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THE CONSUMPTION FUNCTION IN NORWAY
BREAKDOWN AND RECONSTRUCTION

by

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Breakdown and reconstruction*)

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Abstract:

In the mid eighties econometric forecasts and ex post simulations of private consumption in Norway began to show clear signs of "structural breakdown". This evidence lends itself to two interpretations, distinct in their implications for econometric modelling of aggregate consumption. On the one hand consumer behaviour may have been fundamentally changed by the recent deregulation of the financial and housing markets in Norway. Another interpretation stresses the possibility of inherent misspecification of empirical consumption functions. In the latter case one would aim at reconstructing an empirical consumption function, using data which does not include the breakdown period, but which is capable of accounting for the development of private consumption in that period. In this paper we challenge the interpretation that forecast failures provide evidence of "structural shift" by reconstructing the consumption function, using data for the whole sample period.

Moreover, the paper puts emphasis on the temporal properties of the data according to tests proposed by Hylleberg, Engle, Granger og Yeo (1988). The empirical evidence implies that consumption and income are not cointegrated at the zero frequency. On the other hand, cointegration tests - including Johansen's Full information procedure - indicate that consumption, income and wealth are cointegrated.

The paper shows that an error correction model in consumption, income and wealth is invariant of the changes in the credit and in the housing markets. This shows that the predictive failure of existing consumption functions does not constitute evidence against a conditional modelling approach to private households consumption.

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1. Introduction

In a keynesian model of the economy, interest quite naturally focus on the relationship between household expenditure and private disposable income. This explains why the development of macro-econometric models in the early sixties triggered off much empirical research on the macro consumption function. In Norway for example, the consumption function has been a key element in econometric models for the last twenty years. The explanatory power and tracking performance of empirical consumption functions have been judged as satisfactorily compared with more troublesome equations, such as e.g. investment and wage equations. However, in the mid-eighties, forecasts and ex post simulations of private consumption began to drift way off target. This development naturally makes one ask whether the breakdown is evidence of fundamental changes in consumer behaviour or if it is merely disclosing the misspecification of empirical consumption functions. In the latter case one would aim at reconstructing an empirical consumption function, using data which does not include the breakdown period, but which is capable of accounting for the development of private consumption in that period. In this paper we challenge the "structural shift" interpretation by reconstructing the consumption function, using data for the whole sample period.

The rest of this paper is organized as follows. In section 2 we produce evidence for the fact that the existing empirical consumption models have performed badly after 1984. Although the breakdown has occurred in both quarterly and annual models, our main interest in this paper lies with the quarterly models. We therefore investigate the performance of the consumption functions in the two quarterly macroeconometric models in Norway, KVARTS (Jensen and Reymert (1984)) and RIKMOD (cf. e.g. Jansen (1984)).

It has been a long standing claim among time series analysts that investigation of the temporal properties of the individual data series is a natural first step in econometric work (cf. e.g. Granger and Newbold (1974)). The recent literature on cointegration and dynamic representation of cointegrated series has given new relevance to this argument. In particular Granger's representation theorem (Engle and Granger (1987)) implies that the error-correction formulation introduced by Davidson, Hendry, Srba and Yeo (DHSY) (1978) can be interpreted as a dynamic representation of the relationship between consumption and income,i.e. when the long run elasticity of consumption is unity and income is weakly exogenous (Engle et.al. (1983)) for the parameters of interest in the consumption function. Hylleberg, Engle, Granger, and Yoo (1988) (HEGY) extends cointegration theory to seasonally integrated series.

These developments, even if by no means conclusive as to which dynamic consumption model may constitute an adequate characterization of the data, seems to offer valuable insights for the applied econometricians' search for adequate empirical formulations. At least, investigation of integration and cointegration may help eliminate models, i.e. if they involve variables with inconsistent temporal properties.

In section 3 we estimate the properties of the quarterly consumption and income data. Both consumption and income are clearly seasonal, we therefore use the integration tests in HEGY.

