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Keskusteluaiheita
Discussion papers

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INFLATION, INCOME DISTRIBUTION AND STABILITY OF THE SAVING FUNCTION: FURTHER RESULTS***

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A b s t r a c t

A 'money illusion' saving function in which unanticipated inflation affects the saving ratio positively is tested. This gets strong support from Finnish data. Goodness-of-fit statistics are quite high, coefficient estimates of expected sign and highly significant. Standard stability tests of both Brown-Durbin-Evans-type and VPR-type are generally rather favourable. Finally introducing the size distribution of income as the additional explanatory variable has the effect of making the performance of the saving function still slightly better.

1. INTRODUCTION

Since late forties, when several pieces of evidence combined to cast doubt on the simple Keynesian consumption function, the life cycle hypothesis (LCH) by Ando-Brumberg-Modigliani and the permanent income hypothesis (PIH) by Friedman have become well-established models of consumption and saving. Apart from aggregation problems,however, they have turned out puzzling at least in two respects.

First, under certain assumptions you can obtain the 'pure' life cycle-permanent income result, according to which consumption is proportional to lifetime wealth at each instant. But the dependence of current consumption on current income would seem to be greater than can be explained by the pure version of the life cycle-permanent income theory. It is the phenomenon of credit rationing that does not only give an explanation of the strong dependence of current consumption on current income, but also leads to the consumption function with a diminishing marginal propensity to consume out of current income and liquid resources. Thus the structure and working of capital markets may play a crucial role in shaping the consumption and saving behaviour.

Second, aggregate consumption and saving functions have generally been assumed to display homogeneity of degree zero in the price level, nominal income and nominal wealth so that their equal percentage change would leave the real consumption and saving unchanged. But consumers typically do not purchase all their goods simultaneously so that they have immediate knowledge only on the prices of goods they are currently buying. Thus, at least in the first instance, consumers do not have sufficient information to distinguish between relative and general price movements, when both are changing simultaneously[1]. If inflation is unanticipated, then consumers find that the goods they are bying are more expensive than expected and misinterpret this to mean a rise in the relative prices of those goods so that the demand for those goods falls. Only if consumers would have enough time to inspect the prices of all goods, then they would find their mistake. Thus there is a mass illusion that all goods are relatively more expensive so that real saving increases when each consumer attempts to adjust his purchases.

Deaton ([7]) has developed a formal model of savings along these lines and found some support for it by using the U.S. and U.K. quarterly data. This paper reports some further developments and tests of the relationship between saving and unanticipated inflation.

First, we use a different sample, namely Finnish quarterly and annual time series data over the period 1959-1977. In the light of differences in the structure of capital markets between these countries it is interesting to test for the operation of 'money illusion' across countries. Empirical results are presented in section 3. Second, since it is well-known that for various reasons behavioural functions may not stay invariant over time, we conduct a large number of various stability tests and apply varying-parameter regression approach on specifications of the saving function. These are reported in section 4. Third, although the major focus is put on the implications of the absence of complete price information, we are also interested in the determinants of the 'equilibrium' saving function, particularly in the role of the size distribution of income. Empirical results from these further experiments are presented in section 5.

2. THEORETICAL BACKGROUND

This section outlines briefly theoretical background for the saving function in the presence of uncertainty about the current and/or future price level (see, Deaton /7/, Deaton and Muellbauer /9/ for details).

Assume that consumers know only the prices of goods they are currently buying, while the current price level is not known with certainty. If preferences are weakly intertemporally separable and if differences in actual and expected prices are entirely due to mistaken expectations about the general price level, then the following expression for the saving ratio (s/y) can be derived as an approximation

$$(1) \qquad (s/y) = (s^*/y^*) + (\log y - \log y^*) - g(\log P - \log P^*)$$

where variables with (*) refer to anticipated values, $y = Y/P$, $y^* = Y^*/P^*$, P = price level, Y = nominal income and where $g = \sum_{k=1}^{n} w_k^* e_{kk}$ is negative quantity for the desired budget shares w_k^* and for own-price elasticities e_{kk}.

The first term on the RHS of (1) gives the 'equilibrium' saving ratio when expectations of real income and prices are fulfilled. According to the second term consumers cannot react to stimuli they do not perceive so that all unanticipated real income is saved. Finally, the third term suggests that unanticipated prices will increase the saving ratio, ceteris paribus.

It remains to specify how the 'equilibrium' saving ratio is determined. A most simplistic view of the consumption function is to assume it to be of type $C = kY^P$, where Y^P = permanent income. Later on we allow for changes in the term k, but now assume it to be constant for simplicity. Using the specification for the determination of permanent income suggested by Deaton ([7]) the 'equilibrium' consumption function can be written as follows

$$(2) \qquad C_t^* = k\left[Y_t^* + m \int_{-\infty}^{t} e^{(t-s)r}(C_s^* - C_s)ds\right]$$

where r = the rate of interest and m = the rate at which wealth is absorbed into permanent income. This suggests that over any finite time period unanticipated income will affect anticipated income and some of the former will be consumed. Since consumers do not know all prices, plans will be revised as new information becomes available. (2) represents a way of specifying this error-correction mechanism.

Using the approximation $d(s^*/y^*)/dt \cong m((1-k)-(s/y))$ for (2) makes it possible to transform (1) into

$$(3) \qquad d(s/y)/dt = m(1-k) + (q-q^*) - g(p-p^*) - m(s/y)$$

where q and p denote rates of change of real income and price level respectively.

In order to estimate (3) on discrete data it has to be integrated over some finite interval (t,t-h), where h may vary. Utilizing the calculations by Deaton ([7]) makes it possible to express the discrete form of (3) as follows

$$(4) \qquad \Delta_h(s/y)_t = b_0 + b_1(q_{ht}-q_{ht}^*) + b_2(p_{ht}-p_{ht}^*) + b_3(s/y)_{t-h}$$

where b_1 and b_2 should be positive, b_3 negative and all between zero and one. Δ_h is the backwards h^{th} difference operator. Moreover, the coefficient estimates of $(q_{ht} - q^*_{ht})$ and $(p_{ht} - p^*_{ht})$ can be expected to decrease, while that of $(s/y)_{t-h}$ to increase, when the length of differencing (h) is increased.

