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INFLATION, RELATIVE PRICES AND HOUSEHOLD SAVING BEHAVIOR

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INFLATION, RELATIVE PRICES AND HOUSEHOLD

SAVING BEHAVIOR*

by

Matti Virén

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CONTENTS

Abstract

This article contains tests for the direct effects of inflation on household saving. Empirical results with cross-country data from five OECD countries indicate that both unanticipated and anticipated inflation affect the household savings ratio. This latter result, which furthermore turns out to be robust in terms of aggregation and additional variables, is clearly at variance with some previous analyses, particularly those of Deaton.

1. INTRODUCTION

The role of inflation vis-â-vis household saving has been one of the major issues of empirical analysis during the last decade. This has obviously something to do with the fact that there have been considerable variations in both savings ratios and rates of inflation in most of the OECD countries during the 1970's. Empirical analyses have - almost without exception - suggested that inflation has a positive effect on saving. From the theoretical point of view there is much less consensus concerning the possible channels of this relationship. This is because the positive, or ambiguous, relationship is predicted by, for instance, the "real income uncertainty hypothesis" proposed by Juster and Wachtel (1972), the "disequilibrium hypothesis" put forward by Deaton (1977), the "mis-measurement hypothesis" posited by e.g. Hendry and von Ungern-Sternberg (1981), and the "real balance effect" discussed by e.g. Howard (1978) (see also Williams (1979) for an exhaustive survey of these hypotheses). The problem with most of these hypotheses is that they lead to similar specifications which are virtually indistinguishable with pure time-series data. Therefore interpretation of empirical findings becomes rather difficult, as has been pointed out by Howard (1978). Despite this problem, we try here to evaluate the emprical relevance of these hypotheses, particularly concerning Deaton´s hypothesis.

Thus, when carrying out the empirical analysis we use Deaton´s hypothesis as the main frame of reference. This hypothesis, which strictly speaking concerns the effects of unanticipated inflation on the household savings ratio, has, among other things, the advantage of giving specific predictions concerning different elements of inflation, in particular unanticipated and anticipated inflation. This distinction seems to be crucial in terms of different theoretical hypotheses, and we therefore concentrate on testing this distinction in the subsequent empirical analysis. An important feature of these tests is that we allow for a disaggregation of consumer prices. This is done in order to scrutinize the role of possible commodity-specific factors. It is also motivated by a desire to check whether the notion of "shopping habits", which lies behind Deaton´s hypothesis (Deaton (1977)), is of any relevance. Given that the rate of inflation and changes in relative prices are highly correlated, it might be worthwhile checking whether it is, in fact, relative prices which constitute the main determinant of the household savings ratio. Therefore, this possibility is also tested in this study. The relevant tests make use of the Extended Linear Expenditure System (ELES), which is practically the only framework that allows for such relative price effects.

In what follows, theoretical considerations are discussed in section 2, while section 3 is devoted to empirical analyses which make use of quarterly cross-country data from five OECD countries covering the period 1968-1982. Finally, some concluding remarks follow in section 4.

2. THEORETICAL CONSIDERATIONS

As pointed out above we utilize here the saving function specification of Deaton (1977). This specification takes the following form:

$$(S/Y)_t = n(1-k) + b_1(y_t - y_t^e) + b_2(p_t - p_t^e) + (1-n)(S/Y)_{t-1} \tag{1}$$

where S indicates household saving, Y households´ disposable income, y_t and p_t the proportional changes in real income and prices, respectively, y_t^e and p_t^e are the corresponding anticipated values and (1-k) indicates the "equilibrium" savings ratio. According to (1) both the unanticipated real income change rate and the inflation rate affect the household savings ratio positively. Another feature of the model is that savings adjust sluggishly to the corresponding "equilibrium" level, (1-k). This level, which was assumed to be constant by Deaton (1977), depends on such things as the real rate of interest, the distribution of income, and so on (cf. Friedman (1957), p. 216). In the subsequent empirical analysis we allow for (1-k) to depend on the real rate of interest; lack of suitable data, however, prevents tests using income distribution variables. Besides the real rate of interest we also take "anticipated prices" into account by allowing (1-k) to depend on relative prices. This relative price effect is specified according to the Extended Linear Expenditure System (ELES), cf. Lluch, Powell and Williams (1977). The basic idea here is that relative prices affect the relationship between the supernumerary income and the actual income, and this, in turn, affects the household savings ratio. The analysis of Lluch et al reduces to the following specification:

$$(2) \qquad (S/Y)_t = (1-\mu)\left(1 - \frac{\Sigma P_{it}\alpha_i}{Y_t}\right)$$

where S denotes household saving and Y household disposable income, both in nominal terms, the P_i´s denote commodity prices, the α_i´s pre-committed levels of consumption and μ the marginal propensity to consume. Given this specification, the household savings ratio should be higher if the prices of non-necessities are higher than the prices of necessary commodities (necessities have high and non-necessities low α_i values)[1]. Besides estimating equation (2) directly we test the ELES effect by replacing the (1-k) term in (1) by equation (2) and by then estimating jointly the parameters α_i for the rest of the specificaton (1).

