transformation of the heteroskedastic structure due to aggregation.

Third, unmeasured heterogeneity is likely to be present in both panel and pseudo-panel data. In the case of panel data the individual-specific effect for household h is $\alpha(h)$, which is assumed to be constant through time. In the case of pseudo-panel data, the individual-specific effects for a household (h) belonging to the cell (H) at period t, can be written as the
sum of two orthogonal effects: $\alpha(h,t)=\mu(H) + \upsilon(h,t)$. Note that the second component depends on time since the individuals composing the cell $H$ change through time.

The specific effect $\mu$ corresponding to the cell $H$ ($\mu(H)$) represents the influence of unknown explanatory variables $W(H)$, constant through time, for the reference group $H$, which is defined here by the cell selection criteria. $\upsilon(h,t)$ are individual specific effects containing effects of unknown explanatory variables $Z(h,t)$. In the pseudo-panel data the aggregated specific effect $\zeta(H)$ for the cell $H$ is defined as the aggregation of individual specific effects:

$$\zeta(H,t)=\sum \gamma(h,t) \alpha(h,t) = \mu(H) + \sum \gamma(h,t) \upsilon(h,t)$$

where $t$ indicates the observation period and $\gamma$ is the weight for the aggregation of $h$ within cells. Note that the aggregate but not individual specific effects depend on time.

The within and first difference operators estimated with panel data cancel the individual specific effects $\alpha(h)$. The component $\mu(H)$ is also cancelled on pseudo-panel data by the same operators, while the individual effect $\upsilon(h,t)$ may be largely eliminated by the aggregation. Thus it can be supposed that the endogeneity of the specific effect is greater on individual than on aggregated data, as aggregation cancels a part of this effect.

Therefore, with panel data the within and the first-differences operators suppress all the endogeneity biases. With pseudo-panel data the same operator suppresses the endogeneity due to $\mu$, but not that due to $\sum \gamma(h,t) \upsilon(h,t)$. For each individual this part of the residual may be smaller relatively to $\mu$, as cell homogeneity is increased. Conversely, the aggregation into cells is likely to cancel this same component $\upsilon$ across individuals, so that it is not easy to predict the effect of the aggregation on the endogeneity bias.

### 4.2 The Data:

The French family budget surveys are made each five years, the last one in 2000. They detail the pattern of private expenditures of about one thousand households. Classic problems affect the data: errors of measurement, particularly for the income variables, systematic differences with aggregate data (for instance for tobacco), change of the relative prices between the waves (over almost one year)… We use four surveys: 1979, 1984, 1989 and 1995 (the 2000 survey will be added).

The individual data are aggregated according to five cohorts, two education levels and two locations (Paris, other). Only four over the 80 cells have less than 60 households and 10 less than 100 (representing less than 1.3% of the population). The average cells size is 539.
4.3 **RESULTS:**

Estimation on a pseudo-panel containing 20 cells for four surveys is presented in Table 6.

Table 6. Total expenditure elasticities, French Family Budgets

<table>
<thead>
<tr>
<th>Commodity</th>
<th>Between</th>
<th>Within</th>
</tr>
</thead>
<tbody>
<tr>
<td>Food at home</td>
<td>0.523 (.038)</td>
<td>0.949 (.023)</td>
</tr>
<tr>
<td>Alcohol</td>
<td>0.772 (.070)</td>
<td>1.271 (.091)</td>
</tr>
<tr>
<td>Tobacco</td>
<td>0.255 (.150)</td>
<td>1.400 (.091)</td>
</tr>
<tr>
<td>Clothing</td>
<td>1.102 (.153)</td>
<td>-0.674 (.167)</td>
</tr>
<tr>
<td>Furnitures</td>
<td>0.954 (.074)</td>
<td>1.887 (.111)</td>
</tr>
<tr>
<td>Cars purchases</td>
<td>0.796 (.072)</td>
<td>1.349 (.136)</td>
</tr>
<tr>
<td>Transport services</td>
<td>2.031 (.244)</td>
<td>2.430 (.231)</td>
</tr>
<tr>
<td>Leisure services</td>
<td>1.349 (.088)</td>
<td>1.278 (.144)</td>
</tr>
<tr>
<td>Domestic services</td>
<td>1.567 (.147)</td>
<td>0.082 (.086)</td>
</tr>
<tr>
<td>Various</td>
<td>1.331 (.132)</td>
<td>1.258 (.264)</td>
</tr>
</tbody>
</table>


It appears that:

In the cross-section dimension, services are generally luxury expenditures. The ratio of cross-section elasticities between services and the corresponding durable is around 2.