As we noticed, the saving function presented above is based upon the 'disequilibrium' story where uncertainty about the current price level plays a major role. It can be arqued, however, that the most obviuous form of uncertainty about prices is that relating to the future. While information about the current inflation may be quite good when official indices are published frequently, the same does not hold for future inflation. Does this imply the saving function specification different from (4)? Not necessarily!

Suppose that present and future consumption are complements and consider the consequences of the situation in which future price level has been underestimated. Then consumers consume more and save less than they would have desired, had they correctly foreseen the future price level. Consequently, consumer start the 'next' period with smaller assets than they would have liked and increase their saving. It is now unanticipated future inflation that causes saving to rise. Consumption and saving plans are realized, but expectations change from period to period. Under certain assumptions this 'equilibrium' story - LCH with changing expectations - can be shown to lead to the saving function which is unlikely to be identifiable from (4) (Deaton [8] provides some details of this argument).

3. EMPIRICAL RESULTS WITH FINNISH DATA

This section presents empirical results for the saving function specification (4) by using quarterly and annual data from Finland over the period 1959-1976. Time series for consumption has not been constructed in Finland and it lies beyond the scope of this study so that we use private consumption expenditures as our concept of consumption in what follows. The price index to be used is the implicit price index of private consumption expenditures, while the concept of income is a rather wide one, including wages and salaries, social security payments by employers and net income transfers from the public sector.[2)]

Since anticipated rates of changes in real income and price level are not observed variables, links of expectations to observed variables have to be specified. A very simple way of doing this is to assume that they are constant over time so that $q^* = q^0$ and $p^* = p^0$.[3)] Under this hypothesis the OLS estimation results, when quarterly data is used, are reported in Table 1 with different lengths of differencing. Numbers in parentheses are values of t-statistics and ρ = the first-order autocorrelation coefficient obtained by the Hildreth-Lu procedure. The corresponding estimation results, when annual data is used, are reported in equations (9) and (10) of Table 1.

Taking account of the fact that the dependent variable is the difference in the saving ratio, the goodness-of-fit statistics are rather high. All coefficient estimates have right signs and magnitudes and they are highly significant, and there seems to be no serial correlation in error terms.[4)] Moreover, with static expectations the coefficient estimates of q_{ht} and P_{ht} should decrease, while that of $(s/y)_{t-h}$ increase when the length of differencing is increased.

Table 1. Estimation results with different lengths of differencing

equation	constant	q_{ht}	p_{ht}	$(s/y)_{t-h}$	$\bar{R}^2$	D-W	h	ρ
(1)	.090 (5.76)	.591 (5.14)	1.010 (5.88)	-.742 (7.17)	.573	2.203	1/4	0
(2)	.053 (4.10)	.620 (5.64)	.758 (4.61)	-.476 (5.38)	.592	2.073	1/4	-.40
(3)	.078 (4.82)	.477 (5.05)	.545 (4.83)	-.707 (6.72)	.558	2.009	2/4	0
(4)	.077 (4.85)	.476 (5.14)	.542 (4.95)	-.702 (6.72)	.553	1.930	2/4	-.05
(5)	.074 (4.50)	.428 (5.70)	.403 (5.23)	-.722 (6.82)	.599	1.831	3/4	0
(6)	.074 (4.42)	.435 (5.64)	.407 (5.09)	-.722 (6.81)	.590	1.942	3/4	.05
(7)	.098 (5.38)	.323 (4.69)	.336 (5.28)	-.895 (7.77)	.617	1.851	4/4	0
(8)	.098 (5.35)	.325 (4.64)	.337 (5.18)	-.896 (7.77)	.606	1.913	4/4	.03
(9)	.081 (2.13)	.336 (3.14)	.312 (3.01)	-.788 (3.16)	.659	1.572	1	0
(10)	.100 (2.84)	.406 (3.59)	.415 (3.69)	-.978 (4.49)	.673	1.736	1	-.42

Critical t-values: $t_{.05,68} = 1.997$, $t_{.01,68} = 2.653$, $t_{.05,14} = 2.145$, $t_{.01,14} = 2.977$;

lower and upper bounds for D-W are at .05 level: $D\text{-}W_L = 1.52$, $D\text{-}W_U$ - 1.70 (n=72) $D\text{-}W_L = .93$, $D\text{-}W_U = 1.69$ (n=18).

Results in Table 1 do seem to lie in conformity with this pattern. Finally, the coefficient estimates of the constant term and the lagged saving ratio should be negatively correlated which turns out to be the case (see the middle correlation matrix of coefficients in Table 5). Thus the 'money illusion' saving function (4) meets support with Finnish data.

It is interesting to compare results with those obtained by Deaton /7/ by using the same expectations hypothesis and seasonally adjusted data from the U.K. and U.S.[5)] For all there countries the coefficient estimates of unanticipated changes in real income and inflation rate and of the lagged saving ratio are significant and of the expected sign. The performance of the saving function in terms of the goodness-of-fit statistics is better for Finland and the U.K. than for the U.S. The major difference in results lies in the coefficient estimates of the lagged saving ratio. It is much higher (in absolute terms) for Finland than for the U.K. and the U.S. This also 'explains' the higher coefficient estimate of the constant term for Finland, since the coefficient estimates of the constant term and the lagged saving ratio should be negatively related via the term m (reciprocal time horizon). According to the expression (2) the term m describes a mechanism in which replanning occurs in every period when new information becomes available. In the presence of liquidity constraints planning horizons of consumers tend to shorter than with perfect capital markets (Koskela/18/). The different behaviour of the constant and the lagged saving ratio may thus result from differences in the working of capital markets between these countries. We come back to this question later on.

