The Representative Agent Hypothesis: An Empirical Test*

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Abstract

This paper empirically tests the validity of using only *mean* income as a representative variable for the whole population in the aggregate consumption relation and of assuming time-invariance of the coefficients in this relation, as done in macromodels. We use a statistical distributional approach of aggregation to test these properties on the UK-Family Expenditure Survey [1974-1993]. It is observed that the time-invariance assumption is rejected in most cases. A bootstrap test also suggests that in addition to *mean* income, the *dispersion* of income matters significantly for the commodity group *services* in several years and for *total nondurable* in some years, thus invalidating the *representative agent hypothesis*.

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1 Introduction

Consumption accounts for a major share in GDP, and therefore, the behaviour of aggregate consumption is the most studied aggregate relation. There exist several studies in the literature which use the two most important theories, *life cycle hypothesis* and *permanent income hypothesis*, to analyse aggregate consumption.[1] Yet, it is not obvious why these theories, formulated for individual households, should be applicable to aggregate data. In order to transit from microeconomic behaviour to aggregate (macro) behaviour, macroeconomists often opt for the *representative agent hypothesis*. They treat aggregate behaviour as if it is the outcome of a single representative consumer. The economic aggregates are treated as though they necessarily obey the optimizing choices of a single decision maker. As Hartley (1997) puts it: "To fit the requirements of a new classical *representative agent* model, the aggregate curve must be exactly the same as the rigorously derived individual curve". Therefore, the aggregate consumption relation is modelled by substituting 'representative variables' (often the *mean*) for each of the household-specific explanatory variables in the individual behavioural relation. A number of highly restrictive and far-fetched conditions for any real world economy have to be satisfied to ensure the replication of the microeconomic relation as an aggregate consumption relation. Besides, the *representative agent hypothesis* is unjustified and misleading. This is mainly because (i) there is no justification for assuming that the aggregate of all optimizing agents act like an optimizing individual; (ii) the reaction of the *representative agent* to changes in the parameters of the model may not be the same as the aggregate reaction of individuals. Even if the *representative agent* opts for a specific alternative all individuals may choose another alternative.[2] Hence the *representative agent hypothesis* neglects the possibility that the aggregation process itself can generate properties and give insights into the structure of the aggregate consumption relation.

In this paper we adopt the statistical distributional approach of aggregation by Hildenbrand (1998) and Hildenbrand & Kneip (1999, 2002) [henceforth described as HK] to test whether the fiction of a *representative* household is an acceptable approximation or not. To model the change in aggregate consumption HK start with a consumption relation at the micro-level, which may be derived from intertemporal utility maximization without the necessity to specify a micro-relation, particularly the utility function. Then HK explicitly model the aggregation process over a large and heterogeneous population. They achieve a relation for the change in aggregate consumption expenditure which is quite different from the relation at the individual household level. The heterogeneity of the population, which is neglected in the *representative*

[1] See Deaton (1992) and Attanasio (1999) for a detailed survey on aggregate consumption.
[2] See Kirman (1992) on this issue.

agent hypothesis, influences the form of the aggregate consumption relation. In this relation not only the *mean*, which is a commonly used 'representative variable', but also the *dispersion* of income is present. In the empirical demand literature, there also exist models where second and higher moments of the income distribution appear in the aggregate relation. Yet, all these models require a specific nonlinear parametric form at the individual level to account for the individual income heterogeneity [see Blundell and Stoker (2000), Stoker (1993) for a detailed discussion of these models]. The specification of the form of the micro-relation is not needed for HK's distributional aggregation approach. Only a complete set of explanatory variables at the micro-level is needed. This can be derived by applying intertemporal utility maximization so that other variables, such as changes in prices, in wealth, in expected future interest rates, and in future uncertain labour income also appear in the aggregate relation.