Cross-section total expenditures elasticities are similar or greater than the time-series, especially in the case of services, contrary to the durables.

Within elasticities of services seem to be greater for the poor than for the rich as classified by the SIPR index as in section 2 (see for details Cardoso-Gardes, 1996b).

The third result shows that the conclusion drawn in section 2 also concerns within estimates. Nevertheless, estimation by social groups defined by the ISPR index is likely biased by selection biases, and necessitates calculating income changes for each sub-population. Therefore, we will not try to calibrate the influence of social classification on services expenditures by means of different income elasticities, a theme that will be considered in the next section from a different point of view.
5 RELATIVE INCOME EFFECTS:

5.1 THE PRICE EFFECT OF RELATIVE INCOME CHANGES

Consider an Engel curve with positions $A_1$ and $A_2$ of two households that differ in period $t$ by their incomes, $Y_1$ and $Y_2$. Whenever the cross-section income elasticity for this expenditure differs from the time-series, the increase of household $A_1$’s income from $Y_1$ to the level of $A_2$ will increase its consumption according to the time-series elasticity. Thus, its final position, $A_{1,t+1}$, differs from $A_2$. The difference between $A_1$ and $A_2$ in period $t$ corresponds to some type of relative income effect, since household $A_1$ keeps all its characteristics while obtaining a better income position in period $t+1$: the time-series change of its consumption, due to the evolution of its income through time, is not correlated to the change of all latent variables which are related to the relative position of the household in the income distribution. Therefore, the relative income effect may be though as corresponding to the difference between the cross-section and the time-series influence of income changes.

Figure 1. Cross-section and time-series differences

In order to estimate these relative income effects, consider now two income distributions which differ only by the wage of household $h$ (working during $T_h$ hours per month at wage rate $w_h$), with incomes $Y_{1h} = w_{1h} T_h$ greater than the median in the first distribution and $w_{2h} T_h = Y_{2h} > Y_{1h}$ in the second. The second distribution is more unequal than the first. Suppose now that market services can be substitutes to domestic services at a constant wage rate $w$ that is smaller than $w_{1h}$ and $w_{2h}$. In this case, the relative price $w/w_h$ of these market substitutes decreases when the wage rate $w_h$ increases. Therefore, expenditures on market services may increase both by this price effect and by an income effect when the household’s income increases. The difference between the total effect of an income increase and the income effect measures the relative income effect due to the relative price change of services with the wage rate. This difference may be used to define implicitly the change in the relative income position of household $h$. Defining the extend to which an index of
income inequality, such as the Theil index, varies when the income of one household changes, thus allows to measure the coefficient of the consumption change according to the income distribution.

Figure 2: Changes in the income distribution

As an example, suppose that the time-series income elasticity is 1.6 for this service, and the price elasticity –0.8 (according to Frisch’s hypothesis, see Selvanathan, 1993, chap. 6). Income changes for the sole household h by 20% between two periods, which changes the relative price of the service by –16.7%. If the service expenditure amounts in the first period to 100, with h’s income at 2000, the pure income effect amounts to 32 and the price effect due to the relative income variation, to 13. Therefore, the total income elasticity, as measured on cross-section, is 2.25, while the time-series income elasticity, corresponding to the pure income effect, is only 1.6. The relative income elasticity can be estimated as the ratio of the change of consumption due to the relative price increase, to the income change: 0.13/0.2 = +0.65.

5.2 THE INCOME ELASTICITY OF THE COMPLETE PRICES: THE CASE OF FOOD AT HOME, FOOD AWAY AND DOMESTIC SERVICES EXPENDITURES:

Another type of substitution exists between consumption activities that can be obtained, either by acquiring directly market services, or purchasing goods in order to transform them by domestic production. Such is the case with food at home, consumed through domestic production using time and other on monetary inputs, compared to restaurant services. The increase of the household income due to a wage rate increase, also increases the opportunity cost for time spent on any domestic activity, which elevates the complete price of consumption activities requiring time, such as food at home. This imparts a negative effect of income on this domestic activity, facilitating the substitution towards market services.
This effect appears only if the monetary price of the market substitute does not increase with the household income: thus, the effect disappears whenever the income change concerns all households, for instance when the wage rate increases by the same amount, between two periods, for the whole population. On the contrary, the substitution between food at home and food away is likely when comparing the rich and the poor in a cross-section. Thus, the difference between cross-section and time-series income effects affords some information on the relative income effect.