A very simple assumption about g^*_{ht} and p^*_{ht} has been made by keeping them constant, which does not after all sound terribly realistic. Since the hypotheses on the relationship between saving and inflation is based on a theory, in which incorrect inflation-rate expectations play a major role, we have extensively checked the robustness of results to alternative hypotheses of expectations formation. In this connection both adaptive expectations hypothesis and rational expectations hypothesis were used. Results turned out to be qualitatively similar to those obtained by using static expectations. Only coefficient estimates of the unanticipated inflation and the lagged saving ratio were slightly sensitive - while still of the expected sign and significant - to inflation rate expectations specification (results have been reported in Koskela and Virén /20/).[6)]

In order to do some further checks on the saving function specifications we conducted stability tests of various kinds. These are taken up next.

4. THE STABILITY OF THE SAVING FUNCTION

This section reports a number of tests for the 'money illusion' saving function with the purpose of checking whether parameters should be regarded as invariant over time or not. Here the specification with static expectations on q^* and p^*, when the length of quarterly differencing is equal to one, is concentrated on (equation (1) in Table 1). The same stability tests have also been conducted for other specifications with different lengths of differencing and other expectations hypotheses. The results were not too dissimilar with those presented here so that they have not been reported (see Koskela and Virén /20/).

All stability tests are conducted for the estimated equations without the first- and fourth order serial correlation correction because the D-W statistic as well as the Wallis's fourth order serial correlation statistic, d_4, did not show significant serial correlation. The term stability is defined here in the statistical sense of the estimated coefficients of the explanatory variables remaining constant over time so that the null hypothesis H_0 is: $B_1 = B_2 = \ldots = B_T$ and $s_1^2 = s_2^2 = \ldots = s_T^2$, where B denotes the vector of parameters (b_0, b_1, b_2, b_3) and s_t^2 the variance of the error term u_t at time t.

A list of tests used here as well as the relevant references are presented in Table 2, Brown and Durbin and Evans ([2]) and Cameron ([3]) may serve as general references.

Table 2.

name of the test	reference(s)
time trending regression	Brown, Durbin and Evans [2] and Farley and Hinich [10]
cusum test	Brown, Durbin and Evans [2]
cusum squares test	Brown, Durbin and Evans [2]
Harvey-Collier test	Harvey and Collier [14]
Harvey-Phillips test	Harvey and Phillips [15]
Sign-test	Harvey and Collier [14]
Quandt's LR test	Quandt [25], and Quandt [26]
Chow test	Chow [4]
homogeneity test with moving regressions	Brown, Durbin and Evans [2]

Tests listed above differ in terms of their power and alternative hypotheses. It is e.g. well-known that the power of the cusum and Sign-tests is rather weak (Garbade [11], Farley and Hinich [10], Harvey and Collier [14]). In the case of the Quandt's LR test there is a more fundamental problem, which is due to the fact that the distribution of the likelihood ratio, λ_r, under H_0 is not known so that no rigorous test can be derived from it. Anyway, in order to give some idea of the behaviour of parameters over a certain period r, we have computed the corresponding χ^2-statistic, originally proposed by Quandt [25]. Finally, it is worthwhile to point out that the Harvey-Collier test is strictly speaking a test for functional misspecification, and the Harvey-Phillips test is a test for heteroscedasticity of residuals.

Before considering these formal stability tests we may get some useful information about the point in time when any shifts in the regression relation might have occured by computing Quandt's log-likelihood ratio statistic λ_r. This is described in Figure 1. The period r = 44, i.e., 1969 (IV) can be clearly discerned. We use this "a priori" information by computing the Harvey-Phillips test, the Quandt's LR test and the Chow test just with respect to this 'switchpoint period' r = 44. A great deal of variability can be seen in the first values of r. No strong conclusions from this cannot, however, be drawn because the power of the test dimishes when the end-points of the data sample are approached (see Quandt [26]). Also the period of 'oil crisis', 1973 (IV) can be distinguished from the λ_r graph. No stability test statistic, however, computed with respect to this point turned out to be significant.

Stability test statistics are presented in Table 3, in which the column with k=4 correspond to equation (1) in Table 1 (columns labelled by k=3 and k=5 are discussed later on). Letters (f) and (b) have been added whenever recursive residuals and tests based on them have been calculated on forward or backward forecast errors respectively, and H_1 and H_2 refer to the regression relationships when parameters are first-degree and second-degree polynomials in time, and n indicates the length of the time segment used in moving regressions.

Our saving function specification seems to past stability tests rather well, even though there are some mixed results.[7)] The cusum squares and the Harvey-Phillips tests show significant test-statistics. The cusum squares test showing instability, unlike the cusum test, can be interpreted as suggesting that instability is due to a shift in the residual variance rather that to shifts in the values of regression coefficients. The performance of the Harvey-Phillips test-statistic lies also in conformity with this interpretation. This should not be regarded as surprising when one takes into account the sharp shifts in the saving ratio in the 70's, which in turn help to 'explain' the drop in the performance of the saving function in terms of goodness-of-fit statistic, $\bar{R}^2$ (from .624 over the period 1959-1969 to .565 over the period 1970-1976).

What cannot be explained by the shifts in the residual variance, however, is the value of homogeneity statistic for 12 observation moving regression. This result should not be overemphasized partly because of the very short time segment and partly because the value of the test statistic is only just significant at the 5 per cent level.

Thus far, results of various stability tests have been rather favourable to our specification of the saving function. Stability tests reported, however, suffer from a couple of weaknesses for which reason we like to do some further checks.

TABLE 1 QUANDT'S LOG-LIKELIHOOD RATIO.

