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ABSTRACT

One of the important determinants of the response of consumption and saving to the expected real interest rate is the elasticity of intertemporal substitution. Knowledge about the size of this structural parameter is required if the effectiveness of policy interventions, or business cycles, is to be studied by means of macro models. Rather robust evidence from Finnish data suggests that this substitution effect is very small. In the light of this study there does not seem to be much scope for a policy aiming e.g. at making consumers defer consumption by raising their real interest rate expectations.

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CONTENTS

I INTRODUCTION

The effectiveness of monetary policy depends, among other things, on the role of the interest rate in the determination of consumption, saving and investment. While the interest rate sensitivity is subject to continuous theoretical as well as empirical debate, many recent studies of consumption have modelled observed fluctuations as the outcome of optimizing decisions of a representative individual. A basic insight of this so called Euler equation approach is that the expected after-tax real interest rate influences intertemporal choice. A higher interest rate will, ceteris paribus, make the consumer defer consumption. The magnitude of this intertemporal substitution effect is one of the central questions in modern neoclassical macroeconomics.[1]

The Euler equation approach allows a direct estimation of utility function preference parameters without requiring explicit solutions to the consumer's dynamic optimization problem. The intertemporal substitution elasticity can be measured e.g. by the response of the rate of change of consumption to changes in the expected real interest rate. This paper attempts to estimate the elasticity of intertemporal substitution from Finnish data. The analysis will essentially utilize the framework of HALL (1985), who in a recent study using SELDEN's (1978) approach designed and applied a consistent estimation procedure based on the first-order conditions characterizing the optimal consumption plan. The realism of the setting will also be evaluated.

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[1] The extent to which high interest rates reduce consumption is crucial in macroeconomic models assessing the effectiveness of monetary policy. If consumers can be induced to postpone consumption by modest increases in interest rates, then movements of interest rates will make consumption decline whenever other components of aggregate demand rise. Consumption will also move along with real interest rates over the business cycle. The interest sensitivity of consumption will furthermore be crucial in studies evaluating the effects of public borrowing and the tax on interest income. It may also affect estimates of the proportion of total capital stock that is composed of life cycle saving, and how variations in aggregate consumption affect the stock market (see e.g. HALL (1985) and SKINNER (1985) for discussions and further references).

II THEORY

In the literature on consumption and stochastic asset returns the consumer is usually viewed as maximizing the expected value of an intertemporal utility function in which the intertemporal elasticity of substitution is just the reciprocal of the relative risk aversion. If the consumer is highly risk averse, he must have low intertemporal substitution as well; it is not possible to prescribe risk and time preferences separately in this framework.

SELDEN (1978) has developed a more general framework in which it is possible to distinguish between the roles played by risk preferences and time preferences in determining optimal consumption and asset demand (see also SELDEN (1979, 1980)). The approach consists of a generalization of the hypothesis of expected utility maximization over bivariate random variables, or in our context more specifically a period one certain consumption and a period two uncertain consumption. Selden's "ordinal certainty equivalent" (OCE) representation hypothesis is based on a conditional period two (single-attribute) expected utility function and a two-period ordinal time preference index (in contrast to the traditional two period (multiattribute) expected utility model). The OCE representation includes the two-period cardinal expected utility paradigm as a limited special case.

In a world with stochastic asset returns one of the choices facing the consumer is to spend a little less in period one, invest the savings in one asset, and spend the stochastic proceeds in period two. In a recent paper HALL (1985) generalizes his earlier work (HALL (1978)) by relaxing the assumption of a constant real interest rate, and by modelling the risk and time preferences separately as constant, but independent parameters.[2] Assuming an interior

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[2]The former assumption has in fact already been relaxed in earlier papers by e.g. HALL (1981), MANKIW (1981), MUELLBAUER (1982) and WICKENS & MOLANA (1982).

consumption optimum, a stochastic Euler equation will be satisfied, which is shown to imply the following simple and straightforward relation between consumption c and the interest rate r

$$(1) \qquad \Delta c_t = \sigma r_{t-1} + k + \varepsilon_t, \qquad \sigma > 0 \wedge \sigma \neq 1$$

where Δ denotes a log-change, k is a constant that depends on the variances and covariances of c and r, and ε is a normal random variable. r is to be thought of as the mean of the typical consumer's subjective distribution for the real interest rate at the time consumption decisions are made for period t-1. The coefficient σ that governs the influence of the expected real interest rate is precisely the elasticity of intertemporal substitution. A high value of σ means that when the real interest rate is expected to be high, the consumer will defer consumption to the later period. It should perhaps furthermore be emphasized, that the bivariate relation between consumption and the real interest rate does not reveal anything about risk aversion in the OCE framework.

The basic equation (1) for the rate of change of consumption can be estimated from data on consumption and expected real after tax interest rates. A strong testable implication of the theory and the adherent assumptions is that the mean of the rate of growth of consumption is shifted only by the mean of the real interest rate. In the regression, no other variable known in time t-1, or earlier, should help predict the rate of growth of consumption. In order to evaluate the validity of the choice of simplifying assumptions (e.g. various separability assumptions and absence of constraints in capital and labour markets) this orthogonality condition will be carefully tested.

III DATA ISSUES

There are many aspects to the choice of data for this study. These can mainly be related to the measurement of the relevant consumption aggregate and expected real interest rate, and to aggregation issues. As the objective is to estimate a single preference parameter in a representative individual's utility function, real per capita consumption of nondurables and services was chosen as the primary aggregate. Per capita terms were chosen in order to make the data (loosely) consistent with a representative consumer notion, and in order to partly correct for demographic movements.[3] Durables were excluded from the measure of consumption for two reasons. First, it is rather difficult to impute a service flow to the stock of durables, and second, as shown by MANKIW (1982), consumption would follow more general ARMA processes rather than the simple AR(1) process implicitly assumed if durables are included. HALL (1985) uses plain consumption of nondurables, and for sake of comparison this measure will also be employed.[4]

Since no survey data on expected real interest rates in Finland exist, the measurement of these expectations is one of the empirical issues that have to be dealt with in this paper. As nominal interest rates HALL (1985) uses returns from a savings account earning the regulated passbook interest rate, returns from common stock, and returns from government bonds. The choice clearly is a difficult

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[3]While the underlying trend in the population estimate might be close to a trend in the true series, the quarter-to-quarter changes probably contain a lot of noise. In order to evaluate how serious this might be experiments with consumption not deflated by population will be conducted.