Next, we consider the inflation "innovation" term, $p_t - p_t^e$, which is subsequently denoted by $\tilde{p}_t$ (similarly $\tilde{y}_t$ corresponds to $y_t - y_t^e$). As pointed out earlier, its coefficient should be positive, the absolute magnitude depending on the way in which price information is conveyed and hence on the structure of retailing and shopping habits. The more commodities are disaggregated, the greater the possibilities of substitution, and thus the greater the absolute magnitude of b_1. Consequently, the coefficient should be small in absolute terms, if goods are purchased together, say in an hypermarket and large if commodities are bought separately, each in a different store (cf. Deaton (1977), p. 902). The foregoing discussion suggests that the importance of Deaton´s hypothesis might vary very much across commodities; this proposition becomes even more obvious if one thinks about consumers´ ability to distinguish between absolute and relative commodity prices, or price change rates, in the case of important quality differences between commodities.

Now, if we consider formal hypothesis testing we should first test for the hypothesis that, given different hypotheses on expectations formation, b_1 is indeed positive, and accordingly, that the coefficient of the anticipated inflation term is equal to zero. Notice that this latter hypothesis requires that the equilibrium savings ratio does <u>not</u> depend on the anticipated inflation rate. Obviously this is not true because it is precisely the real rate of interest which determines this anticipated inflation. One could, of course, argue that the corresponding income and substitution effects work in opposite directions and thus that the total effect is negligible. However, if the real rate of interest is introduced into the model, anticipated inflation should not have an independent role in this specification.

As far as the disaggregation issue is concerned, we test, on the one hand, whether Deaton´s hypothesis is robust, in particular, with respect to the role of anticipated vis-à-vis unanticipated inflation. On the other hand, we analyze to what extent inflation effects are due only to commodity-specific factors, and test whether the overall performance of Deaton´s hypothesis can be increased by using different commodity group prices instead of the aggregate price index.

3. ESTIMATION RESULTS

3.1. Data

Cross-country data from 5 OECD countries (Australia, France, Japan, the United Kingdom and the United States) covering, with some minor exceptions, the period 1968.2-1982.3 are used in the empirical analysis. The data are seasonally adjusted and unweighted; purchases of consumer durables are included in our measure of consumption. A detailed description of the time series utilized can be found in the Appendix.

3.2. Results of regression analysis

We start by analyzing the role of (anticipated) relative prices. As a general frame of reference we use the following specification which is derived from equations (1) and (2):

$$(3) \qquad (S/Y)_t = \sum_{j=1}^{5} D_j + \sum_{i=1}^{6} a_i(P_{it}/Y_t) + a_7\tilde{p}_t + a_8\tilde{y}_t + a_9r_t + a_{10}\Delta U_t + a_{11}(S/Y)_{t-1} + u_t$$

where the D_j´s denote individual country intercepts (which are used with pooled cross country data), the P_{it}´s are the implicit deflators of consumption expenditure in commodity group i (i = 1,2,...,6), Y_t is the nominal (households´ disposable) income, $\tilde{y}_t$ and $\tilde{p}_t$ are the unanticipated real income change rate and the inflation rate, respectively, r_t is the real rate of interest, ΔU_t is the first difference of the unemployment rate, and, finally, u_t is the error term. As stated earlier, the a_i´s should correspond to the subsistence consumption of the i^{th} good (strictly speaking, $a_i = (1-\mu)n\alpha_i$). Given (1) and (2), the respective coefficient estimates should have negative signs.

As far as the real income change rate and inflation innovations are concerned, we have applied two methods of constructing proxies for these unobservables. First, we have used (following e.g. Deaton (1977)) the constant expectations hypothesis, which implies that p_t^e and y_t^e are assumed to be constant over time in individual countries. This, in turn, implies that $\tilde{p}_t$ and $\tilde{y}_t$ are simply the corresponding log differences; the constant values of p^e and y^e are lumped into the the constant term, or in this case, into individual country intercepts. As an alternative way of constructing these innovation proxies we have used the ARIMA-technique. Thus, we have fitted an ARIMA model for each country into the y_t and p_t (and p_{it}) series, and used the residuals of these ARIMA models as proxies for these unanticipated changes. In order to save space, the corresponding estimates are not presented here - they are, however, available upon request from the author. The important point about the ARIMA residuals is the fact that they are white noise, which is in accordance with the rational expectations framework.