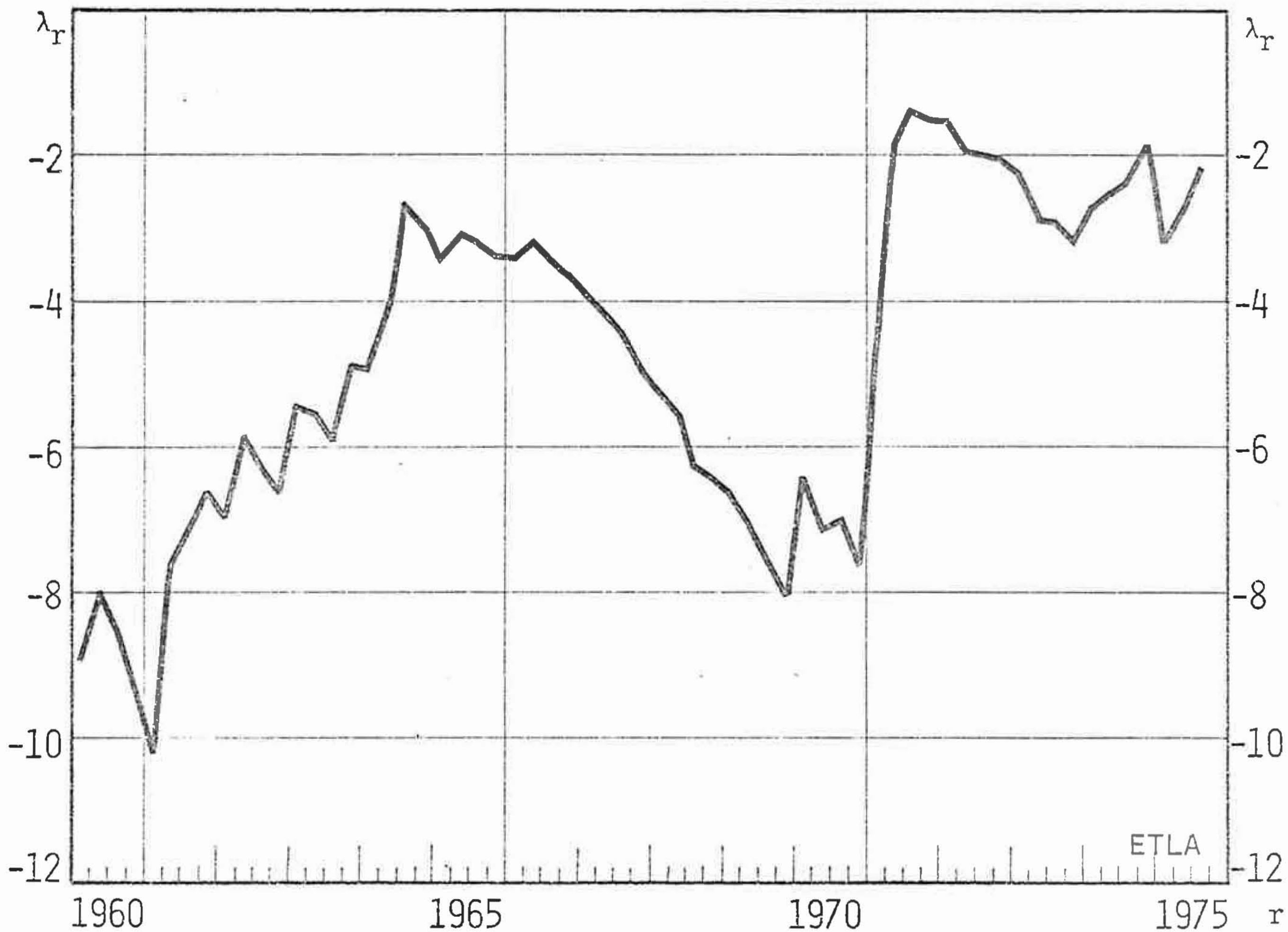


Table 3. Stability test statistics

test statistic	k=4	k=3	k=5	critical test values 5 %/1 % (k=4)
time trending regressions				
$F(H_1$ vs $H_0)$	1.223	.522	1.266	2.51/3.62
$F(H_2$ vs $H_0)$	1.409	.530	1.379	2.10/2.82
cusum test (b)	.320	.481	.627	.948/1.143
cusum test (f)	.502	.343	.465	.948/1.143
cusum squares test (b)	.254	.370	.247	.149/.181
cusum squares test (f)	.285	.347	.280	.149/.181
Harvey-Collier test (b)	-.753	.606	-1.372	1.998/2.650
Harvey-Collier test (f)	.565	-1.004	.549	1.998/2.650
Harvey-Phillips test,				
r=44 , (b)	2.653	4.338	2.759	1.98/2.49
(f)	.304	.233	.311	1.79/2.30
Sign-test (b)	.970	.181	.916	1.96/2.58
Sign-test (f)	.243	.542	.550	1.96/2.58
Quandt's log-likelihood ratio, r=44	36.89	42.21	38.64	12.59/16.81
Chow-test, r=44	1.522	.472	1.303	2.52/3.65
homogeneity statistics for moving regressions				
n=12	1.890	.800	1.535	1.79/2.30
n=24	1.208	.471	1.004	2.25/3.12
n=36	1.396	.399	1.199	2.52/3.65

First, it may be argued that they are based on an inappropriate alternative hypothesis. Particularly, this is be true with the Quandt's LR test and the Chow test in which the alternative hypothesis supposes that coefficients are constant within a sub-period, but change by a discrete amount across sub-periods. If this is not the case, then it is appropriate to re-estimate coefficients with a methodology that allows continous but gradual time variation in the coefficients. Second, the stability tests conducted are typically the tests for the stability of the whole regression relationship so that the source of variability is not identified. Varying parameter regression (VPR) technique provides one approach to both of these problems by making it possible to estimate time-varying coefficients. Moreover, it is important to point out that the simulation results presented by Garbade /11/ show that VRP techniques are generally more powerful than methods based on recursive residuals.[8)]

Allow now for the coefficient vector of the 'money illusion' saving function, $B = (b_0, b_1, b_2, b_3)$, to be subject to sequential variation over time. There are many possibilities to try to model the evolution of B_t, but we suppose that it follows a random walk with zero drift through time. Garbade /11/ shows that this assumption is fairly robust to misspecification as long as the true dynamics of the B_t series is highly autocorrelated. The saving function with static expectations and with h = 1/4 can now be written as

(5a) $$\Delta_{\frac{1}{4}}(s/y)_t = b_{0t} + b_{1t}q_{\frac{1}{4}t} + b_{2t}p_{\frac{1}{4}t} + b_{3t}(s/y)_{t-\frac{1}{4}} + u_t$$