In this paper we concentrate on the effect of income, which is observed in cross-section data. The other effects, appearing in the aggregate relation, are either unobserved in the data set, we use (i.e., wealth[3]), or unobservable in principle. The coefficients relating to the two parameters of the income distribution, i.e., *mean* and *dispersion*, appearing in the aggregate relation, can be estimated from cross-section data *independently* from each other, *independently* from the coefficients of other effects, and *separately* for each year. Therefore, this approach offers some advantages over the classical aggregate time series analysis of consumption which is based on the *representative agent hypothesis* :

(i) It is possible to test directly the validity of the *representative agent hypothesis* by testing the significance of the coefficient relating to the dispersion of the income distribution. The test is conclusive against *representative agent* if one finds significance of this coefficient. Yet, it is not necessarily conclusive in favour of *representative agent* if the test suggests insignificance of the coefficient. This is due to the fact that we need not specify the functional form of the micro-relation and we consider only one explanatory variable, not the full set.

(ii) The assumption of time invariance of the coefficients, implicitly present in *representative agent* models, can be easily verified by performing stationarity tests of these coefficients, as these coefficients are estimated *separately* for each year.

(iii) The specification of an individual behavioural relation is not required. Therefore, the imposition of any ad hoc structure on unobservable variables can be avoided.

[3]In the Family Expenditure Survey only a part of the asset variable, i.e., property income, is given.

The goal of this paper is to explore the validity of the *representative agent hypothesis* by performing the tests described in (i) and (ii). The paper is organized as follows: in section 2 we briefly describe the methodology which we use for the comparison with the *representative agent* model. Section 3 depicts the data and the results, and finally, in section 4 the conclusions are drawn.

2 Methodology

In their distributional approach of aggregation HK (2002) take into account the heterogeneity of the population in consumption expenditure, income, and household attributes such as age, household composition, etc. For a large and heterogeneous population H_t aggregate consumption expenditure is defined as

$$C_t := \frac{1}{\#H_t} \sum_{h \in H_t} c_t^h \tag{1}$$

where c_t^h denotes the consumption expenditure of household h in period t, and $\#H_t$ denotes the number of households in the population H_t.

Let $\bar{c}_t^a(y)$ denote the mean consumption expenditure of the subpopulation $H_t(y, a)$ where y denotes current labour income and a denotes a profile of observed household attributes such as age, employment status etc. Then the function $y \to \bar{c}_t^a(y)$ is the cross-section Engel curve of the subpopulation $H_t(a)$. Hence, aggregate consumption expenditure C_t can be expressed as

$$C_t = \int \bar{c}_t^a(y) distr(y_t^h, a_t^h | H_t) \tag{2}$$

where $distr(y_t^h, a_t^h | H_t)$ is the joint distribution of household income y_t^h and attribute profile a_t^h across the population H_t.

It is a well known empirical fact that the Engel curve $\bar{c}_t^a(\cdot)$ changes over time. For modelling the change in C_t over time, one has to model the change of the Engel curve $\bar{c}_t^a(\cdot)$ and of the joint distribution $distr(y_t^h, a_t^h | H_t)$ over time. In HK it is shown how the change of the Engel curve can be modelled (using the assumption of *structural stability*). In this case the Engel curve can be parametrized in the following way:

$$\bar{c}_t^a = f(y, a, \theta_t(a)).$$

Regarding the evolution of the joint distribtion $distr(y_t^h, a_t^h | H_t)$ over time HK use two hypotheses. The first hypothesis says that the standardized log-current income distribution changes very slowly (local time-invariance). The second hypothesis describes the modelling of the attribute distribution in a similar spirit.

Hence, the relative change in aggregate consumption expenditure[4] can be decomposed into several effects by a Taylor series expansion:

$$\begin{aligned} \frac{C_t - C_{t-1}}{C_{t-1}} &= \beta_{t-1}(m_t - m_{t-1}) + \gamma_{t-1}(\frac{\sigma_t - \sigma_{t-1}}{\sigma_{t-1}}) \\ &+ \text{ remainder term} \end{aligned} \tag{3}$$

where m_t and σ_t denote the *mean* and the *standard deviation* of the log-current labour income distribution, respectively. The remainder term captures the change in the parameter $\theta_t(a)$ which in turn is determined by the modelling methodology of the microunits' behaviour. For example, if individual consumption expenditure c_t^h is modelled by expected intertemporal utility maximization under the life cycle budget constraint with stochastic labour income and no credit restriction, the remainder term includes changes in wealth, in the expected value of future interest rates, in expected future labour income, and in preferences as well as second order terms (from the Taylor series expansion). Therefore, the remainder term should not be considered as negligible or as an i.i.d error term. The first two terms in the aggregate relation (3) capture the effect of the changing distribution of real current labour income.