Furthermore, the ARIMA technique makes it possible to test formally the importance of anticipated and unanticipated inflation.

There is one apparent problem in distinguishing between the anticipated and unanticipated inflation and real income change rate terms due to the fact that there have been introductions of an indexation rules in several countries during the period of investigation (see e.g. Braun (1976) for details). One might assume that these rules would at least invaliditate the constancy of parameters. Empirical evidence, suggests, however, that this problem might not be very serious.[2)]

The additional variables, r_t and ΔU_t, are introduced into (3) mainly in order to check the robustness of results. The real interest rate variable, r_t, can be interpreted as a determinant of the equilibrium savings ratio, 1-k, possibly along with relative prices, while the unemployment rate variable, ΔU_t, might capture the effects of income or employment uncertainty. Both of these variables should have positive signs, even though strictly speaking the sign of the real interest rate is ambiguous due to conflicting income and substitution effects.

In this context we use a six-commodity classification for consumption expenditure. The relevant items are listed below:

1. Food, beverages and tobacco
2. Clothing and footwear
3. Gross rent, fuel and power
4. Transport and communication
5. Furniture and household operations
6. Other goods and services

This classification makes sense where the ELES-type models are concerned but not necessarily when analyzing the effects of inflation on the demand for consumer durables and liquid assets[3)]. However, the data do no allow for a more detailed classification of goods. As stated above, these data were obtained for five countries: Australia, France, Japan, the United Kingdom and the United States, the period of estimation being, with some minor exceptions, 1968.2-1982.3. When the onservations are pooled, there are altogether 268 observations (263 when the ARIMA models are used), which obviously solves the degrees of freedom problem in the context of disaggregated price indices. Although this paper concentrates on the results with pooled cross-country data, all analyses have also been carried out using individual country data. The results obtained are, however, presented only in cases, where they are at variance with those of the pooled data[4)].

We now turn to the estimation results of equation (3) presented in Table 1. This Table includes OLS estimates for the coefficients of the "relative prices", $(P_i/Y)_t$, inflation and income change rate innovations, $\tilde{p}_t$ and $\tilde{y}_t$, the first difference of the unemployment rate, ΔU_t, the real rate of interest, r_t, the lagged savings ratio and the individual country intercepts, D_j. When proxying the innovations terms, ARIMA innovations have been used for equations (3) and (4) and constant expectations for equations (5) and (6) so that $\tilde{p}_t$ and $\tilde{y}_t$ are in this latter case simply the corresponding log differences. r_t in equation (2) is constructed using the fitted value of the ARIMA model with respect to $\Delta \log P_t$. The diagnostic statistics at the bottom of Table 1 are Durbin´s m-statistic for first-order autocorrelation (in the presence of a lagged dependent variable) while LM(4) is the Breusch LM statistic for residual AR(4) process, cf. Breusch (1978)[5)].

The estimation results clearly indicate that the ELES framework does not contribute at all to explaining the changes in the savings ratio over time (this result accords fairly well with the findings of Lluch et al

Table 1. OLS Estimates of equation (3)