(5b) $$u_t \sim N(0,s^2)$$

(5c) $$B_t = B_{t-1} + e_t$$

(5d) $$e_t \sim N(0,s^2Q)$$

The random vector e_t is assumed to be serially uncorrelated and uncorrelated with the random scalar u_t. Q is a relative covariance matrix for e_t with s^2 as scaling variance. Both Q and s^2 are unknown and have to be estimated. Thus the parameter process is assumed to incorporate permanent and transitory changes. E.g. if $Q_{00} = Q_{11} = Q_{22} = Q_{33} = 0$, then B_t does not vary over time and VPR collapses to ordinary regression problem.[9)]

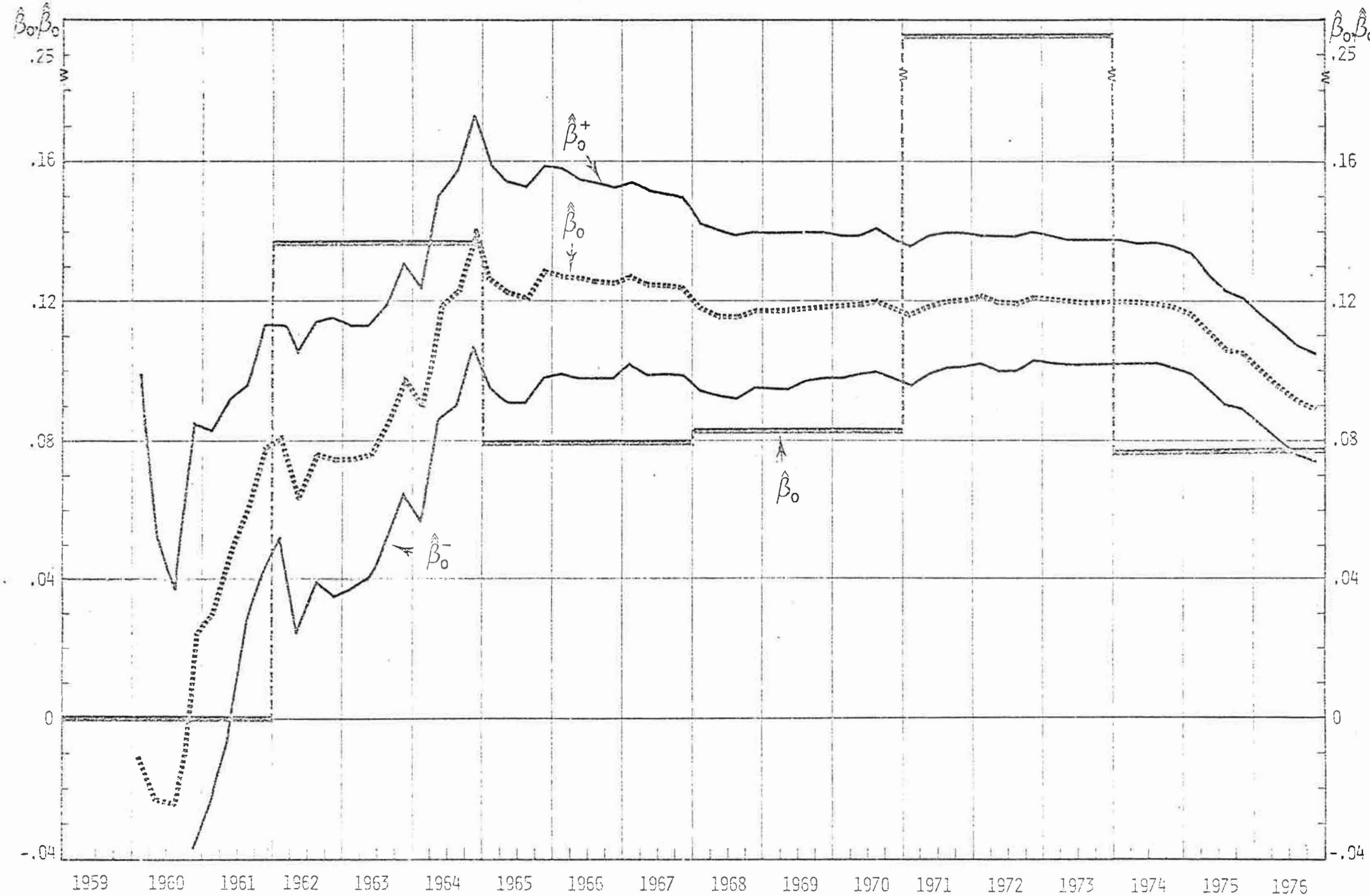
The respective likelihood function was maximized with the following diagonal terms of the relative covariance matrix Q: $Q_{00} = .007$, $Q_{11} = Q_{22} = Q_{33} = 0$, all the off-diagonal terms being equal to zero.[10)] Moreover, it was found that the corresponding innovation terms, $w_t = \Delta(s/y)_t - x_t\hat{B}_{t|t-1}$, where x_t denotes the vector of explanatory variables and $\hat{B}_{t|t-1}$ the estimate of B_t based on observations $\Delta(s/y)_0, \Delta(s/y)_1, \ldots, \Delta(s/y)_{t-1}$, driving the Kalman filter were white noise (according to the Box-Pierce Q-statistic with five lags ($Q_5 = 9.44 < 11.07 = \chi^2_{.05,5}$)). Hence the optimality of the Kalman filter is guaranteed (see Mehra /21/, p. 177).

Thus the constant term may be subject to permanent time variation implying shifts over time, while this would not seem to be the case with other coefficients. Their variability could then be characterized as being due to transitory time variation with a fixed permanent component according to this interpretation. Of course, still the possibility remains that the result concerning the relative covariance matrix Q is due to functional misspecification of the saving function. We cannot, however, discriminate between these two hypotheses.