[4]HANSEN & SINGLETON (1983) have drawn attention to the fact, that by estimating models with nondurables and services and nondurables one is implicitly considering two different assumptions about the separability of preferences. Similarly, the choice among asset returns will amount to choosing among different models of the return generating process. Different time periods corresponding to different monetary policy regimes may furthermore yield different real rate processes (e.g. HUIZINGA & MISHKIN (1985)).

one, but in the present study the average lending rate of banks r_l is chosen as the primary interest rate facing the representative consumer.[5] Following Hall the average deposit rate r_d (as well as the average return on high interest time deposits r_{td}), the effective yield on common stock r_s as well as the interest on (medium to long-term) government bonds r_b will be used. All yields are tax-exempt, except for the yield on common stock. This yield was consequently corrected with one minus the effective marginal tax rate.[6]

Three rough proxies for the expected rate of change of the price level where constructed. In the case of perfect foresight realized quarter-to-quarter (year-to-year for r_{td}) log-changes (as decimals at an annual rate) in the respective implicit deflator for consumption were utilized. These deflators were also used for all deflated variables. Furthermore static and autoregressive (AR(4)) expectations were employed. Thus three times five different proxies for the expected real after-tax interest rate were obtained.

It might be noted that the differences between on the one hand different expectations models and on the other hand different nominal yields were not overwhelmingly big, and that all series followed roughly the same (markedly autoregressive) pattern over time. Furthermore, all series were everything else but constant, and no trend was visible for the whole period considered. All data are seasonally adjusted quarterly data from the Bank of Finland. At one occasion it was possible to utilize unadjusted data as well (see fn. 17).

Finally, the potentially important bias inherent in studies of this sort related to time aggregation ought to be pointed out. The Euler

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[5] If, however, there has been excess demand for loans this interest rate will underestimate the relevant interest rate.

[6] The effective marginal tax rate was calculated as in WILLMAN (1985). In the sample it varies between .20 and .38.

equation involves consumption at two points in time, but due to data limitations one is reduced to using time averages of consumption. If the length of the period used by the consumer for planning purposes differs from that used in the estimation, time aggregation biases will be present.

IV EMPIRICAL RESULTS

Some basic analyses

In this section empirical estimates of the elasticity of intertemporal substitution σ based on equation (1) are presented (it is assumed that k does not change significantly over time). In order to get a picture of the general fit of the model and of the constancy of the preference parameter equation (1) is first estimated using ordinary least-squares (OLS).[7] As there are good reasons to doubt the accuracy of the OLS-estimate of σ, too much attention to the size of σ should, however, not at this stage be paid.

Estimation results are presented in Table 1. The estimated constant, the squared multiple correlation coefficient and the Durbin - Watson test statistic are included as purely descriptive diagnostics. Data cover (at the most) the period 1961IV-1984IV. The empirical evidence of (1) as a specification is not altogether conclusive; several findings merit attention. The specifications with r_1 and r_{td} display first-order autocorrelation in the residuals, while one cannot reject the null hypothesis of no autocorrelation in the specifications with r_d, r_s and r_b.[8 9] In the first two cases referred to other test statistics thus might be

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[7]It would be possible to argue, that while one is estimating a parameter in the representative individual's utility function - and not a parameter of a consumption function - an unstable σ would indicate changing preferences rather than misspecification. As HALL (1985) only reports the estimate of σ (and the corresponding standard error), a more systematic evaluation of these issues might be called for.

[8]Note, however, that the estimated ρ:s are biased towards zero (SAWA (1978)), and that the significance levels are not exact (see e.g. DUFOUR & ROY (1985)).

[9]As pointed out by e.g. DAVIDSON & HENDRY (1981) the criterion of uncorrelated residuals is a rather weak test of model adequacy. At least if the consumer processes information rationally, the error ε must be serially uncorrelated, but indefinite numbers of models which are "data coherent" on this criterion are bound to exist.

TABLE 1

Estimation results with seasonally adjusted real per capita consumption of nondurables and services (OLS)