	(1)	(2)	(3)	(4)	(5)	(6)
(P_1/Y)	4.917 (0.21)	.714 (0.03)	8.569 (0.49)	4.318 (0.26)	7.374 (0.61)	4.134 (0.35)
(P_2/Y)	10.190 (0.27)	5.610 (0.15)	22.524 (0.81)	20.762 (0.78)	20.524 (1.05)	29.564 (1.50)
(P_3/Y)	-9.727 (0.28)	-4.986 (0.15)	28.545 (1.12)	30.410 (1.25)	28.827 (1.61)	24.851 (1.44)
(P_4/Y)	-22.131 (1.42)	-23.022 (1.45)	-3.918 (0.34)	-7.435 (0.67)	-1.796 (0.22)	-9.058 (1.09)
(P_5/Y)	10.226 (0.21)	11.430 (0.24)	-28.167 (0.78)	-25.474 (0.75)	-28.498 (1.13)	-28.771 (1.18)
(P_6/Y)	.808 (0.03)	3.937 (0.16)	-24.099 (1.31)	-20.527 (1.17)	-21.739 (1.68)	-17.889 (1.41)
$\tilde{p}$			.141 (1.86)	.207 (2.81)	.258 (5.43)	2.38 (5.00)
$\tilde{y}$			.594 (14.69)	.596 (15.39)	.662 (25.78)	.660 (26.74)
r		-.406 (2.46)		-.298 (2.48)		.035 (1.08)
ΔU		.836 (2.27)		1.142 (4.33)		.831 (4.49)
$(S/Y)_{-1}$	.602 (12.15)	.555 (11.03)	.782 (20.05)	.726 (18.77)	.862 (29.04)	.830 (28.18)
Australia	.053 (7.57)	.059 (8.31)	.029 (5.31)	.036 (6.59)	.006 (1.47)	.007 (1.50)
France	.105 (2.08)	.118 (2.37)	.008 (0.21)	.028 (0.78)	-.025 (0.95)	-.096 (0.37)
Japan	.080 (7.75)	.087 (8.38)	.044 (5.45)	.053 (6.74)	.015 (2.49)	.020 (3.33)
UK	.046 (7.57)	.051 (8.30)	.025 (5.28)	.030 (6.43)	.057 (1.63)	.045 (1.03)
USA	.038 (3.51)	.043 (3.97)	.012 (1.42)	.018 (2.22)	-.005 (0.85)	-.024 (0.42)
R^2	.8746	.8807	.9331	.9399	.9675	.9702
m	-.445 (4.47)	-.405 (3.96)	-.016 (0.20)	.015 (0.19)	.195 (2.89)	.194 (2.90)
LM(4)	49.280	43.084	29.695	30.545	16.528	16.026
F	.567	.427	.618	.753	1.149	1.105

t-ratios are shown in parentheses. F indicates a joint significance test for the (P_i/Y)-terms. ARIMA innovations have been used for $\tilde{p}$ and $\tilde{y}$ in equations (3) and (4) and constant expectations in equations (5) and (6).

(1977), p. 84-85). The coefficient estimates of the (P_i/Y)-terms are so imprecise that the corresponding F-statistics (see the last row in Table 1) do not allow us to reject the hypothesis that these coefficients are identically equal to zero. The result seems to be robust with respect to additional variables ΔU_t and r_t as well as to the treatment of the innovation terms $\tilde{p}_t$ and $\tilde{y}_t$. There is only a slight caveat due to the Breusch LM-statistics, which indicate that there is some higher order autocorrelation in the residuals (mainly of the fourth order suggesting that the seasonal adjustment of the data has not been appropriate)[6]

Given these results, we can proceed to analyze the effects of unanticipated and anticipated inflation assuming that the "equilibrium savings ratio" is either constant or depends on the real rate of interest. The following equation is now used as a frame of reference:

$$(4) \qquad (S/Y)_t = \sum_{j=1}^{5} D_j + \sum_{i=1}^{6} b_i \tilde{p}_{it} + b_7 \tilde{y}_t + b_8 r_t + b_9 \Delta U_t + b_{10}(S/Y)_{t-1} + e_t$$

where the $\tilde{p}_{it}$'s denote the innovation terms of the relative rates of change of the individual commodity prices. The OLS estimation results are presented in Table 2: columns (1) to (4) correspond to ARIMA innovations and columns (5) to (8) to pure log differences, which, in turn, correspond to the idea of constant expectations. Now, consider first the performance of the basic Deaton specification (cf. (1) and (5) in Table 2). Clearly the estimation results give some support for this specification: the parameter estimates are of expected sign and magnitude and fairly precisely estimated. The additional variables ΔU_t and r_t, which behave well in accordance with 'a priori theorizing', do not invalidate this result. However, the coefficient estimate of $\tilde{p}_t$ becomes smaller and less precise when the log differences (of the constant expectations specification) are replaced by the corresponding ARIMA innovations. This result could be due to the fact the assumption of constant inflation rate expectations is better than the corresponding ARIMA assumption or because the implicit assumption of Deaton´s specification that anticipated inflation has no effect on savings is simply inappropriate! As we shall see later on, it is just this latter assumption which receive support from the data.

As far as the individual commodity prices are concerned, it can be observed that the coefficient estimates are to some degree in accordance with the theoretical justification of Deaton´s hypothesis. That is, the coefficient estimates for durables and services (p_5 and p_6, respectively) have the right signs and are fairly precise while the opposite seems to be true for the other commodities. The fact that, for instance, the coefficient estimate of p_2 (clothing) is negative suggests that the performance of Deaton´s hypothesis is, after all, far from perfect.