Using the definitions of β_t and γ_t given in HK one can define the corresponding coefficients for the subpopulation H_t^a as

$$\beta_t^a := \frac{\sum_{h \in H_t^a} \partial_x \bar{c}_t^a(x_t^h)}{\tilde{C}_t^a} \tag{4}$$

and

$$\gamma_t^a := \frac{\sum_{h \in H_t^a} (x_t^h - m_t^a) \partial_x \bar{c}_t^a(x_t^h)}{\tilde{C}_t^a} \tag{5}$$

where $x_t^h = \log y_t^h$, $\tilde{C}_t^a := \sum_{h \in H_t^a} c_t^h$, and m_t^a is the mean of x_t^h in the subpopulation H_t^a. Therefore, the values of β_t and γ_t for the whole population are obtained as

$$\beta_t = \frac{\sum_a \tilde{C}_t^a \beta_t^a}{\sum_a \tilde{C}_t^a} \tag{6}$$

and

$$\gamma_t = \frac{\sum_a \tilde{C}_t^a (\gamma_t^a + (m_t^a - m_t)\beta_t^a)}{\sum_a \tilde{C}_t^a}. \tag{7}$$

It can be seen from (4) and (5) that β_t and γ_t depend only on the average derivative of the cross-section Engel curve of the subpopulation H_t^a. It follows from the hypothesis of local-time invariance of the distribution of log-current income x_t^h that [for proof see HK (2002)]

$$\beta_{t-1}(m_t - m_{t-1}) + \gamma_{t-1}(\frac{\sigma_t - \sigma_{t-1}}{\sigma_{t-1}}) = \beta_{t-1}(\frac{Y_t - Y_{t-1}}{Y_{t-1}}) + \bar{\gamma}_{t-1}(\frac{\sigma_t - \sigma_{t-1}}{\sigma_{t-1}}) \tag{8}$$

[4]Note that all variables considered, i.e., consumption expenditure and income are deflated by the general price index.

where Y_t is the mean of the curent income distribution, i.e., $Y_t = \frac{1}{n}\sum_{h=1}^{n} y_t^h$, where n is the number of households[5]. The coefficient $\bar{\gamma}_t$ is defined as

$$\bar{\gamma}_t := \gamma_t - \frac{\beta_t}{Y_t} \sum_{h \in H_t} (x_t^h - m_t)$$

If we recall the definition of *elasticity* it is clear from (3) and (8) that β_t, γ_t, and $\bar{\gamma}_t$ can be interpreted as elasticities. The coefficient β_t is the elasticity with respect to mean current income, which means that it can be regarded as an aggregate income elasticity of consumption expenditure. The coefficients γ_t and $\bar{\gamma}_t$ are elasticities with respect to the *dispersion* σ_t of log-current income under the ceteris paribus conditions of a constant median income and constant mean income, respectively. This income dispersion elasticity is a new concept. We use the nonparametric direct average derivative estimator [DADE] (see Stoker [1991]) to estimate $\hat{\beta}_t$, $\hat{\gamma}_t$, and $\hat{\bar{\gamma}}_t$.[6]

In the *representative agent* model one substitutes 'representative variables' (*mean*) for the explanatory variables present in the behavioural relation at the household level and considers this as the aggregate relation. Then the effect of the change in current labour income is represented by a single term involving only the change in the 'representative variable' (*mean* current labour income) Y_t, e.g., $\alpha(\frac{Y_t - Y_{t-1}}{Y_{t-1}})$. In this framework the coefficient α depends on the partial derivative $\partial_Y C(Y_t, \cdots)$, which is unknown due to the presence of unobservable variables, e.g., expectations, in the aggregate relation $C(\cdot)$. The usual practice is the substitution of these unobservables by proxies. Additionally, the coefficient α has to be estimated from time-series data with the implicit assumption of time-invariance.