Expectations	r	σ	k	R^2	D-W	ρ1	ρ2	ρ3	ρ4	P-W	E	J-B	C	C^2	CHOW	period
Perfect foresight	r_l	.045 (.043) (.043)	.006 (.002) (.002)	.012	2.79	-.379	.164	-.136	-.018	.204	18.41**	20.2**	-	**	.90	1961IV - 1984IV
	r_d	-.020 (.038) (.038)	.003 (.003) (.002)	.006	2.01	-.034	-.123	-.112	.182	.275	4.80	1.56	*	**	3.91*	1972III- 1984IV
	r_{td}	.063 (.054) (.044)	.008 (.002) (.003)	.015	2.79	-.396	.159	-.155	-.032	.208	17.80**	18.3**	-	**	.56	1962IV - 1984IV
	r_s	.042 (.030) (.037)	.008 (.002) (.003)	.026	2.39	-.200	.043	-.074	.105	.229	9.69*	2.14	-	**	.89	1966III- 1984IV
	r_b	-.011 (.033) (.034)	.004 (.002) (.001)	.002	2.03	-.045	-.134	-.121	.186	.275	4.90	1.74	*	**	4.14*	1972III- 1984IV
Static	r_l	.071 (.042) (.043)	.006 (.002) (.002)	.030	2.77	-.389	.171	-.165	-.008	.204	18.79**	22.1**	-	**	.37	1961IV - 1984IV
	r_d	.041 (.038) (.038)	.007 (.003) (.002)	.024	2.04	-.048	-.213	-.188	.222	.275	2.40	4.31	-	**	.98	1972III- 1984IV
	r_{td}	.057 (.054) (.042)	.008 (.002) (.003)	.013	2.75	-.378	.162	-.161	-.037	.208	16.46**	19.3**	-	**	.20	1962IV - 1984IV
	r_s	.050 (.029) (.032)	.009 (.002) (.002)	.038	2.38	-.190	.049	-.094	.108	.229	16.13**	3.26	-	**	.67	1966III- 1984IV
	r_b	.035 (.033) (.033)	.005 (.002) (.001)	.023	2.04	-.049	-.205	-.175	.225	.275	2.65	4.43	-	**	1.09	1972III- 1984IV
Auto-regressive	r_l	.098 (.067) (.061)	.006 (.002) (.002)	.023	2.73	-.368	.173	-.162	-.021	.204	16.65**	23.4**	-	**	.35	1961IV - 1984IV
	r_d	.054 (.059) (.060)	.007 (.003) (.003)	.017	2.07	-.061	-.187	-.175	.205	.275	2.55	4.12	-	**	1.00	1972III- 1984IV
	r_{td}	.073 (.056) (.045)	.009 (.002) (.003)	.019	2.76	-.380	.161	-.165	-.038	.208	16.73**	19.5**	-	**	.50	1962IV - 1984IV
	r_s	.062 (.040) (.043)	.009 (.002) (.003)	.034	2.35	-.176	.065	-.085	.091	.229	16.28**	3.01	-	**	.90	1966III- 1984IV
	r_b	.046 (.049) (.049)	.005 (.002) (.001)	.018	2.06	-.060	-.182	-.168	.212	.275	2.85	4.32	-	**	1.14	1972III- 1984IV

Standard errors are shown in parentheses below estimated coefficients. Upper standard errors are ordinary least-squares standard errors, lower are White's (1980) heteroskedasticity-corrected quasi "standard errors". R^2 is the squared multiple correlation coefficient unadjusted for degrees of freedom, and D-W is the Durbin-Watson test statistic for first-order autocorrelation. ρ_i:s (i = 1,..., 4) are sample autocorrelation coefficients of order i, and P-W is two times the Prothero & Wallis (1976) / Ljung Box (1978) version of the Box & Pierce (1970) approximate standard error. E denotes the test for heteroskedasticity of Engle (1982 calculated with four lags, and J-B is the Jarque & Bera (1980) test for normality. C and C^2 denote the cusum and the cusum of squa test, respectively, and - is used if the hypothesis of no structural break could not be rejected at the 5 % level of significance. CHOW is the Chow stability test with the supposed structural break set at 1973IV. * indicates that a null hypothesis could be rejected at the 5 % level of significance and ** that it could be rejected at the 1 % level. All estimates are based on quarterly data, and n is the number of observations.

affected. This is probably the case at least with the test for normality of residuals, which otherwise does not reject the null hypothesis. Furthermore, at least in the case of r_1, r_{td} and r_s signs of heteroskedasticity can be detected. The cusum of squares test can be interpreted as signalling departures from homoskedasticity as well.

Most caution is perhaps needed when interpreting the stability test statistics. In general the cusum test seems unable to detect any systematic movement in σ. It is, however, well known that the test is not very powerful at all instances (see e.g. GARBADE (1977)). The cusum of squares on the other hand always hints at some haphazard movement in σ. Nevertheless, this test is known to give "significant" results more often than an exact test would, and one cannot exclude that mere deviations from independence of residuals or homoskedasticity have been picked up. Taken together, the two tests could be interpreted as suggesting that the possible instability is due to a shift in the residual variances rather than to a shift in σ. The Chow test does not in general reject the stability hypothesis, but the heteroskedasticity adds some uncertainty to these results (see e.g. TOYODA (1974) and SCHMIDT & SICKLES (1977)).[10] One interpretation of the evidence presented in Table 1 could be that specifications involving r_1 and r_{td} simply are less adequate than the other ones. Fairly extensive testing of this proposition by means of the DAVIDSON & MACKINNON (1981) J-test, however, compellingly failed to support such a hypothesis (the results are available from the author upon request). Further examination seems warranted, and this will be conducted in the next section.

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[10] Rather similar conclusions are reached by KOSKELA & VIRÉN (1984a) within HALL's (1978) specification using data from Finland and ten other OECD countries.

In order to utilize the maximum number of degrees of freedom to obtain consistent estimates of σ model (1) was - notwithstanding the possible changes in preferences - reestimated using a consistent instrumental variables estimator (IV).[11] [12] Estimation results are presented in Table 2. The estimated elasticity of intertemporal substitution is uniformly rather small. Minimum, average and maximum values of σ obtained are .0003, .068 and .127, respectively. Estimates differ in a seemingly unsystematic manner roughly as much according to different expectations mechanism as according to different asset yields (although there seems to be a tendency to obtain the smallest σ with static expectations). All IV-estimates of σ differ clearly from the corresponding OLS-estimates, but the order of magnitude of σ is roughly the same.