Anyway, disaggregating the aggregate price index into 6 commodity price indices does clearly increase the explanatory power of the equations. In the case of the constant expectations specification the importance of disaggregation was tested in the following way: $\Delta \log P_t$ was replaced by the corresponding Divisia indices $\Sigma\, w_{it} \Delta \log P_{it}$, which should have the same coefficients if the disaggregation does not matter. F-statistics allow, however, for rejecting this hypothesis ($F_{5,255} = 3.986$ and $F_{5,253} = 2.792$; the first statistic refers to the equation without the r_t and ΔU_t-terms while the second refers to the equation including these terms). One could also clearly reject the hypothesis that the coefficients of

Table 2. OLS Estimates of Equation (4)

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
$\tilde{p}_1$			-.065	-.043			.034	.048
			(1.13)	(0.78)			(0.90)	(1.29)
$\tilde{p}_2$			-.102	-.097			-.111	-.104
			(1.81)	(1.77)			(3.20)	(3.09)
$\tilde{p}_3$			-.025	-.024			.052	.008
			(0.35)	(0.35)			(1.23)	(0.19)
$\tilde{p}_4$			.003	.019			-.009	-.005
			(0.13)	(0.76)			(0.64)	(0.39)
$\tilde{p}_5$			.151	.135			.171	.186
			(2.14)	(2.01)			(3.78)	(4.26)
$\tilde{p}_6$			.178	.185			.120	.076
			(2.08)	(2.25)			(2.19)	(1.42)
$\tilde{p}$	.144	.211			.263	.244		
	(1.92)	(2.88)			(5.77)	(5.15)		
$\tilde{y}$	.588	.589	.603	.605	.658	.660	.673	.672
	(14.86)	(15.48)	(14.81)	(15.50)	(26.15)	(27.03)	(26.75)	(27.61)
r		-.297		-.258		.012		.014
		(2.61)		(2.26)		(0.50)		(0.57)
ΔU		1.120		1.084		.818		.824
		(4.32)		(4.20)		(4.44)		(4.46)
$(S/Y)_{-1}$	.786	.730	.797	.746	.863	.843	.880	.863
	(20.86)	(19.40)	(21.22)	(19.83)	(30.49)	(30.41)	(31.36)	(31.54)
Australia	.029	.035	.027	.033	.006	.007	.004	.006
	(5.43)	(6.67)	(5.17)	(6.28)	(1.49)	(1.61)	(1.06)	(1.33)
France	.035	.046	.033	.043	.010	.012	.007	.009
	(5.52)	(7.01)	(5.26)	(6.60)	(2.03)	(2.27)	(1.45)	(1.76)
Japan	.043	.052	.041	.049	.015	.018	.012	.015
	(5.55)	(6.84)	(5.27)	(6.46)	(2.57)	(3.17)	(2.09)	(2.67)
UK	.025	.030	.024	.028	.006	.006	.003	.003
	(5.46)	(6.54)	(5.27)	(6.21)	(1.70)	(1.41)	(0.89)	(0.87)
USA	.016	.019	.015	.018	.001	.001	-.001	-.000
	(4.84)	(5.92)	(4.60)	(5.53)	(0.32)	(0.23)	(0.27)	(0.17)
R^2	.9321	.9388	.9357	.9417	.9669	.9694	.9693	.9717
m	-.032	.002	-.013	.015	-.203	-.209	-.165	-.183
	(0.42)	(0.03)	(0.17)	(0.20)	(3.07)	(3.20)	(2.44)	(1.83)
LM(4)	28.310	29.597	32.027	30.205	16.767	16.819	19.175	17.296

t-ratios are shown in parentheses. Equations (1)-(4) correspond to ARIMA innovations and (5)-(8) to constant expectations.

$\tilde{p}_{it} = \Delta \log P_{it}$ in Table 2 are identically equal to zero ($F_{6,256} = 9.173$ and $F_{6,254} = 8.078$). Testing for the importance of disaggregation is a more complicated matter in the case of ARIMA innovations because using $\tilde{p}_t$, on the one hand, and the $\tilde{p}_{it}$´s, on the other, leads to non-nested specifications. We have adopted the conventional F-test framework in carrying out these tests (see Mizon and Richard (1983) and Dastoor (1983) for details). The respective F-statistics for the additional $\tilde{p}_{it}$ terms, given Deaton´s basic specification with the aggregate price index, suggest that disaggregation pays off - even though only at the margin ($F_{6,250} = 2.197$ and $F_{6,248} = 2.323$). Finally, it once again emerges that we can reject the hypothesis that the disaggregated inflation innovation terms are identically equal to zero ($F_{6,251} = 3.034$ and $F_{6,249} = 3.570$)[7].