In order to test the significance of permanent time variation the standard LR test with the null hypothesis H_0: $Q_{00} = 0$ (under the maintained hypothesis that the other diagonal elements are zero) is used. The likelihood ratio statistic $-2\ln\lambda$ turns out to be .396, which is not significant at the 5 per cent significance level for χ^2-distribution with one degree of freedom. With $Q = 0$, however, the LR statistic will be more concentrated toward the origin than χ^2-distribution. Thus using this distribution to determine the critical values of $-2\ln\lambda$ leads to a conservative test of stability (see Garbade /11/, p. 55-56).

Anyway, we computed the smoothed estimates of b_{0t}, denoted by $\hat{\hat{\beta}}_0$, which correspond to the maximum likelihood covariance matrix with $Q_{00} = .007$ (under the maintained hypothesis that the other diagonal elements are zero). These estimates, bordered by plus and minus one standard error of estimates, are presented in Figure 2 together with the OLS piecewise estimates of b_0 across different sub-periods of 12 quarters, denoted by $\hat{\beta}_0$. The smoothed estimates of b_{0t} indicate some variation of the constant term over time. The constant term consists of the 'equilibrium' saving ratio, $1-k$, the term m as well as the expected values of real income and price level changes so that variation is due to one of these components. In the light of the behavior of the standard errors during the period, 1960-64, it is not completely unjustified to say that the constant term has been rather stable over the whole observation period. On the other hand, the piecewise regression estimates of b_0 behave differently, particularly over the period 1971-1973, being obviously a result of some transitory elements. In this connection the null hypothesis of parameter stability can be rejected when piecewise linear variation in B_t is the alternative hypothesis (computing the Chow statistic for 6 nonoverlapping periods with 12 observations in each gives the value of F-statistic $2.910 > 2.530 = F_{.05,4,60}$). A similar result was already given by the homogeneity statistic with 12 observations (see Table 3).

Figure 2. Smoothed and piecewise estimates of the constant term



5. SOME FURTHER EXPERIMENTS

On the whole, the picture that emerges from stability analyses conducted in earlier section, is rather favourable to our specification of the saving function. The lagged saving ratio, while being slightly sensitive to specification of expectations hypotheses on q^* and p^*, has played a major part in it. Recently, however, Lawson [22] has argued for the notion that the current desired saving ratio would be determined independently of discrepancies between past saving ratios and their desired levels. We start this section by evaluating the role of the lagged saving ratio in equation (4).

On the other hand, the VPR-application can be interpreted as suggesting that to the extent that there are permanent drifts in coefficients over time, then the constant term would be the most likely candidate for them. We end this section by reporting some empirical experiments, in which the constant term k of the basic specification (4) is allowed to change.

5.1. The lagged saving ratio as an error-correction mechanism

As has been pointed out earlier, the term m describes a mechanism by means of which replanning occurs at every period when new information about present and/or future price levels becomes available. Suppose now that such a replanning does not occur so that the lagged saving ratio is dropped out from the saving function (4) (m=0). The representative set of results with and without the $(s/y)_{t-h}$ term are reported in Table 4.

Dropping the lagged saving ratio is seen to have remarkable consequences. First, the performance of the saving function, measured by $\bar{R}^2$, falls dramatically. The null hypothesis H_0: $b_3 = 0$ is definitively rejected by using F-test in all cases. Second, there are now serious problems in error terms. The saving function without the lagged saving ratio suffers not only from serial correlation (see Table 4), but also from the strong heteroscedasticity of residuals. This is to be seen by looking at the stability tests presented in Table 3 (column labelled by k=3), in which the cusum squares and the Harvey-Phillips tests show clear instability. Finally, the coefficient estimate of the unanticipated inflation becomes smaller and insignificant and the constant term drops in magnitude and the correlation matrix of coefficients behaves as expected (see the downmost column of Table 5).

All in all, dropping the lagged saving ratio from (4) leads to very poor results. Most of the power of the saving function disappears. We do not regard this as surprising, since in the presence of uncertainty about the current and/or future price levels replanning, as new information becomes available, would quite naturally seem to be an essential ingredient of the story, particularly in Finland, where credit rationing has occasionally taken place (see Tarkka [29]).

Table 4. Estimation results with and without the constant and (s/y)-terms

equation	constant	$q_{-\frac{1}{4}t}$	$q_{\frac{1}{4}t}$	$(s/y)_{t-\frac{1}{4}}$	$\bar{R}^2$	D-W	F_H *)
(1)	.090 (5.76)	.591 (5.14)	1.010 (4.88)	-.742 (7.17)	.573	2.208	..
(2)	-.017 (2.98)	.763 (5.15)	.486 (1.91)		.261	2.984	48.892
(3)		.809 (6.15)	.852 (3.44)	-.171 (4.79)	.374	2.863	33.116
(4)		.558 (4.03)	-.099 (0.58)		.177	2.863	34.748

*) F_H indicates the test statistic for the null hypothesis H_0: b_0 and/or $b_4 = 0$, the critical test values are: $F_{.01,2,68} = 4.97$ and $F_{.01,1.68} = 7.06$.

Table 5. Correlation matrices of coefficients

	$\hat{b}_0$	$\hat{b}_1$	$\hat{b}_2$	$\hat{b}_3$	$\hat{b}_4$
$\hat{b}_0$	1.000				
$\hat{b}_1$	-.259	1.000			
$\hat{b}_2$	-.282	.167	1.000		
$\hat{b}_3$	-.412	.199	-.368	1.000	
$\hat{b}_4$	-.846	.098	.411	-.113	1.000
$\hat{b}_0$	1.000				
$\hat{b}_1$	-.331	1.000			
$\hat{b}_2$	.135	.133	1.000		
$\hat{b}_3$	-.958	.212	-.355	1.000	
$\hat{b}_0$	1.000				
$\hat{b}_1$	-.457	1.000			
$\hat{b}_2$	-.770	.228	1.000		

A correlation coefficient .235 (.306) is just significant at .05 (.01) level with 68 df.

5.2. Income distribution and the saving function

Thus far the term k in $C = k\,Y^p$ has been assumed to be constant for simplicity. We allow now for its variability over time without yet specifying its determinants.