All elasticities are, however, rather imprecisely estimated. In fact, no estimated σ reaches the two standard error level. This could, of course, result from an improper choice of instruments, but as noted earlier (fn. 11), no clearly "better" instruments than the ones utilized here were found. A perhaps more plausible explanation is that the relative variation in the expected real return is insufficient to allow a precise estimation of σ. In fact it was not altogether impossible to increase the estimation efficiency; when the data were filtered using the Hildreth & Lu method (HL) quite a few precisely estimated elasticities were

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[11] As the realization of r is not known in period t-1, the consumption innovation is correlated with r. Thus any variable in the period t-1 information set would seen eligible for the instrumentation of r. HALL (1985) has, however, pointed out, that when c_{t-1} represents an average, any variable measured in period t-1 can be correlated with ε_t. Accordingly, r_{t-2}, Δc_{t-2} and the realized inflation rate π_{t-2} were used as instruments for r throughout the whole paper (a great number of instruments were evaluated, but neither σ nor the standard error of σ turned out to be very sensitive to the choice of instruments).

[12] The estimates based on autoregressive expectations might still pose a problem (NELSON (1975)). Even IV-estimates will not in general (unless $H_0 : \sigma = 0$) be consistent in these cases (PAGAN (1984)).

TABLE 2

Estimation results with seasonally adjusted real per capita consumption of nondurables and services (IV)

Expectations	r	σ	k	R^2	D-W	period	n
Perfect foresight	r_1	.127 (.075)	.006 (.002)	.012	2.81	1961IV - 1984IV	93
	r_d	.077 (.075)	.009 (.005)	.006	2.12	1972III- 1984IV	50
	r_{td}	.067 (.057)	.009 (.003)	.015	2.79	1962IV - 1984IV	89
	r_s	.053 (.040)	.009 (.003)	.026	2.41	1966III- 1984IV	74
	r_b	.0003 (.0457)	.005 (.002)	.002	2.06	1972III- 1984IV	50
Static	r_1	.079 (.044)	.006 (.002)	.030	2.77	1961IV - 1984IV	93
	r_d	.057 (.039)	.008 (.003)	.024	2.02	1972III- 1984IV	50
	r_{td}	.062 (.054)	.008 (.002)	.013	2.75	1962IV - 1984IV	89
	r_s	.053 (.030)	.009 (.002)	.038	2.38	1966III- 1984IV	74
	r_b	.051 (.033)	.005 (.002)	.023	2.01	1972III- 1984IV	50
Auto-regressive	r_1	.110 (.071)	.006 (.002)	.023	2.73	1961IV - 1984IV	93
	r_d	.080 (.063)	.009 (.004)	.017	2.06	1972III- 1984IV	50
	r_{td}	.077 (.058)	.009 (.002)	.019	2.76	1962IV - 1984IV	89
	r_s	.066 (.041)	.009 (.003)	.034	2.35	1966III- 1984IV	74
	r_b	.063 (.051)	.005 (.002)	.018	2.05	1972III- 1984IV	50

For an explanation of symbols see table 1.

obtained.[13] Since these estimates of the intertemporal elasticity of substitution inevitably are inconsistent, they are, however, assigned to the Appendix (Table A.1). The constant k is irrespectively of the estimation technique generally very precisely estimated.

What do estimates of σ from different periods reveal?

In the previous section some signs of movements or temporary distortions in the intertemporal elasticity of substitution were documented. In order to further investigate this issue using consistent estimates of σ a rolling regression with the instrumental variables estimator was performed. The expected real lending rate assuming perfect foresight was chosen for the analysis. Each segment contained 40 observations, and the first estimated period was 1961IV-1971III. Estimates of σ are plotted against time in Figure 1 (the time index refers to the latest observation used).

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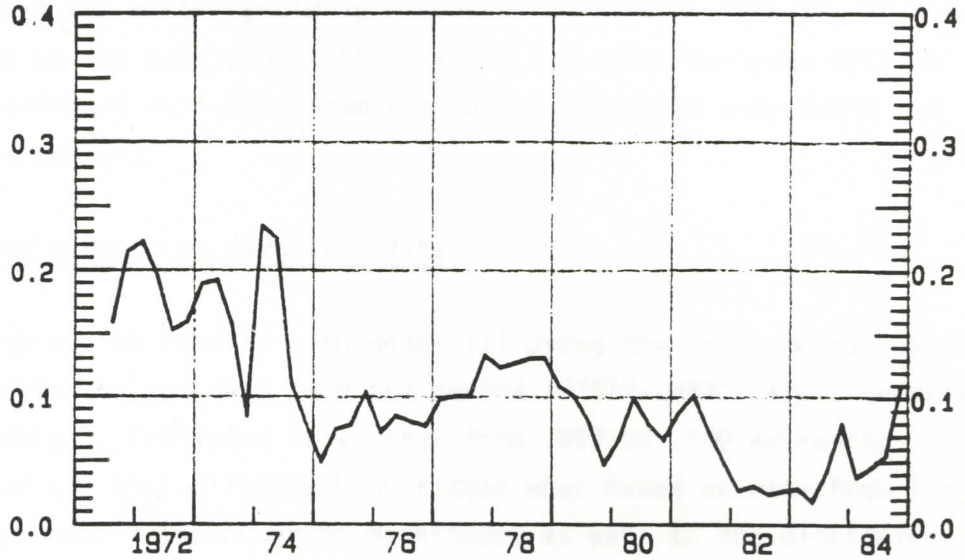
[13] MANKIW & SHAPIRO (1986) have recently demonstrated that the use of highly autocorrelated regressors in models like (1) necessitates the use of slightly higher critical levels of significance. Estimated first-order autoregressive coefficients of the r:s range from .205 to .953, but although these are biased towards zero (SAWA (1978)), this potential bias does not seem to be of crucial importance.

FIGURE 1

Estimate of the intertemporal elasticity of substitution from moving regression



Estimates of σ range from .018 to .235. The graph of σ can be interpreted to suggest a downward trend in σ, or more precisely a more rapid decline in the early periods followed by a slowly declining, or constant, σ from 1975 onwards (using data from e.g. the 80s only yielded σ = .096).[14] The possibility that the observed movements are manifestations of a temporary distortion in the period 1973IV-1974IV cannot, however, be left out of consideration.[15]

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[14] A tentative explanation for this could lie in demographic changes. Even if individual preferences were stable over time, preferences of individuals of different ages differ. As the age structure of the population changes, the age distribution represented by the representative consumer will change.