Defining the consumption change due to the endogenous evolution of the complete price, as the difference between the cross-section effect of income changes (measured by the coefficient \( \beta_{i}^{cs} \)) and the time-series (not biased) one (\( \beta_{i}^{ts} \)), one obtains the equation for the change in expenditure on good \( i \):

\[
\ln C_i = \beta_{i}^{ts}.d\ln Y + \gamma_i . d\ln \pi_i = \beta_{i}^{cs}.d\ln Y
\]

which allows to calculate the relative income elasticity of the complete price \( \pi_i \) under Frisch hypothesis calibrating the own price elasticity to half of the (time-series) income elasticity:

\[
E_{\pi_i/Y} = 2(1 - \beta_{i}^{cs} / \beta_{i}^{ts}).
\]

This formula gives the following income elasticities of the complete price for food at home and food away: 1.00 and –3.13 for the US (PSID data, see Gardes, 2003), 0.95 and –0.61 for France. For all services expenditures, estimates of Table 5 give \( E_{\pi_i/Y} = -0.64 \) and –1.8 for Domestic Services.

The high value of these elasticities implies that the demand for services are likely highly influenced by changes on the income distribution, as soon as the own price elasticity for services is non null. Thus, services expenditures may have been lowered in European countries by an increase of the low wages since the sixties. We calibrate these effects in the subsequent sections.

5.3 **Substitution effects between domestic activities and market substitutes:**

Suppose now that the complete price writes:
\[ \pi_i = p_{m_i} + t_i \omega \]  

(6)

with \( p_{m_i} \) the monetary price for goods used in consumption activity \( i \), \( t_i \) the time spent to consumption \( i \) and \( \omega \) the opportunity cost for time. The later is supposed to be proportional to the ratio of the minimum wage rate (for market services) to the household wage rate, so that (independently of any tax correction and for constant hours of work) the income elasticity of the opportunity cost is –1.

Formula (6) gives, whenever the monetary price does not depend on the household’s income:

\[ \frac{\partial \pi_i}{\partial Y} = t_i \left( \frac{\partial \omega}{\partial Y} \right) = E_{\pi_i/Y} \cdot \left( \frac{\pi_i}{Y} \right) = \frac{E_{\pi_i/Y}}{1 - E_{\pi_i/Y}} \cdot \left( \frac{p_{m_i} + t_i \omega}{Y} \right), \]  

so that:

\[ t_i = \left[ \frac{E_{\pi_i/Y}}{1 - E_{\pi_i/Y}} \right] \cdot \left( \frac{p_{m_i}}{\omega} \right) \]  

(7)

Formula (7) indicates the minimum time for activity \( i \) such that the household would prefer too buy the market service, instead of accomplishing himself the domestic production corresponding to consumption \( i \):

\[ t_i \succ \left[ \frac{E_{\pi_i/Y}}{1 - E_{\pi_i/Y}} \right] \cdot \left( \frac{p_{m_i}}{\omega} \right) \Leftrightarrow \text{market substitute} \succ \text{domestic production} \]  

(8)

For example, if the monetary price for food at home in France is 5 euros per person, and the opportunity cost for one hour of domestic activity 10 euros, the maximum duration of the domestic food at home activity must be 4.5 hours. For \( \omega = 50 \) euros, this maximum duration is 1.9 hours. For the U.S., with \( \frac{E_{\pi_i/Y}}{1 - E_{\pi_i/Y}} \) set at 15, \( p_{m_i} = 3 \) $ and \( \omega = 8 \) $ (respectively 50 $), the limit \( t_i \) is 5.6 hours (respectively 0.9). These figures seem plausible.

\[ ^4 \text{Note that they are sensible to the proximity of } E_{\pi_i/Y} \text{ to the unity.} \]
The limit duration \( t_i \) increases with the ratio of monetary price to the opportunity cost of time, and also with the income elasticity of the complete price for activity \( i \). As the latter is positively related to the income inequality in the population, the decision rule (8) indicates that *market services expenditures may be greater for the more unequal countries* (at least when income inequality is measured by the ratio of maximum to minimum income, for instance by the inter-decile ratio).