All in all, the results thus far presented have not been totally unfavourable for Deaton´s saving function specification. It is only that the coefficient of the inflation rate variable has appeared to be rather sensitive to different proxies for unanticipated inflation and that disaggregating the inflation rate leads to partially unsatisfactory estimation results. We have not, however, performed any strong test with respect to Deaton´s hypothesis. Therefore, we next turn to test the hypothesis that anticipated inflation does not affect the household saving ratio, or that it does so only via the real rate of interest. These hypotheses are next tested by using the ARIMA technique. Table 3 contains the coefficient estimates for the predicted values of p_t and p_{it} as well as the corresponding innovation terms. These results clearly indicate that the role of anticipated inflation is far from negligible, as the following F-statistics for the joint significance of the p^e_{it}-terms in equations (3) and (4) in Table 3 show: $F_{6,244} = 6.790$ and $F_{6,242} = 6.348$.

Table 3. Estimates of the anticipated and unanticipated inflation terms

	(1)		(2)		(3)		(4)	
$\tilde{p}_1$ & p_1^e					-.051 (0.90)	.328 (3.67)	-.031 (0.56)	.341 (3.87)
$\tilde{p}_2$ & p_2^e					-.156 (2.86)	.031 (0.37)	-.143 (2.67)	.006 (0.08)
$\tilde{p}_3$ & p_3^e					-.002 (0.03)	.225 (2.57)	-.015 (0.21)	.184 (2.12)
$\tilde{p}_4$ & p_4^e					.006 (0.24)	-.037 (1.12)	.015 (0.65)	-.047 (1.43)
$\tilde{p}_5$ & p_5^e					.178 (2.63)	.187 (1.54)	.173 (2.61)	.218 (1.83)
$\tilde{p}_6$ & p_6^e					.112 (1.35)	.078 (0.56)	.104 (1.28)	-.093 (0.61)
$\tilde{p}$ & p^e	.163 (2.28)	.624 (5.47)	.184 (2.56)	.537 (3.74)				
$\tilde{y}$	.596 (15.87)		.605 (16.17)		.621 (16.06)		.619 (16.30)	
r			.012 (0.09)				-.107 (0.85)	
ΔU			.897 (3.45)				.959 (3.66)	
$(S/Y)_{-1}$	.723 (19.28)		.709 (19.06)		.734 (19.34)		.721 (19.31)	
Australia	.022 (4.27)		.025 (4.40)		.017 (3.17)		.023 (4.00)	
France	.032 (5.22)		.026 (5.23)		.026 (4.19)		.033 (4.89)	
Japan	.044 (5.92)		.048 (6.41)		.038 (5.18)		.044 (5.93)	
UK	.015 (3.20)		.017 (3.17)		.009 (1.71)		.015 (2.54)	
USA	.010 (3.12)		.011 (2.94)		.007 (2.03)		.010 (2.73)	
R^2	.9393		.9420		.9449		.9479	
m	-.037 (0.48)		-.019 (0.25)		.008 (0.10)		.015 (0.20)	
LM(4)	22.436		20.380		23.567		37.787	
F	12.548		4.334		4.941		4.255	

t-ratios are shown in parentheses. F indicates a F-test statistic for the equality of the coefficients of $\tilde{p}$ and p^e, or the $\tilde{p}_i$'s and the p_i^e's.

Equation (1) in Table 3 shows that anticipated inflation even has a larger (positive) effect on the savings ratio than unanticipated inflation.[8] This result holds even if the real rate of interes is introduced into the equation. Further, equations (3) and (4) indicate that this result also holds when general anticipated inflation is decomposed in terms of different commodity groups.[9] In this case, it is only for consumer durables prices that unanticipated inflation has an effect which is in accordance with Deaton´s hypothesis while anticipated inflation performs clearly better with almost all commodity prices. Why does anticipated inflation have such a strong effect? There seems to be no clear answer to this question. One possible answer is that consumers respond to an increase in the inflation rate by accumulating more financial assets or by purchasing in advance such storeables as houses (cf. Fortune (1981)). Besides this "real balance effect"-type hypothesis one can also mention the possibility that the practice of fixing nominal wages for a finite period so that real wages are eroded by inflation leads to increased saving when the inflation rate is increasing (cf. Bulkley (1981) for details). Finally, one could imagine that in the case of fixed down payment ratios, or of fixed nominal borrowing limits, under credit rationing, an increase in the rate of inflation forces households to accumulate more financial assets thereby increasing saving (cf. Jackman and Sutton (1982), p. 120). The high values of the coefficients of p_3^e (corresponding to the rate of price change of housing) in Table 3 are, in fact, well in accordance with this hypothesis.[10]

In the light of these results it seems easy to explain why the constant expectations hypothesis has been so superior in the context of Deaton´s specification (cf. Deaton (1977) and Koskela and Virēn (1982a,b)).