Let us emphasize again two important advantages of the cross-sectional approach of aggregation, mentioned in the introduction which allow us to use this approach to test for the validity of the *representative agent hypothesis*. First, the coefficients β_t, γ_t, and $\bar{\gamma}_t$ are estimated for each time-period t, i.e., every year in our analysis, *separately*, which means that they can be time-varying. This is in contrast to the approach of macromodels, where the coefficients are implicitly assumed to be time-invariant. The second advantage is the possibility of estimating the effect of the changing income distribution *independently* of the other effects. From the definitions of β_t^a and γ_t^a in (4) and (5) it follows that these coefficients are estimated separately from cross-section data in each period. Therefore, it is not necessary to make any statistical regularity assumption on the remainder term consisting of expectational variables. Additionally, the problem of collinearity can be avoided. Above all, the question whether only the *mean* as a 'representative variable' for the whole population

[5]Under the assumption that the log-current income distribution is symmetric, m_t can be regarded as the median.

[6]See Chakrabarty and Schmalenbach (2002) for a more detailed description of the estimation procedure.

is sufficient or not, can be answered by testing the significance of the coefficient relating to the *dispersion* of income.

We use a bootstrap procedure to conduct a hypothesis test of the significance of the coefficients. In the case of $\bar{\gamma}$ we test the null hypothesis $H_0 : \bar{\gamma} = \bar{\gamma}_0 = 0$ against the alternative that $H_1 : \bar{\gamma} \neq 0$. A rejection of H_0 implies the significance of the dispersion effect. In that case the R.H.S. of equation (8) cannot be reduced to $\beta_{t-1}(\frac{Y_t - Y_{t-1}}{Y_{t-1}})$, which implies the rejection of the *representative agent hypothesis*. We consider the bootstrap distribution of $\hat{\bar{\gamma}}^* - \hat{\bar{\gamma}}$ [see Härdle & Hart (1992) and Hall & Wilson (1991)], where $\hat{\bar{\gamma}}^*$ is the value of $\hat{\bar{\gamma}}$ computed for a resample drawn from the original sample with replacement. As compared to the usual procedure of testing H_0 against H_1 based on the difference $\hat{\bar{\gamma}}^* - \bar{\gamma}_0$, the test based on the resampling of $\hat{\bar{\gamma}}^* - \hat{\bar{\gamma}}$ increases the power of the test significantly.[7] Therefore, for a test at the 5% level we compute a number b such that

$$Pr(|\hat{\bar{\gamma}}^* - \hat{\bar{\gamma}}| > b) = .05$$

Then we reject the *representative agent hypothesis*, i.e., H_0 in favour of H_1 if the absolute value of $\hat{\bar{\gamma}}$ is greater than b, i.e., $|\hat{\bar{\gamma}}| > b$. This is done similarly for the coefficients β_t and γ_t.

In order to test for the time invariance of the coefficients β_t, γ_t, and $\bar{\gamma}_t$ we perform some parametric and nonparametric tests as well as the Dickey-Fuller test for nonstationarity. In order to apply the parametric and nonparametric tests we divide the total time periods into two subperiods and then conduct these tests [details are given in the result section].[8]

3 Data and Results

We use data from the UK Family Expenditure Survey for the time period 1974-1993[9]. We consider five commodity groups *food*, *fuel* & *light*, *services*, *clothing* & *footwear*, and *total nondurable*. Our income variable is disposable non-property income. Consumption expenditure and income are deflated by the general price index of the respective month in which the household was surveyed. The attribute profile a consists of age and employment status of the head of the household, household composition, and the number of working

[7] There exists another bootstrap procedure which adjusts for scale. Yet, as the variance of these coefficients is not easy to estimate we disregard this guideline. Also this procedure will have an effect on the conclusions only if the difference between H_0 and H_1 is somewhat equivocal [Hall & Wilson (1991)].

[8] Yet, in order to guard against this arbitrary breaking point, we also took some other neighbouring years as breaking points. The overall conclusions remained unchanged.

[9] The year 1978 is excluded because it is impossible to construct the income variable due to problems in the data base.

persons in the household[10].