[15] During the early 70s several exceptional changes occurred in Finland. Among these were the transition to negative (large) real rates of interest, quick and comprehensive demographical changes, and rapid rises in housing costs. If, e.g., the representative consumer's utility function depends on income distribution, estimates of σ might not be invariant to changes in policy rules that induce redistribution of income. Model (1) was also estimated with a dummy variable for this period, but this extension failed to account for any changes.

Since some signs of model inadequacy were detected in the previous section, and because the above tests "should be regarded as yardsticks for the interpretation of data rather than leading to hard and fast decisions" (BROWN et. al. (1975), p. 150), no claim that the matter would be closed will be made. However, in order to be on the safe side - if possible - further analyses will be performed with data from the post-1974 period only (with one exception).

Estimates with post-1974 data

Estimation results with model (1) using the instrumental variables estimator and data from the period 1975IV-1984IV are presented in Table 3. Estimates of σ range from .069 to .130 averaging .092, and are thus slightly higher than when based on data from longer periods.[16] The order of magnitude, as well as the difficulty to obtain precise estimates, are, however, unaltered. The highest estimates are obtained by assuming perfect foresight, and the lowest with autoregressive expectations, but in general outcomes vary only little and in an unsystematic manner. Differences in autoregression among residuals are minor, and autocorrelation does not seem to be alarmingly high.

In order to evaluate the robustness of the above results, some additional sets of estimates will be commented briefly. Estimates so far have been based on seasonally adjusted data (because the data were obtainable in this form only). As utility presumably depends on actual consumption rather than on consumption adjusted by X-11, equation (1) was reestimated with another, seasonally unadjusted, data set. These estimates revealed only marginal

[16] This is in accordance with the predictions of the theoretical model of BECKER & BARRO (1986), who argue that real interest rates and growth rates of consumption per capita would be unrelated in the long run.

TABLE 3

Estimation results with seasonally adjusted real per capita consumption of nondurables and services 1975IV-1984IV (IV)

Expectations	r	σ	k	R^2	D-W
Perfect foresight	r_l	.130 (.102)	.004 (.002)	.015	2.48
	r_d	.126 (.100)	.010 (.005)	.013	2.48
	r_{td}	.113 (.085)	.009 (.004)	.024	2.37
	r_s	.119 (.086)	.011 (.005)	.044	2.55
	r_b	.094 (.080)	.003 (.002)	.026	2.49
Static	r_l	.081 (.058)	.004 (.002)	.029	2.33
	r_d	.080 (.058)	.008 (.003)	.027	2.33
	r_{td}	.082 (.075)	.008 (.004)	.019	2.33
	r_s	.076 (.053)	.009 (.004)	.060	2.38
	r_b	.074 (.052)	.004 (.002)	.040	2.35
Auto-regressive	r_l	.074 (.090)	.004 (.002)	.018	2.34
	r_d	.069 (.089)	.007 (.004)	.015	2.34
	r_{td}	.106 (.082)	.008 (.004)	.031	2.35
	r_s	.079 (.080)	.009 (.005)	.060	2.39
	r_b	.073 (.078)	.004 (.002)	.034	2.36

n = 40. For an explanation of symbols see table 1.

changes in the estimates of σ compared to those reported in Table 3, and are hence reported in the Appendix (Table A.2).[17]

For the sake of comparison with the analyses of HALL (1985), model (1) was also estimated using seasonally adjusted real consumption of nondurables. On an average results did not differ much from the results presented in Table 3 (the mean of the estimated σ:s was .085), but the dispersion of the estimates was somewhat higher (from σ = .043 to σ = .168). For details see Table A.3 in the Appendix.[18]

How reliable is the model?

In the previous sections rather striking robustness in the estimates of σ with respect to different interest rates, expectations mechanisms, consumption measures and data sets was documented. More fundamentally, however, the reliability of the estimates should be regarded as conditional on the validity of the estimated model.[19] Although the model is a generalization of Hall's earlier model, it still is very simple, and several strong assumptions about inter- and intratemporal separability and capital and labour market conditions are carried over. The orthogonality

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[17]This consumption series was constructed by the Central Statistical Office of Finland, and covers the post-1974 period only. Seasonal dummies (Q_1, Q_2 and Q_3) were employed, and the minimum, average and maximum estimates of σ were .059, .099 and .128, respectively. In passing one might note, that the seasonal fluctuation in the rate of change of real per capita consumption is rather strong (see Table A.2 for details).

[18]In fact estimates of σ were roughly unchanged when estimated from seasonally adjusted real consumption of nondurables and services and from per capita consumption of nondurables. These estimates are not reported for space reasons, but they can be obtained from the author upon request.

[19]The approach itself can, of course, also be questioned, as well as the use of aggregate data. For a general discussion see e.g. MANKIW (1986), an for more specific criticism of the two topics see BLINDER (1983), BLINDER & DEATON (1985), MANKIW et. al. (1985) and DIAMOND (1985), HAYASHI (1985), KURZ (1985), respectively.

condition implied by the model, however, allows one to test the validity of these assumptions.

A large number of potentially significant variables could be considered. In order to avoid condescending to data mining, only variables with a theoretical justification will be used. The orthogonality tests draw heavily on the large amount of parallelling tests of Hall's original model, and in order to economize with space only some representative references - rather than the rigorous justifications - for each variable will be given. Each variable will be included in (1) separately, once lagged one period, and once lagged one, two, three and four periods. The estimated coefficient of a one period lagged auxiliary variable X will be denoted β.