Obviously, this is because the constant expectations specification takes into account some effects of anticipated inflation.[11)] As the F-statistics in Table 3 indicate, one cannot, however, restrict the coefficients of unanticipated and anticipated inflation terms to be equal. Thus, there is little justification for specifying the saving function solely in terms of the actual inflation rate, as is typically done (cf. e.g. Davidson et al (1978)).

4. CONCLUDING REMARKS

The main finding of this study is that it is not only unanticipated inflation which affects (positively) household saving but also anticipated inflation. This result is strikingly robust with respect to the aggregation of the price index and with respect to the treatment of the real rate of interest and the unemployment rate variables. Given this result, one can conclude that there is indeed a lot to be done in analyzing inflation and household saving behavior. This study suggests moreover that the problems exist not only because of the great number of competing theoretical explanations but also because of various commodity specific factors.

FOOTNOTES

1) The relative price argument implied by the ELES would be particularly useful if we used a simple Keynesian consumption function: $C = a_0 + b_0(Y/P)$ as frame of reference. This function would imply that the average propensity to consume declines as real income increases. If, instead, the parameter a_0 depended on relative prices as in (2), changes in relative prices might conceivably compensate for the increase in real income. This is because increases in the relative prices of commodities with large subsistence components might produce an increase in the average propensity to consume. (cf. Lluch and Williams (1975)). Obviously, constancy of the consumption structure (implied by the constant α_i´s) is not a very convincing assumption. One way of solving this problem is to assume that the pre-commited levels of consumption (α_i) depend on some stock of habits, which, in turn, can be expressed by a "habit formation" specification of the Linear Expenditure System (see e.g. Pollack (1976)). Thus, equation (2) takes the form $(S/Y)_t = (1-\mu)(1 - (\Sigma\ \alpha_i P_{it} Q_{it-1})/Y_t)$, where Q indicates the volume of consumption. As far as the corresponding estimation results are concerned, see footnote 6.

2) As far as the ARIMA models are concerned this fact came out when fitting the models into different subperiods of data; as for the behavioral equations, see footnote 7. Analyzing otherwise the direct effects of indexation on household saving is a tedious thing. For instance, the effects of indexation via reduced uncertainty about the rate of return can be shown to be ambiguous (see e.g. Lippman and McCall (1981)).

3) As far as durables are concerned, this classification is problematic because, for instance, transport also includes purchases of motor vehicles. A further problem is the fact that houses and "other durables" are typically treated in a very different way in national accounts; private consumption expenditure includes the flow of services from houses (i.e. gross rent) but not purchases of houses. For "other durables" the opposite is true.

4) We have also tested for the pooling restrictions implied by equations (3) and (4). It turned out that these pooling restrictions could in some cases be rejected, in particular with the ARIMA innovation models. Thus, all of the parameters a_i and b_i do not seem to be constant over different countries. For instance, for equations (1), (2), (5) and (6) in Table 2 the following F-statistics were obtained: $F_{8,243} = 4.107$, $F_{10,233} = 5.082$, $F_{8,243} = 1.590$ and $F_{10,246} = 1.929$.

5) When computing the m- and LM(4)-statistics, the caps in sequence in movements from one country to the next were taken into account.

6) When (3) is estimated applying the "habit formation" hypothesis, i.e. with $(P_{it}/Q_{it-1}/Y_t)$ instead of (P_{it}/Y_t), results change substantially. So, for instance, corresponding of Table 1 the following estimates are obtained:

	(1')	(2')	(3')	(4')	(5')	(6')
(P_1Q1_{-1}/Y)	-.929	-.941	-1.064	-1.050	-.924	-.927
	(12.66)	(13.43)	(11.97)	(12.50)	(12.19)	(12.80)
(P_2Q2_{-1}/Y)	-.896	-.767	-.999	-.834	-.812	-.653
	(3.43)	(3.05)	(3.90)	(3.41)	(3.11)	(2.59)
(P_3Q3_{-1}/Y)	-.980	-.935	-1.080	-.960	-.932	-.982
	(8.64)	(8.04)	(9.07)	(7.83)	(8.00)	(8.53)
(P_4Q4_{-1}/Y)	-.706	-.619	-.859	-.769	-.743	-.617
	(6.49)	(5.80)	(7.34)	(6.77)	(6.68)	(5.70)
(P_5Q5_{-1}/Y)	-.217	-.310	-.448	-.572	-.327	-.378
	(1.36)	(1.97)	(2.64)	(3.46)	(2.00)	(2.42)
(P_6Q6_{-1}/Y)	-.610	-.651	-.806	-.830	-.634	-.677
	(7.99)	(8.87)	(8.23)	(8.97)	(7.98)	(8.89)
$\tilde{p}$			.158	.212	.124	.075
			(2.95)	(3.97)	(2.66)	(1.37)
$\tilde{y}$			-.157	-.129	.009	.007
			(2.73)	(2.38)	(0.29)	(0.24)
r		-.082		-.189		.062
		(0.83)		(1.91)		(1.73)
ΔU		.931		.916		.927
		(4.80)		(4.89)		(4.94)
$(S/Y)_{-1}$	.057	.057	-.066	-.072	.059	.047
	(1.74)	(1.74)	(1.12)	(1.29)	(1.43)	(1.19)
R^2	.9692	.9692	.9679	.9716	.9669	.9702
m	-.070	-.056	-.154	-.133	-.154	-.134
	(0.86)	(0.72)	(1.91)	(1.70)	(1.87)	(1.71)
LM(4)	23.428	19.267	20.099	15.997	21.437	19.216