The estimated values of the coefficients β_t, γ_t, and $\bar{\gamma}_t$ for all five commodities are presented in Table 1. It is observed that the $\hat{\bar{\gamma}}_t$ values are very low compared to the values of the two other parameters. The values of $\hat{\bar{\gamma}}_t$ increase with the increase in the income elasticity. The lowest values of $\hat{\beta}_t$, $\hat{\gamma}_t$, and $\hat{\bar{\gamma}}_t$ are found for *food*, the highest values of these three parameters are found for *services*. In Table 1 the critical values b are given for those coefficients which are significant at the 5% or 10% level, respectively. We draw 499 bootstrap resamples and therefore, b is the 24th largest value of the 499 values of $|\hat{\bar{\gamma}}^* - \hat{\bar{\gamma}}|$ at the 5% level and the 49th largest at the 10% level. According to the above mentioned rule, i.e., if the absolute values of these coefficient estimates are greater than the corresponding critical values b, the coefficients are considered to be significantly different from zero. We can, therefore, say that $\hat{\beta}$ is significant at the 5% level for all years for all commodities. For *services*, *clothing & footwear*, and *total nondurable* $\hat{\gamma}$ is significant at the 5% level for all years, for *food* and *fuel & light* it is significant at the 5% or 10% level for most years. For *services*, the commodity group with highest income elasticity, *dispersion* matters. The coefficient $\hat{\bar{\gamma}}$ is significantly different from zero for several years. For *total nondurable*, the commodity group of major policy concern, this coefficient is significant for some years. Hence, it follows that in general the effect of income-dispersion cannot be neglected in the aggregate relation. This implies that *mean* income as a 'representative variable' is not sufficient for capturing the aggregate effect of income. Therefore, the *representative agent hypothesis* has to be rejected. Yet, for other commodities, such as *food*, *fuel & light*, and *clothing & footwear* (except two years) *mean* income as a 'representative variable' can capture the effect of income sufficiently since $\hat{\bar{\gamma}}$ is not significantly different from zero. Therefore, with respect to the income effect, the *representative agent hypothesis* can be regarded as a valid approximation. Yet, recall that we cannot conclude whether the *representative agent hypothesis* as a whole is a valid approximation because we did not consider the effects of other explanatory variables and the identity of the macro-relation form with the form of the micro-relation.

Another property which is implicitly assumed in time-series *representative agent* models is the time-invariance of the coefficients in the aggregate relation. Yet, from Table 1 we can see that the values of the coefficients differ across the years. Therefore, we also test the validity of the time-invariance property of these coefficients. In Table 2 we present the results obtained from nonstationarity tests. The results from all these tests can be enumerated as:

1. We use a parametric (unpaired t-test) and two nonparametric (Mann-Whitney U-test and Kolmogorov-Smirnov [K-S]) tests to indicate whether the distributions of two subsamples are identical with reference to the

[10] For more details see Chakrabarty & Schmalenbach (2002).

mean or other location parameters. These tests suggest differences in the distributions of the two subsampes for (a) $\hat{\hat{\gamma}}$ of *food*, (b) $\hat{\gamma}$ of *services* and *total nondurable*, and (c) $\hat{\beta}$ of *clothing & footwear*.

2. We also use the Dickey-Fuller (DF) test due to the low power of the nonparametric tests and specific distributional assumption of the parametric test described in (1). Additionally, we can examine the nature of the nonstationarity. This test suggests the following:

   i. The DF t- test with drift indicates nonstationarity, i.e., presence of a unit root for (a) $\hat{\hat{\gamma}}$ of *fuel & light*, *clothing & footwear*, and *total nondurable*, (b) $\hat{\beta}$ of *total nondurable*, and (c) $\hat{\gamma}$ of *fuel & light* at 5% level.

   ii. The DF t-test with trend suggests nonstationarity for (a) all three parameters of *fuel & light* and (b) $\hat{\beta}$ of *total nondurable* which is identified as difference-stationarity (DS) by the DF F-test. Therefore, for these parameters the trend arises due to a drift in the nonstationary random walk, not due to a deterministic trend.

3. To support the findings in (1) we tested for the presence of a deterministic trend because for the parameters mentioned in (1), for which we could not find DS. We found trend-stationarity for all four parameters.

We can, therefore, claim that the time-invariance assumption, often made in the literature, is not valid in general, which is supported by the nonstationarity of most of the parameter values. Note that not all of these coefficients are relevant, even if they are nonstationary, because they are found to be insignificant (Table 1).