Straightforward variables to include are the log-change in consumption Δc_{t-1} (HALL (1978)), the expected real interest rate $r_{1,t-2}$ (HALL (1985)), the nominal interest rate $R_{1,t-1}$ (MANKIW (1982)) and the expected inflation rate π^e_{t-1} (BLINDER & DEATON (1985)). The unexpected inflation rate π^{ue}_{t-1} has its origin in the price confusion effect proposed by DEATON (1977). Both π^e and π^{ue} were obtained from an univariate AR(4) process for the inflation rate. A connection between the growth rate of consumption and the (log of the) real per capita disposable income y_{t-1} or the change in income Δy_{t-1} could arise if consumption and leisure are not separable in the utility function (NELSON (1985)). Constraints in the labour or the capital market would motivate variables like the unemployment rate u_{t-1}, the change in the unemployment rate Δu_{t-1}, and the (log-) change in the real after-tax wage rate Δw_{t-1} (KING (1985a)). Liquidity constraints could also motivate the (log-) changes in a liquidity variable ΔL_{t-1} (measured as M2 divided by disposable income) (DALY & HADJIMATHEOU (1981)). Furthermore, the consumer could respond to the relative price of consumption goods p^r_{t-1} (BLINDER & DEATON (1985)), and to changes in wealth ΔW_{t-1} (proxied by the log-change in a real per capita stock price index) (BILSON (1980)). Finally, a time trend could represent e.g. changes in tastes (BLINDER & DEATON (1985)).

Estimation results with model (1) and additional regressors are reported in Table 4. Maximum time periods are utilized to gain as much power as possible in the orthogonality test. In general, model (1) cannot be said to be at variance with Finnish aggregate time series data from the period 1960-1984. In 13 cases out of 14 the β:s are small and convincingly insignificant. This evidence is somewhat more uniformly in favour of the orthogonality proposition than earlier evidence in the literature, thus suggesting that the relaxation of the constant real rate of interest assumption is a crucial one. Furthermore, unlike other studies with a variable interest rate (MANKIW (1981), MUELLBAUER (1982)) income is not found to influence consumption significantly.[20]

There is, however, one important exception from the otherwise uniform evidence. The one period lagged rate of change of consumption is beyond doubt significant (further lags did not quite pass the 5 % level of significance), and the estimated coefficient is furthermore relatively large. Unfortunately, the evaluation of lags of Δc in (1) is the most general test conceivable for model adequacy (KING (1985a), p. 255), and a rejection of the orthogonality condition gives little insight into the cause of failure.

Broadly speaking the significant lag could stem from separability, rationing and (or) data issues. Exhaustive answers are hardly to be expected, but a short discussion seems warranted. The assumption of intertemporal separability is crucial to the empirical tests of models of this type, although the validity of this assumption is

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[20]HAYASHI (1985) has recently argued that a significant income variable could be attributable to a correlation between consumption and income measurement errors or consumption and leisure taste shocks. The latter correlation can occur despite the maintained hypothesis that consumption and leisure are separable in the utility function. "Significant" findings in the literature might furthermore partly be attributable to the factors pointed out by MANKIW & SHAPIRO (1986), and to the use of data averages when temporal aggregation problems are present (TIAO & WEI (1976)).

TABLE 4

Estimation results with additional regressors (IV)

X	σ	β	k	R^2	D-W	F	period	n
Δc_{t-1}	.187	-.433	.009	.148	2.02	4.58**	1960IV - 1984IV	97
	(.070)	(.101)	(.002)					
$r_{l,t-2}$	.254	-.066	.006	.010	2.78	1.18	1960IV - 1984IV	97
	(.446)	(.263)	(.002)					
$R_{l,t-1}$	.113	-.001	.012	.016	2.82	.57	1961IV - 1984IV	93
	(.089)	(.002)	(.017)					
π^e_{t-1}	.058	-.046	.010	.023	2.80	1.37	1961IV - 1984IV	93
	(.400)	(.300)	(.026)					
π^{ue}_{t-1}	.128	.113	.006	.024	2.73	1.38	1961IV - 1984IV	93
	(.072)	(.086)	(.002)					
y_{t-1}	.134	-.0004	.009	.022	2.83	.69	1960IV - 1983IV	93
	(.091)	(.0020)	(.012)					
Δy_{t-1}	.141	.006	.006	.023	2.84	.53	1960IV - 1983IV	93
	(.074)	(.009)	(.002)					
u_{t-1}	.138	-.001	.008	.024	2.83	.37	1960IV - 1984IV	97
	(.072)	(.001)	(.004)					
Δu_{t-1}	.145	-.002	.006	.028	2.84	.36	1960IV - 1984IV	97
	(.072)	(.002)	(.002)					
Δw_{t-1}	.146	-.064	.007	.021	2.82	.16	1960IV - 1980IV	81
	(.082)	(.139)	(.002)					
ΔL_{t-1}	.143	-.086	.006	.029	2.86	.70	1960IV - 1983IV	93
	(.074)	(.069)	(.002)					
p^r_{t-1}	.155	.033	-.031	.071	2.80	.85	1960IV - 1984IV	97
	(.121)	(.163)	(.180)					
ΔW_{t-1}	.138	.001	.006	.014	2.82	1.17	1962III- 1984IV	90
	(.088)	(.031)	(.002)					
T_{t-1}	.128	-.003	.008	.024	2.83	-	1960IV - 1984IV	97
	(.078)	(.007)	(.004)					

All estimates are based on seasonally adjusted real per capita consumption of nondurables and services and the expected real lending rate assuming perfect foresight. F denotes a F-test for the hypothesis that the estimated coefficients of the auxiliary variable lagged one, two, three and four periods all equal zero. ** denotes that the hypothesis can be rejected at the 1 % level of significance. For an explanation of other symbols see table 1.

far from clear. The further assumption that preferences for consumption and leisure within a given period are separable is also questionable, although the clear insignificance of the income variables lessens doubts for this matter.[21]

The implicitly assumed absence of constraints in the capital market and in the labour market may constitute a more serious deficiency. Recent empirical work suggests that a fraction (say 20 per cent) of the consumers are liquidity constrained in the sense of credit rationing or differential interest rates (see HAYASHI (1985) for a comprehensive survey). It seems highly likely that liquidity constraints have been predominant in Finland for most of the time considered, and some fraction of the consumers may well have been affected in the 80s as well.[22] [23] Labour supply rationing may also affect consumption via various channels, and some empirical evidence supporting this view has been presented (see e.g. KOSKELA & VIRÉN (1984c), (1985)).