We first calibrate the elasticity of the budget share over an income inequality index. Then, we calculate the difference in the Theil index between France and the U.S. and compute by three different methods the resulting difference in the budget share for services.

### 5.4 Changes in the Income Distribution and Opportunity Costs:

Suppose that the income distribution differ between two countries by the sole higher income class, with populations \( \Pi_1 \) (for instance France) and \( \Pi_2 \) (U.S.), and \( Y_{Dk} \) the \( k \)th income decile:

\[
(\Pi_1) \ Y_{D9} = 1.5 \ Y_{D5} = 3 \ Y_{D1} \\
(\Pi_2) \ Y_{D9} = 2 \ Y_{D5} = 4 \ Y_{D1}
\]

The Theil index measuring the income inequality increases by \( 0.1 \ln 1.33 \approx 2.9\% \) between countries 1 and 2 (supposing that the distributions of income over the ninth decile are proportional in the two countries). Therefore, because of the greater inequality in country 2, the opportunity cost for domestic production for the rich in terms of the minimum wage rate is smaller by 25% in country 2 (1/4 compared to 1/3).

The own price elasticity can be calibrated, under Frisch hypothesis, as minus half of the income elasticity, around \(-0.75\) for all services (according to the income elasticity for all services in Table 5, 6 and in Cardoso-Gardes, 1996a). Suppose also that the budget share for all services is 0.5 for the rich. Then, the change in the budget share is equal to \(-0.75 \cdot 0.5 \cdot -0.25 = +0.094\) for 10% of the population. This aggregates to \(+3.1\%\) for the whole population if the rich are supposed to consume 30% (respectively a proportion \( p \)) of the total expenditure in services. Thus, the marginal propensity of the aggregate budget share over the Theil index \( \Gamma \) is \( \partial \ w_{serv}/\partial \Gamma = 0.031/0.029=1.07 \) (respectively \( 0.094p/0.029 \)). Therefore, the influence of income distribution is much under-estimated in an Oaxaca decomposition of budget shares (see Table 6 and Kalwij, tables 10 and 12). Moreover, the Theil index is calculated, in
these estimations, for the whole population instead of a comparison between very high and very low-income classes.

5.5 INCOME INEQUALITY IN FRANCE VS THE U.S. AND MARKET SERVICES EXPENDITURES:

We use three different methods to assess the influence of the income distribution on the budget share for services. In the first calculation, we calibrate the differential of the budget share as concerns the Theil index $\partial w_{serv}/\partial \Gamma$ as in section 4.4, considering the changes in the relative price for services due to income inequality, and compute the change in the budget share corresponding to the observed change in the Theil index. Between countries $i$ and $j$:

$$dw_{serv} = \left(\partial w_{serv}/\partial \Gamma\right)_{i}d\Gamma_{ij}.$$  

In the second method, we calculate directly the change in the relative price corresponding to the ratios of high and low income in the two countries, and use the price elasticity $e_{ps}$ to compute the change in services expenditures:

$$dw_{serv} = e_{ps}\ln(p_{i}/p_{j}).$$

Finally, the third method uses the difference in the Theil indexes corresponding to the ratio of high and low incomes in each country, then compute the change in complete prices for services using then income elasticity of complete prices calculated in section 4.2:

$$d\pi_{i}/\pi_{i} = E_{\pi/Y}dY/Y.$$  

Then, the change in the budget share depends on the price effect of complete prices on the budget share:

$$dw_{serv} = w_{serv}E_{\pi_{serv}/\pi_{i}}(d\pi_{i}/\pi_{i}).$$

Thus, the first method computes directly the effect of a change in then Theil index on the budget share, while the second estimates the price effect due to the difference in the income distributions, and the third estimates the effect of the income distribution on the relative price using the relation between relative income and the complete price for services. Note that these three method are not completely independent, especially the first and the second.