Differentiating the equation (2) (p.4) with respect to time and using the approximation $(s/y) \cong \log Y - \log C$ yields the approximation $d(s^*/y^*)/dt \cong m((1-k)-(s/y)) - d(\log k)/dt$ for (2). Finally, differentiating the saving ratio equation (1) (p.4) with respect to time and utilizing the expression for $d(s^*/y^*)/dt$ leads up to

$$(6) \qquad d(s/y)/dt = m(1-k) + (q-q^*) - g(p-p^*) - m(s/y)-n$$

where n denotes rates of change of k.

Next we have to specify the variable(s) on which the desired consumption-permanent income ratio k depends. A number of explanations have been suggested to account for this, like the real rate of interest, uncertainty effects of inflation, real balance or Pigou-effect and the size distribution of income. In what follows we focus on the size distribution of income partly because of the lack of data to evaluate alternative explanations and partly because there is some preliminary evidence with Finnish data on the behalf of this explanation (see Koskela and Sullström /19/).[11)]

Assuming that k depends on the size distribution of income D e.g. linearly and transforming (6) to a form, appropriate on discrete data, leads up to the following specification of the saving function

$$(7) \qquad \Delta_h(s/y)_t = b_0 + b_1 q_{ht} + b_2 p_{ht} + b_3 (s/y)_{t-h} + b_4 D_{ht} + b_5 d_{ht}$$

where d denotes rate of change in D and where static expectations on q^* and p^* have been assumed. The coefficients b_1, b_2, b_4 and b_5 should be positive - if a rise in D means a rise in the inequality in the size distribution of income - and b_3 negative.

To the extent that the role of the size distribution of income in the saving function is due (at least partly) to credit rationing, b_1 should decrease and $-b_3$ increase when the income distribution variable is introduced. Under these circumstances there is a tendency for planning horizons of consumers to become shorter and unanticipated changes in real income to affect, not saving, but consumption.

We substituted aggregate 'indices of inequality' for the terms D_{ht} and d_{ht} and picked up from a large number of inequality measures available such ones which could be regarded as representing a certain aspect to inequality of income distribution. By this means we tried to ensure that results do not come from a certain special way of measuring inequality. The following measures of inequality were used: the coefficient of variation (V), the Gini-index (G) and the ratio of harmonic to arithmetic mean income (H), the so-called Dalton's h.[12)] Here we report only results with V and G.

The inequality indices were only available on annual basis. We also constructed corresponding indices of inequality on quarterly basis by deriving quarterly variations of these indices within a year from the (seasonally adjusted) unemployment rate. It turned out that rates of changes in inequality indices were statistically insignificant both with and without their levels, and sometimes they had also different signs from the level variable. Therefore, we do not report results in which rates of changes in inequality indices have been used as explanatory variables.[13)]

Results by using annual data are presented in equations (1)-(6) of Table 7. Comparing these with the corresponding results of Table 2 (p. 10) suggests that the performance of equations, measured by $\bar{R}^2$, is increased by introducing inequality measures as explanatory variables, but the increase is so slight that the null hypothesis H_0: $b_4 = 0$ cannot be rejected ($F(1) = 3.414 < 4.67 = F_{.05,1,13}$). This is not, however, a 'fair' test since in equation with a constant term and an income distribution variable the latter is more a substitute for the former than an additional explanatory variable. Again dropping the lagged saving ratio decreases the performance of the saving function sharply and gives rise to serial correlation. Even though H_0 cannot be rejected, other evidence is favourable for the role of income distribution variable: estimates of b_1 and b_3 behave as expected; estimates of the constant term and the inequality index should be negatively correlated which is the case (see the upmost correlation matrix of coefficients in Table 5); income distribution variables are of expected sign and significant (except G_t in equation (2)) at the 5 per cent level.[14)]

As far as the results with quarterly data - also presented in Table 6 - are concerned, they do not contradict with the evidence obtained by using annual data. Comparing estimates with the corresponding estimates in Table 1 shows that introducing the income distribution variable has the effect of decreasing the size and value of t-statistic for the constant term as one would expect. On the other hand, inequality indices, while having 'right' signs, are significant only when the constant term is dropped.[15)]

5. CONCLUDING REMARKS

We have analyzed the relationship between aggregate private saving, inflation and income distribution by starting from the notion that real economic decisions depend on relative prices, but consumers do not in reality have enough

Table 6. Estimation results with income inequality indeces

equation	constant	q_{ht}	p_{ht}	$(s/y)_{t-h}$	V_{ht}	G_{ht}	$\bar{R}^2$	D-W	h
(1)	-.079 (1.14)	.309 (2.84)	.525 (4.19)	-.924 (3.83)	.157 (2.62)		.709	1.951	1
(2)	-.201 (1.48)	.328 (2.86)	.377 (3.55)	-.919 (3.54)		.641 (2.13)	.671	1.558	1
(3)		.291 (2.68)	.468 (4.04)	-.987 (4.15)	.098 (3.18)		.703	1.553	1
(4)		.321 (2.68)	.350 (3.22)	-.900 (3.33)		.213 (2.46)	.642	1.372	1
(5)	-.139 (1.47)	.501 (3.69)	.236 (1.68)		.092 (1.14)		.426	2.613	1
(6)	-.177 (0.97)	.511 (3.69)	.147 (1.30)			.305 (0.79)	.399	2.356	1
(7)	.065 (2.22)	.604 (5.22)	1.106 (4.83)	-.754 (7.23)	.022 (0.98)		.572	2.233	1/4
(8)	.064 (2.07)	.605 (5.22)	1.039 (4.97)	-.761 (7.23)		.060 (0.98)	.573	2.228	1/4
(9)		.673 (5.87)	1.260 (5.60)	-.660 (6.74)	.065 (5.24)		.548	2.443	1/4
(10)		.670 (5.86)	1.063 (4.98)	-.672 (6.77)		.032 (5.33)	.552	2.428	1/4

Critical t-values: $t_{.05,13} = 2.160$, $t_{.01,13} = 3.012$, $t_{.05,67} = 1.998$, $t_{.05,67} = 2.654$.