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[21] Neither evidence of intertemporal separability (e.g. HOTZ et.al. (1986) nor evidence of intratemporal separability (e.g. ABBOTT & ASHENFELTER (1976)) is compelling. For a more exhaustive discussion see DEATON & MUELLBAUER (1980), chapter 5.

[22] E.g. the system of housing financing might be regarded as causing liquidity constraints nowadays. Roughly 20 per cent of the households in Finland thus potentially might be liable to constraints (KOSONEN (1986)). A possibility of credit rationing also exists if certain households are not able to borrow at all (STIGLITZ & WEISS (1981)). Expected but uncertain future (temporary) borrowing constraints may furthermore influence current consumption (KOSKELA & VIRÉN (1984b)).

[23] Some empirical evidence for Finland was obtained by applying the testing procedure developed by FLAVIN (1981). Of the consumers 6.8 % - 10.5 % (depending on the consumption measure employed) were found to face binding constraints. The t-values of these estimates, however, never exceeded .5, and the estimates were not constant over time. As the t-values furthermore may be biased upwards (see the discussion in e.g. HAYASHI (1985), KING (1985a), NELSON (1985) and WALSH (1985)), and the estimates are conditional on the hypothesis of constant real interest rates, no clearly convincing evidence seems to emerge.

A more stringent test of the occurrence of constraints than the orthogonality test may nevertheless be performed. Two subsamples were created by noting wether an observation belonged to a period in which the unemployment rate was above its trend value or under its trend value. Thus one could evaluate the significance of labour market conditions by testing for the equality of σ across subsamples. The Chow statistic rejects homogeneity of the two samples at the 5 % level of significance, but not at the 1 % level ($F(2, 93) = 3.72$). This adds some uncertainty to the validity of the assumption of absence of constraints in the labour market. A similar test for capital market constraints was performed on the basis of periods of excess demand for versus excess supply of bank loans as identified by TARKKA (1986). The homogeneity hypothesis could not be rejected in this case ($F(2, 93) = .73$). However, as there are some intuitively acceptable reasons to expect that constraints in labour and capital markets are connected it would be misleading to claim too much for these results. KING (1985b) finds empirical evidence of a positive correlation by using data from the U.K. and the U.S.A.

Finally, the significant lag could be caused by factors that stem from deficiencies or abnormalities in the data rather than from implications from economic theory.[24] The important difficulties due to the use of aggregate data, and time aggregation bias have already been mentioned. A closer examination of the significance of Δc_{t-1} revealed that the coefficient was significant only when observations from 1964 or earlier years were included. The size, as well as the significance, of the estimated coefficient dropped markedly when 1965I, or a later period, was used as first observation. More precisely, the significance was dependent on three observations (1961III, 1963IV and 1965I) which contained the

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[24]The basic permanent income - life cycle hypothesis predicts that Δc should be serially uncorrelated. The sample autocorrelation at lag one, however, is -.358, which is highly significantly different from zero. Thus it is clear that Δc can be predicted from past values. No autocorrelation is, however, found when data from 1965 onwards are used.

three fastest consumption growths in the whole data. In 1963IV this was furthermore paired with the highest inflation peak in the sample.[25]

As results in previous sections suggested that the most relevant period to consider in order to gain knowledge about the size of σ and about the validity of the model today would be the post-1974 period, it might be concluded that the findings that entirely rely on pre-1965 data should not be given too much weight. Thus one is not able to reject the orthogonality condition by using this set of aggregate time series data.

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[25] In 1965I the calculation of the population figure was revised. Since estimation results with data undeflated by population were practically identical to those reported in Table 4 the revision, however, seems to be of minor importance.

V CONCLUDING REMARKS

The identification of the structural parameters is, i.a., a necessary prelude to the construction of macro models that would allow us to analyze effects of policy interventions and to study business cycles. One such preference parameter is the intertemporal elasticity of substitution, which measures the magnitude of the response of consumption to a change in real interest rate expectations. The empirical evidence from aggregate Finnish time series data suggests that this substitution effect is rather small - the order of magnitude is .1. This estimate is furthermore rather robust to a broad menu of alternative operationalizations of the estimated model, which itself survives thorough attempts to reject it. In the light of this evidence there does not seem to be much scope for a policy aiming at making consumers defer consumption by affecting their real interest rate expectations.

The estimated intertemporal elasticity of substitution is of roughly the same size as earlier consistent estimates indicate (most of the previous studies, however, suffer from various biases; see HALL (1985) for details and references).[26] Since some evidence suggesting that the elasticity has not been constant through time is found, the small differences between studies might be caused by differences in the considered time periods. HAYASHI (1985) has recently pointed out, that differences also may arise from differences in the nature of the loan market if liquidity constraints occur. A topic for further research clearly would be to allow for heterogeneity among consumers with respect to e.g. possible occurrence of liquidity constraints and with respect to age. This, however, would require access to comprehensive micro data.

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[26] Minimum, average and maximum values found in this study (using post-1974 data) are .043, .092 and .168, respectively. HALL (1985) reports consistently estimated values (based on aggregate time series data from the U.S.) of -.455, .032 and .346, and SKINNER (1985) documents (using cross-sectional data from 1972/73 from the U.S.) values of -.392, .382 and 1.448.