First, we compare only the two extreme decile incomes in the U.S. and in France (Table 3 in Kalwij et al.). The Theil index $\Gamma$ is greater in the U.S. by 0.019, which imparts a greater budget share for services of 0.020 (for $\partial w_{serv}/\partial \Gamma=1.07$). The income inequality between higher and lower income classes, which influences the opportunity cost of domestic production, concerns also households pertaining to the 8th or the 7th decile classes. Also, the Theil index of 0.019 does not take into account the distribution of income above the 9th decile. So, the influence of the Theil index difference due to the comparison between the extreme income classes may be rather a figure that is very similar to the actual difference observed, around 7.5% (Table 3 in Kalwij et al.).
By a more direct method based directly on the price effect, we could calculate the difference between the budget shares in the two countries by the formula: \( \text{d}w_{\text{serv}} = -0.75 \cdot \ln(\pi_i^F/\pi_i^U.S) = -0.75 \cdot \ln(3.87/4.76) \cdot p = +0.078 \) for \( p=0.5 \) and \( 0.047 \) for \( p=0.3 \).

The third method consists to estimate the change in the relative income of some group of households corresponding to a change of the Theil index, then computing the change in the complete price corresponding to the relative income elasticity of this price, and using the own price elasticity of services to calculate the change in their budget shares. The change in Theil index writes, supposing \( x_i = x + \text{dx}_i \):

\[
\Delta \Gamma = \sum_i \ln\left(\frac{x+\text{dx}_i}{x}\right) \approx \sum_i \frac{\text{dx}_i}{x} \approx \frac{Y_9 - Y_1}{Y_1} = \ln(4.76/3.87) = 0.207
\]

Thus, with \( E_{\pi_i/Y} = -0.64 \) for all services (respectively \(-1.8 \) for domestic services) as previously estimated, we obtain:

\[
\text{d}w_{\text{serv}} = \frac{\partial w_{\text{serv}}}{\partial \ln \pi_i} \cdot E_{\pi_i/Y} \cdot \frac{dY_r}{Y_r} = w_{\text{serv}} \cdot E_{\text{serv}/\pi} \cdot E_{\pi_i/Y} \cdot \frac{dY_r}{Y_r}
\]

\[
\Rightarrow \text{d}w_{\text{serv}} = 0.5 \cdot (-0.75) \cdot (-0.64) \cdot 0.207 = +0.0497 \text{ and } \text{d}w_{\text{dom,serv}} = 0.025 \cdot (-0.8) \cdot (-1.8) \cdot 0.23 = 0.0075.
\]

Finally, income inequality explains more than a half of the difference (amounting to 7.5%) between the French and the U.S. budget shares for services. We conclude that the difference between the American and European budget shares for services may be correctly accounted using the different ratios of the maximum and minimum incomes in these countries, these ratios being proxies for the relative price of services for the richer households which are the more likely consumers of domestic services in industrial countries\(^5\).

\[^5\] A calculation of the effects of the income distribution in Europe and the U.S. necessitates precise estimates of the income elasticity of complete prices for all types of services expenditures, an estimate of the difference between the ratio of maximum and minimum wages in the two countries (or the difference in the Theil index due to the extreme income classes), and the proportions of services for each income class.
Conclusion

Using dynamic specification of the consumption model with age cohort instrumentation seems to be an efficient method to obtain plausible estimates of long and short term elasticities based on cross-section data. Obtained long term total expenditure elasticities for services are significantly greater than short term ones, both for estimations using individual data or aggregate time-series. This is very important for income effects calibration. When computed on cross-section the under-estimation of expenditure elasticities can be as high as 40%.

Consequently, the computed dynamic parameters of consumption function increase significantly the role of income effects in the decomposition analysis of the inter-temporal change in budget shares between goods and services when compared with the static frame. In particular the price effects play much smaller relative role in the observed total change in budget shares.

The deeper analysis of income effects shows that the difference between the American and European budget shares for services is well explained by relative income effect through different opportunity costs for market services in the two countries. This result has an important policy implication: the increase of domestic services consumption needs essentially the lower opportunity cost for relatively rich families.

More results of dynamic elasticities estimates, by detailed sub-populations, can be obtained using the proposed dynamic consumption model estimation on cross-section data. The further research, using time-series of cross-section (pseudo-panel) data, will give more evidence about relationships between static and dynamic elasticities.
REFERENCES