Lower and upper bounds for D-W are at .05 level: $D\text{-}W_L = .82$, $D\text{-}W_u = 1.87$ (n=18), $D\text{-}W_L = 1.49$, $D\text{-}W_u = 1.74$ (n=72).

information to distinguish between relative and general price movements. This suggested a 'money illusion' saving function, in which unanticipated inflation affects the (difference in) saving ratio positively.

Empirical results with Finnish data can be briefly summarized as follows: The goodness-of-fit statistics were quite high, coefficient estimates of expected sign and highly significant. In order to do some further checks we conducted standard stability tests of both Brown-Durbin-Evans type and of VPR-type. On the whole, results of various stability tests were rather favourable to the specification of saving function. There seemed to be, however, signs of instability due to a shift in the residual variance and/or the constant term over the whole period. Therefore we finally carried out some further experiments by allowing the consumption-permanent income ratio to vary over time. More particularly, we introduced the size distribution of income as the additional explanatory variable. On the whole the evidence was favourable - even though not too strong - for the role of income distribution variable in the saving function.

Among agendas for further research we mention one. The major difference in results with the 'money illusion' saving function specification between Finland, the United Kingdom and the United States was in the coefficient estimates of the lagged saving ratio which were much higher in absolute terms for Finland than for the U.K. and the U.S. Its size depends on the mechanism by means of which consumers' replanning occurs in every period when new information about the current and/or future price levels becomes available. Under imperfect capital merkets there is a tendency for planning horizons to become shorter when credit markets become 'tighter' and vice versa. Allowing for the dependence of error-correction mechanism on the 'tightness' of credit markets is an area where research is needed.

FOOTNOTES:

1) For empirical evidence on the positive relationship between the variability in the general level of prices and in the relative prices, see e.g. Cukierman and Wachtel /5/.

2) The quarterly data has been constructed in the context of the quarterly model of the Bank of Finland and it is only available from the period 1958 (I) on. Sources of annual data are: National Accounting in Finland in 1952-1964, Statistical Reports No. 53. Helsinki 1968; National Accounts in 1965-1977, Helsinki 1978. All published by the Statistical Central Office. Quarterly data is seasonally adjusted.

3) This specification was used by Deaton /7/. Note that now k, q^0 and p^0 cannot be separately identified.

4) We also computed the following regression from residuals, u_t: $u_t = a_1 u_{t-(1/4)} + a_2 u_{t-(4/4)} + e_t$ and postulated the null hypothesis:

H_0: $a_1 = a_2 = 0$. The following F-statistics were obtained for various

equations of Table 1: (1) 1.198, (3) .263, (5) .134 and (7) .118. The corresponding critical value at the 5 % significance level is 3.15 so that the null hypothesis could be rejected in no cases.

5) The results by Davidson et. al (/5/) lie in conformity with those obtained by Deaton (/6/) as far as the British evidence on the relationship between saving and inflation is concerned.

6) Some experiments were also carried out with a so called threshold hypothesis. A most simple version of this is to suppose that if the inflation rate is lower than some level, say p_T, then consumers can distinguish between relative and general price movements, while in the case of $p > p_T$ the earlier argument holds. Preliminary results were encouraging (see Koskela and Virén /20/). This would seem to be an interesting area for further research.

7) For applications of these test procedures, see e.g. Cameron /3/, Hackl and Katzenbeisser /13/ and Kahn /17/.

8) For various kinds of VPR-applications, see e.g. Garbade and Wachtel /12/, Laumas and Mehra /21/, and Rausser and Laumas /27/.

9) We are interested in obtaining the so-called filtered and smoothed estimates of B_t (see Garbade /11/ for details). The filtered estimates of B_t make it possible to evaluate the nature of an instability in individual coefficients. They do not, however, use forward observations and may thus be quite erratic in early periods of the sample interval. Therefore, the smoothed estimates may be more informative (Garbade/11/).

10) Our computer program, RSTOCH/RECUCHK, did not allow for the direct estimation of Q. Estimation was performed with different Q's and the value of Q was chosen, which led to the maximum of the log-likelihood

$$(1) \qquad L = -\left(\frac{1}{2}\right)\{(T-p)\log \sum_{s=p+1}^{T} w_{s/s-1}^{2}/(1+x_s M_{s/s-1} x_s') + \sum_{s=p+1}^{T} \log(1+x_s M_{s/s-1} x_s')\}$$

where $w_{s/s-1}$ denotes a prediction error. The covariance matrix is defined $s^2 M_{s/t} = E(\hat{b}_{s/t}-b_s)(\hat{b}_{s/t}-b_s)'$ and p is the first period for which the filtered estimate $\hat{b}_{p/p}$ is computed. Thus the program carries out a required number of estimations within a given range of Q and the Q which leads to the highest value of (1) is picked up (see Mäkeläinen /24/, for the details of this computer program).

11) More particularly, there is some supporting evidence for the hypothesis that equalizing the size distribution of income would increase aggregate consumption. On the other hand, tests with the U.S. data suggest that equalizing the size distribution of income would either leave the aggregate consumption unchanged or diminish it slightly (see Blinder /1/).

12) These inequality indices are reported in Koskela and Sullström /19/. V is most sensitive to inequality associated with 'exceptionally rich', H to inequality associated with 'poverty' and G is most sensitive to inequality associated with a wide spread of the less extreme incomes without much tendency for the majority of them to be bunched within quite a narrow range.

13) A full set of result is available upon request.

14) The inequality index H did not show up very well; it had even the 'wrong' sign and it was statistically insignificant. But this measure is particularly sensitive to inequality associated with 'poverty' so that estimates with H might be interpreted as suggesting that k tends to be constant at 'low' levels of income (Koskela and Virén /20/).

15) It has been argued by Hendry and Ungern-Sternberg /14/ that conventionally measured disposable income is not a good approximation to the economic definition of income, since the real value of liquid assets is not being held constant (see also Siegel /28/. This hypothesis provides a way of explaining the saving behaviour since inflation correction removes an upward trend in the saving ratio. It is obvious, however, that the effect of inflation on disposable income goes partly via the income distribution (See Deaton /8/ for a critique of this mismeasurement hypothesis).

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