where symbols are same as in Table 1, for simplicity country dummies have not been displayed. Clearly, the "habit formation" terms $(P_{it}Q_{it-1}/Y_t)$ show very good performance both in terms of the magnitude and precision of the coefficient estimates indicating that relative prices do - after all - affect household saving in the case the structure of pre-commited consumption differs from the structure of actual consumption. Clearly, it is not necessary to consider the "habit formation" specification as a part of the Deaton specification; it is perhaps only the (aggregate) inflation rate variable $\tilde{p}$ which adds something to the "habit formation" model.

7) In this context we arranged a some sort of test for the effects of indexation by scrutinizing whether the coefficients of the unanticipated inflation rate variable, $\tilde{p}_t$, differ between countries. The following F-statistics were then obtained for equations (1), (2), (5) and (6) in Table 2: $F(1)_{4,255} = 2.437$, $F(2)_{4,253} = 3.146$, $F(5)_{4,255} = 1.960$ and $F(6)_{4,253} = 1.852$ while the critical values are 2.41 (5 percent level) and 3.41 (1 percent level). Thus, the assumption on the equality of the coefficients of $\tilde{p}_t$ over countries does not seem to be completely incompatible with data.

8) Juster and Wachtel (1972) and Howard (1978) have also tested the role of anticipated inflation. Howard´s results using data from 5 countries were rather mixed while Juster´s and Wachtel´s results for survey data from USA indicated that anticipated inflation actually reduces household saving!

9) A similar result comes out when (4) is estimated with static expectations, i.e. with $\tilde{p}_t = \Delta \log P_t - \Delta \log P_{t-1}$ and $\tilde{y}_t = \Delta \log y_t - \Delta \log y_{t-1}$. Then the following parameter estimates are obtained for $\tilde{p}$ and $\tilde{y}$ (using equations (5) and (6) in Table 2 as a frame of reference):

	(5')	(6')
$\tilde{p}$	.186	.192
	(3.52)	(3.30)
$\tilde{y}$	.232	.239
	(6.10)	(6.22)

These estimates are clearly much smaller than those obtained with constant expectations (see Table 2); the same is true with the t-ratios. Thus, one can conclude that the performance of Deaton´s specification relies very much on constant expectations, which, in turn, does not really justify speaking about the effects of unanticipated inflation.

10) An analogous result is obtained by Shiba (1978).

11) Cf. also Lawson (1980), who shows that replacing the constant expectations in Deaton´s specification by a Bayesian "adaptive expectations" scheme affects the final estimation results considerably. For instance, the coefficient of the lagged savings ratio goes to unity in this case.

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APPENDIX

Variables and Data Sources

C_i Private consumption expenditure by object in constant prices; data source: OECD Quarterly National Accounts Bulletin 1968-1982, Table 7.

CV_i Private consumption expenditure by object in current prices; data source: same as for C_i.

P_i The implicit deflator of private consumption expenditure.

S Household (including non-profit institutions) saving; data source: OECD Quarterly National Accounts Bulletin 1968-1982, Table 4.

Y Households´ disposable income; data source: $Y = CV + S$.

i Interest rate for long-term bonds; data source: OECD, Main Economic Indicators, various publications.

r The real rate of interest; data source: $r = i/4 - \Delta \log P_t^e$. (With constant expectations r simply equals i; with ARIMA expectations $\Delta \log P_t^e$ corresponds to the predicted value of the respective ARIMA model).

U The rate of unemployment; data source: OECD, Main Economic Indicators, various publications.

The data sample covers the period 1968.2-1982.3 for France and United States, 1968.2-1982.2 for United Kingdom, 1969.4-1982.2 for Australia, and 1970.2-1981.1 for Japan. All the data are seasonally adjusted using the additive X11-adjustment. For Japan and United Kingdom the adjustment was made by the author.

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