### APPENDIX:

Table A1. Short term and long term elasticities from partial adjustment model

<table>
<thead>
<tr>
<th>Expenditure</th>
<th>Short term elasticity</th>
<th>Long term elasticity</th>
<th>beta estimated</th>
<th>Short term elasticity</th>
<th>Long term elasticity</th>
<th>beta fixed=0.30</th>
<th>Av. budget Shares</th>
</tr>
</thead>
<tbody>
<tr>
<td>1. Food and non-alcoholic beverages</td>
<td>0.77135</td>
<td>0.42667</td>
<td>0.40783</td>
<td>0.82673</td>
<td>0.40190</td>
<td>0.2458057</td>
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</tr>
<tr>
<td>2. Alcoholic beverages and tobacco</td>
<td>0.94782</td>
<td>-0.91981</td>
<td>0.15021</td>
<td>0.86073</td>
<td>-0.88915</td>
<td>0.0392308</td>
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</tr>
<tr>
<td>3. Clothing and Footwear</td>
<td>1.04621</td>
<td>0.42667</td>
<td>0.28498</td>
<td>1.05002</td>
<td>0.40190</td>
<td>0.0663557</td>
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</tr>
<tr>
<td>4. Private Transport Goods</td>
<td>0.95721</td>
<td>0.87517</td>
<td>0.23906</td>
<td>0.97055</td>
<td>0.89189</td>
<td>0.1066406</td>
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<tr>
<td>5. Furnishing and Appliances</td>
<td>1.05567</td>
<td>0.36073</td>
<td>0.22043</td>
<td>1.07416</td>
<td>0.35771</td>
<td>0.0715124</td>
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<tr>
<td>6. Entertainment Goods</td>
<td>1.10911</td>
<td>0.53667</td>
<td>0.25983</td>
<td>1.12350</td>
<td>0.53478</td>
<td>0.0762236</td>
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<tr>
<td>7. Personal Goods</td>
<td>1.06553</td>
<td>-1.65855</td>
<td>0.18418</td>
<td>1.12681</td>
<td>-1.70662</td>
<td>0.0201892</td>
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<tr>
<td>8. Home Energy</td>
<td>0.96503</td>
<td>-0.33451</td>
<td>0.22359</td>
<td>0.92013</td>
<td>-0.32585</td>
<td>0.0712655</td>
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<tr>
<td>9. Food and beverages away from home</td>
<td>1.09959</td>
<td>3.93790</td>
<td>0.17036</td>
<td>1.15215</td>
<td>3.93724</td>
<td>0.0456740</td>
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<tr>
<td>10. Holiday Services</td>
<td>1.40237</td>
<td>6.84263</td>
<td>0.25032</td>
<td>1.46859</td>
<td>6.81459</td>
<td>0.0237067</td>
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<tr>
<td>12. Household Services</td>
<td>1.85925</td>
<td>8.53185</td>
<td>0.41833</td>
<td>1.69016</td>
<td>8.76469</td>
<td>0.0160930</td>
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<tr>
<td>14. Personal Services</td>
<td>1.14716</td>
<td>10.6953</td>
<td>0.19486</td>
<td>1.15091</td>
<td>10.7288</td>
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<td>15. Public Transport Services</td>
<td>0.98506</td>
<td>6.22082</td>
<td>0.095598</td>
<td>1.17144</td>
<td>6.01218</td>
<td>0.0126436</td>
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<tr>
<td>16. Private Transport Services</td>
<td>0.92650</td>
<td>3.70915</td>
<td>0.31040</td>
<td>0.92579</td>
<td>3.70581</td>
<td>0.0215261</td>
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<tr>
<td>17. Communication Services</td>
<td>1.08832</td>
<td>4.34332</td>
<td>0.12627</td>
<td>1.06150</td>
<td>4.28033</td>
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<tr>
<td>19. Entertainment Services</td>
<td>1.39543</td>
<td>0.42667</td>
<td>0.31304</td>
<td>1.37831</td>
<td>0.40190</td>
<td>0.0054109</td>
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<tr>
<td>20. Miscellaneous goods and services</td>
<td>1.00529</td>
<td>2.14704</td>
<td>0.26316</td>
<td>1.00394</td>
<td>2.14395</td>
<td>0.0967719</td>
<td></td>
</tr>
</tbody>
</table>


Specification: equation (3), dynamic 1, Table 5
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Ronald Schettkat and Lara Yocarini (Jan. 2003)

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Working Papers: (See list below)
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Working papers are downloadable at http://www.uva-aias.net/lower.asp?id=194


2. Laura Blow, Household Expenditures Patterns in the UK

3. Adriaan Kalwij & Wiemer Salverda, Changing Household Demand Patterns in the Netherlands: Some Explanations

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