4 Conclusions

Using a statistical distributional approach of aggregation this paper attempts to test the validity of assuming the existence of a *representative agent*. We are particularly concerned with the effect of income in the aggregate consumption relation. We examined two properties of the *representative agent hypothesis* i.e., the assumption of time-invariance of the coefficients in the aggregate relation and the use of only *mean* income as a ‘representative variable’ for the income distribution. It is found that the coefficients relating to the effect of the income distribution in the aggregate relation are nonstationary in most cases. This is especially interesting because the stochastic processes not only of the explanatory variables, but also of the time-varying coefficients are necessary for making predictions by using this aggregation model, a fact which is ignored in the macromodels. Additionally, the *dispersion* of income plays a significant role in the aggregate consumption relation, as shown by the significant values of the dispersion elasticity for several years, especially for the commodity groups *services* and *total nondurable*. Therefore, with respect to the income effect the *representative agent hypothesis* may be a reasonable approximation in the case of *food*, *fuel* & *light*, and *clothing* & *footwear* in this particular data set, but not in the case of *services* and *total nondurable*. This empirical test can be replicated for other data sets to judge about the *representative agent hypothesis* as an acceptable approach.

Year	Food			Fuel & light			Services		
	$\hat{\beta}$	$\hat{\gamma}$	$\hat{\bar{\gamma}}$	$\hat{\beta}$	$\hat{\gamma}$	$\hat{\bar{\gamma}}$	$\hat{\beta}$	$\hat{\gamma}$	$\hat{\bar{\gamma}}$
1974	0.290** (0.097)	0.070** (0.039)	-0.022	0.215** (0.074)	0.064* (0.059)	-0.004	0.953** (0.203)	0.333** (0.297)	0.029
1975	0.264** (0.258)	0.056	-0.027	0.157** (0.100)	0.011	-0.038	0.866** (0.212)	0.346** (0.168)	0.075
1976	0.263** (0.263)	0.079* (0.073)	-0.004	0.191** (0.065)	0.027	-0.033	0.871** (0.198)	0.417** (0.194)	0.144
1977	0.238** (0.224)	0.060* (0.059)	-0.015	0.155** (0.074)	0.057** (0.047)	0.011	0.900** (0.332)	0.420** (0.201)	0.148** (0.128)
1979	0.254** (0.237)	0.067	-0.017	0.265** (0.065)	0.108** (0.067)	0.021	0.960** (0.252)	0.435** (0.174)	0.120* (0.115)
1980	0.263** (0.203)	0.096	0.005	0.229** (0.058)	0.060	-0.019	0.762** (0.198)	0.382** (0.172)	0.118* (0.112)
1981	0.226** (0.217)	0.044	-0.032	0.183** (0.068)	0.106** (0.077)	0.044	0.863** (0.117)	0.338** (0.074)	0.046
1982	0.290** (0.268)	0.085	-0.003	0.222** (0.088)	0.093** (0.056)	0.026	0.883** (0.133)	0.406** (0.115)	0.138** (0.098)
1983	0.189** (0.179)	0.036	-0.024	0.243** (0.075)	0.090** (0.071)	0.013	0.900** (0.161)	0.422** (0.090)	0.138** (0.082)
1984	0.295** (0.291)	0.078* (0.072)	-0.013	0.281** (0.067)	0.088** (0.044)	0.001	1.029** (0.280)	0.431** (0.148)	0.112* (0.096)
1985	0.265** (0.248)	0.033	-0.056	0.235** (0.054)	0.093** (0.052)	0.014	0.887** (0.161)	0.342** (0.092)	0.045
1986	0.258** (0.228)	0.071	-0.020	0.192** (0.054)	0.111** (0.073)	0.043	0.874** (0.269)	0.464** (0.178)	0.155** (0.146)
1987	0.225** (0.216)	0.059* (0.056)	-0.027	0.200** (0.052)	0.072** (0.047)	-0.005	0.858** (0.233)	0.438** (0.164)	0.109
1988	0.246** (0.235)	0.071* (0.065)	-0.031	0.271** (0.109)	0.094** (0.076)	-0.018	1.005** (0.157)	0.633** (0.238)	0.217** (0.215)
1989	0.230** (0.217)	0.041	-0.054* (0.050)	0.177** (0.055)	0.048* (0.048)	-0.025	0.907** (0.180)	0.399** (0.088)	0.024
1990	0.244** (0.219)	0.080	-0.021	0.172** (0.055)	0.076** (0.057)	0.001	1.017** (0.384)	0.573** (0.207)	0.127
1991	0.232** (0.211)	0.055	-0.044	0.208** (0.080)	0.105** (0.076)	0.016	0.842** (0.118)	0.388** (0.183)	0.029
1992	0.238** (0.201)	0.066** (0.058)	-0.030	0.167** (0.066)	0.095** (0.053)	0.027	0.900** (0.160)	0.475** (0.194)	0.118
1993	0.249** (0.246)	0.085** (0.079)	-0.014	0.224** (0.223)	0.090** (0.079)	0.001	0.920** (0.920)	0.456** (0.402)	0.092
mean	0.250	0.065	-0.023	0.210	0.078	0.004	0.905	0.426	0.104