APPENDIX 1

TABLE A.1

Estimation results with seasonally adjusted real per capita consumption of nondurables and services (HL)

Expectations	r	σ	k	ρ	R^2	D-W	period	n
Perfect foresight	r_l	.074 (.031)	.006 (.001)	-.400	.059	2.02	1961IV - 1984IV	93
	r_d	-.016 (.037)	.004 (.003)	-.040	.004	1.95	1972III- 1984IV	50
	r_{td}	.071 (.036)	.009 (.002)	-.390	.042	2.00	1962IV - 1984IV	89
	r_s	.050 (.026)	.009 (.002)	-.200	.051	2.00	1966III- 1984IV	74
	r_b	-.007 (.032)	.004 (.002)	-.050	.001	1.96	1972III- 1984IV	50
Static	r_l	.075 (.031)	.006 (.001)	-.380	.060	2.00	1961IV - 1984IV	93
	r_d	.039 (.037)	.007 (.003)	-.060	.023	1.95	1972III- 1984IV	50
	r_{td}	.052 (.037)	.008 (.002)	-.370	.023	1.99	1962IV - 1984IV	89
	r_s	.053 (.025)	.009 (.002)	-.190	.057	1.99	1966III- 1984IV	74
	r_b	.034 (.032)	.005 (.002)	-.060	.023	1.95	1972III- 1984IV	50
Auto-regressive	r_l	.088 (.048)	.006 (.001)	-.360	.036	1.99	1961IV - 1984IV	93
	r_d	.053 (.056)	.007 (.003)	-.070	.018	1.96	1972III- 1984IV	50
	r_{td}	.065 (.038)	.008 (.002)	-.380	.032	1.98	1962IV - 1984IV	89
	r_s	.062 (.034)	.009 (.002)	-.170	.045	1.99	1966III- 1984IV	74
	r_b	.044 (.046)	.005 (.002)	-.060	.018	1.97	1972III- 1984IV	50

The first-order autoregressive filter is denoted ρ. For an explanation of other symbols see table 1.

TABLE A.2

Estimation results with seasonally unadjusted real per capita consumption of nondurables and services 1975IV-1984IV (IV)

Expectations	r	σ	k	Q_1	Q_2	Q_3	R^2	D-W
Perfect foresight	r_l	.118	-.012	.021	.046	-.007	.824	2.29
		(.097)	(.003)	(.005)	(.005)	(.005)		
	r_d	.123	-.006	.022	.046	-.007	.823	2.28
		(.096)	(.005)	(.005)	(.005)	(.005)		
	r_{td}	.128	-.007	.022	.046	-.006	.833	2.27
		(.081)	(.004)	(.004)	(.004)	(.004)		
	r_s	.108	-.005	.021	.046	-.007	.835	2.36
		(.081)	(.006)	(.004)	(.004)	(.004)		
	r_b	.083	-.012	.022	.046	-.007	.834	2.29
		(.076)	(.003)	(.004)	(.004)	(.004)		
Static	r_l	.080	-.011	.022	.045	-.006	.832	2.19
		(.055)	(.003)	(.004)	(.004)	(.004)		
	r_d	.080	-.008	.022	.045	-.006	.831	2.19
		(.055)	(.004)	(.004)	(.004)	(.004)		
	r_{td}	.059	-.009	.022	.045	-.006	.834	2.18
		(.075)	(.004)	(.004)	(.004)	(.004)		
	r_s	.071	-.007	.021	.045	-.007	.838	2.24
		(.050)	(.004)	(.004)	(.004)	(.004)		
	r_b	.071	-.012	.022	.045	-.007	.836	2.21
		(.049)	(.003)	(.004)	(.004)	(.004)		
Auto-regressive	r_l	.118	-.011	.022	.045	-.006	.831	2.15
		(.087)	(.003)	(.004)	(.004)	(.004)		
	r_d	.114	-.006	.022	.045	-.006	.831	2.14
		(.086)	(.005)	(.004)	(.004)	(.004)		
	r_{td}	.124	-.007	.022	.045	-.006	.835	2.18
		(.078)	(.004)	(.004)	(.004)	(.004)		
	r_s	.105	-.005	.021	.045	-.007	.840	2.21
		(.076)	(.006)	(.004)	(.004)	(.004)		
	r_b	.101	-.012	.022	.045	-.007	.837	2.18
		(.074)	(.003)	(.004)	(.004)	(.004)		

n = 40. For an explanation of symbols see table 1.

TABLE A.3

Estimation results with seasonally adjusted real consumption of nondurables 1975IV-1984IV (IV)

Expectations	r	σ	k	R^2	D-W
Perfect foresight	r_l	.168 (.126)	.004 (.002)	.015	2.52
	r_d	.164 (.124)	.011 (.006)	.014	2.51
	r_{td}	.081 (.082)	.008 (.004)	.023	2.34
	r_s	.136 (.102)	.012 (.006)	.030	2.57
	r_b	.114 (.095)	.004 (.002)	.023	2.53
Static	r_l	.071 (.047)	.005 (.002)	.045	2.27
	r_d	.069 (.046)	.008 (.003)	.043	2.27
	r_{td}	.061 (.071)	.007 (.004)	.012	2.32
	r_s	.069 (.044)	.009 (.004)	.064	2.29
	r_b	.066 (.044)	.004 (.002)	.053	2.28
Auto-regressive	r_l	.048 (.103)	.004 (.002)	.012	2.32
	r_d	.043 (.098)	.006 (.005)	.010	2.32
	r_{td}	.076 (.078)	.008 (.004)	.015	2.34
	r_s	.062 (.092)	.008 (.006)	.040	2.35
	r_b	.052 (.088)	.004 (.002)	.024	2.34

n = 40. For an explanation of symbols see table 1.

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