Table 1: Estimated values of β, γ, and $\bar{\gamma}$ for 5 commodity groups for each year using the DADE procedure. Critical values calculated by bootstrap are given in parentheses for the significant values.
**Significant at 5% level (absolute value is greater than 5% critical value, given in parentheses).
*Significant at 10% level (absolute value is greater than 10% critical value, given in parentheses).

Year	Clothing & footwear			Total nondurable		
	$\hat{\beta}$	$\hat{\gamma}$	$\hat{\hat{\gamma}}$	$\hat{\beta}$	$\hat{\gamma}$	$\hat{\hat{\gamma}}$
1974	0.829** (0.224)	0.268** (0.128)	0.004	0.572** (0.103)	0.218** (0.073)	0.035
1975	0.824** (0.244)	0.369** (0.217)	0.111	0.538** (0.119)	0.204** (0.128)	0.036
1976	0.857** (0.155)	0.351** (0.242)	0.083	0.558** (0.047)	0.218** (0.069)	0.043* (0.041)
1977	0.940** (0.411)	0.394** (0.238)	0.110	0.554** (0.194)	0.202** (0.093)	0.034
1979	0.695** (0.178)	0.254** (0.211)	0.026	0.557** (0.062)	0.219** (0.041)	0.036* (0.030)
1980	0.857** (0.280)	0.369** (0.150)	0.073	0.531** (0.109)	0.203** (0.060)	0.020
1981	0.718** (0.138)	0.229** (0.211)	-0.014	0.563** (0.065)	0.203** (0.040)	0.013
1982	0.927** (0.209)	0.410** (0.158)	0.129** (0.111)	0.606** (0.085)	0.250** (0.071)	0.067** (0.053)
1983	0.710** (0.164)	0.288** (0.115)	0.064	0.541** (0.068)	0.204** (0.042)	0.033* (0.033)
1984	0.849** (0.192)	0.331** (0.163)	0.068	0.618** (0.104)	0.243** (0.109)	0.051
1985	0.836** (0.208)	0.322** (0.117)	0.043	0.579** (0.081)	0.230** (0.055)	0.037
1986	0.714** (0.155)	0.330** (0.116)	0.078	0.586** (0.139)	0.260** (0.072)	0.053* (0.053)
1987	0.782** (0.177)	0.320** (0.090)	0.020	0.558** (0.077)	0.212** (0.046)	-0.002
1988	0.726** (0.131)	0.313** (0.098)	0.013	0.571** (0.056)	0.281** (0.080)	0.045
1989	0.749** (0.143)	0.360** (0.105)	0.051	0.545** (0.079)	0.235** (0.056)	0.009
1990	0.663** (0.202)	0.316** (0.104)	0.025	0.533** (0.141)	0.252** (0.061)	0.018
1991	0.693** (0.129)	0.183** (0.101)	-0.113** (0.093)	0.537** (0.073)	0.193** (0.078)	-0.037
1992	0.661** (0.124)	0.257** (0.119)	-0.009	0.534** (0.062)	0.226** (0.057)	0.011
1993	0.740** (0.739)	0.379** (0.332)	0.088	0.563** (0.563)	0.248** (0.198)	0.025
mean	0.777	0.318	0.045	0.560	0.226	0.028

Table 1: Continued

Critical values b are given in parentheses.

Parameters for five commodity groups		Test Statistics						
		Unpaired t-test $\|t\|$-value	U-test	K-S test	DF (with drift) t-ratio	DF(with trend) t-ratio	DF(with trend) F-ratio	Deterministic trend ($\|t\|$)-value
Food	$\hat{\beta}$	1.267	28	0.489	-6.535	-7.515	28.540	-0.002 (1.581)
	$\hat{\gamma}$	0.491	39	0.178	-8.925	-8.648	37.670	-0.00002 (0.030)
	$\hat{\tilde{\gamma}}$	2.937*	16*	0.576*	-4.183	-5.424	14.790	-0.001* (1.860)
Fuel & light	$\hat{\beta}$	0.512	40	0.269	-3.770	-3.651*	6.680*	0.0002 (0.095)
	$\hat{\gamma}$	1.333	31	0.389	-2.748*	-3.213*	5.160*	0.002 (2.219)
	$\hat{\tilde{\gamma}}$	0.368	41	0.200	-3.062*	-3.045*	4.650*	0.001 (.924)
Services	$\hat{\beta}$	0.402	42	0.178	-4.832	-4.998	12.580	0.002 (0.665)
	$\hat{\gamma}$	2.130*	20*	0.668*	-4.279	-5.640	15.990	0.007* (2.523)
	$\hat{\tilde{\gamma}}$	0.215	37	0.267	-5.851	-5.714	16.620	0.0003 (0.013)
Clothing & footwear	$\hat{\beta}$	2.689*	19*	0.589*	-4.040	-7.775	30.230	-0.009* (3.163)
	$\hat{\gamma}$	0.625	36	0.378	-5.353	-5.962	17.780	-0.002 (0.755)
	$\hat{\tilde{\gamma}}$	1.745	26	0.478	-3.647*	-4.857	11.801	-0.004 (1.754)
Total nondurable	$\hat{\beta}$	0.674	42	0.233	-3.690*	-3.585*	6.460*	-0.001 (0.508)
	$\hat{\gamma}$	2.079*	22*	0.578*	-4.411	-5.482	15.060	0.002* (1.950)
	$\hat{\tilde{\gamma}}$	1.914	27	0.467	-3.629*	-4.439	9.862	-0.002 (2.052)

Table 2: Tests for time invariance of parameters

* Indicates nonstationarity

- Unpaired t-test: Null hypothesis H_0: means are equal in two subsample; alternative H_1: means are not equal.

- U-test: Null hypothesis H_0: distributions are identical in two subsamples; alternative H_1: distributions differ in terms of location parmeter. H_0 is accepted if the test statistic $>$ critical value. 1% and 5% critical values are 16 and 24 respectively.

- K-S test: Null hypothesis H_0: distributions are identical; alternative H_1: distributions differ in any manner. H_0 is accepted if the test statistic $<$ critical value. 1% and 5% critical values are .69 and .57 respectively.

- DF (with drift): The equation estimated is of the type: $\Delta z_t = \alpha + \eta z_{t-1} + \epsilon_t$. Null hypothesis $H_0 : \eta = 0$, i.e., unit root; alternative H_1: stationarity. Unit root is accepted if the t-ratio $>$ DF critical value. 1% and 5% critical values are -3.75 and -3.00 respectively.

- DF (with trend): The equation estimated in this case is the following: $\Delta z_t = \alpha + \eta z_{t-1} + \rho t + \epsilon_t$. For t-ratio the null hypothesis is unit root, i.e., $\eta = 0$ and the criteria to accept the null is the same as above. 1% and 5% critical values are -4.38 and -3.60 respectively.
  For F-ratio: null hypothesis $H_0 : \eta = 0$ & $\rho = 0$; alternative $H_1 : \eta < 0$. This tests for difference stationarity (DS) against trend stationarity (TS). H_0 is accepted if F-ratio $<$ DF critical value. 1% and 5% critical values are 7.24 and 10.61 respectively.

- In the final column we also estimate the equation $z_t = \alpha + \rho t + \epsilon_t$ to detect the significance of the deterministic trend when the F test rejects DS.

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