Section 4 sums up the implications from our investigation of the temporal data properties for dynamic modelling, namely that models based only on income and consumption are not statistically well founded. A wider information set is necessary. In section 5 we discuss alternative empirical long run equilibrium relationships, and section 6 shows that we are able to establish empirically stable dynamic consumption functions for the whole sample period.

2. The evidence for a breakdown

In this section we briefly discuss the evidence for breakdown in Norwegian consumption functions. The discussion is based on re-estimation of the consumption functions in the two quarterly macroeconometric models of the Norwegian economy KVARTS (Jensen and Reymert (1984)) and RIKMOD (Jansen (1984)).

The KVARTS consumption function was originally estimated in Biørn and Jensen (1983) who fitted an extended linear expenditure system (ELES) model for Norwegian quarterly data. The empirical formulation of the macro consumption function is:

$$\text{(2-1)} \qquad \hat{C}_t = \{\hat{\beta}_0 + \sum_{i=0}^{7} \hat{\beta}_{1-i} Y_{t-i} + \sum_{i=0}^{3} \hat{\beta}_{2-i}\, \Delta CR_{t-i}\} \underline{\hat{\beta}}_3\, \underline{Q}_t$$

C is total private consumption in fixed prices and Y is households' disposable income. ΔCR is a measure of increases in credits to households, $\underline{Q}$ is a composite term containing seasonals and a dummy relating to the introduction of VAT in 1970.[1)]

The RIKMOD function is almost identical to the formulation of the permanent income hypothesis in Evans (1969):

$$\text{(2-2)} \qquad (\hat{\frac{C}{Y}})_t = \hat{\alpha}_0 + \hat{\alpha}_1 \left[\sum_{i=0}^{3} (1 - 0.25\ i) \frac{\Delta Y_{t-i}}{Y_{t-i-1}} \right]$$
$$+ \hat{\alpha}_2\ 0.25 \sum_{i=1}^{4} \frac{C}{Y}_{t-i} + \underline{\hat{\alpha}}_3\, \underline{Q}_t$$

Table 2.1 shows the evidence for breakdown in these two consumption functions. F_{CH} is the post sample Chow-test and $\chi^2(20)$ is the Chi-square test of forecast accracy (Chow (1960), Hendry (1979)). Both tests are clearly significant. A caveat applies here, because the appendix shows that there is some doubt whether we have actually succeeded in replicating the KVARTS model for the sample period 1968 4 to 1983 4.

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1) The modelling group in the Central Bureau of Statistics kindly provided us with the income and credit data used in the re-estimation of the consumption function in KVARTS.

Table 2.1 Stability tests 1984 1 - 1988 4

	$F_{CH}(T_2,T_1-k)$	Critical values 2)	$X^2(T_2)$	Critical values 2)
RIKMOD 84 1-88 4	$F_{CH}(20,54) = 4.51$	1.77	$X^2(20) = 124$	31.41
KVARTS 1) 84 1-88 4	$F_{CH}(20,52) = 2.59$	1.78	$X^2(20) = 58.54$	31.41

$F_{CH}(T_2,T_1-k)$ = Chow test of parameter stability for T_2 periods out of sample. T_1 is the number of observations and k the number of variables. Distributed as $F(T_2, T_1 - k)$ on the null. Chow (1960).

$X^2(T_2)$ = Post-sample goodness of fit test based on the one step ahead prediction residuals. Distributed approximately as $X^2(T_2)$ on the null (cf. Hendry (1979)).

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1) See appendix for details.

2) Critical values at the 5% level.

While the evidence for breakdown seems to be clear cut, there are contending interpretations of the prediction errors. Firstly, forecasts errors can be evidence of a genuine structural break, i.e. a fundamental change in consumer behaviour. The fact that Norwegian credit and housing markets were deregulated in the mid-eighties is the main motivation for this view. On the other hand, although the forecast errors are evidence of parameter instability in the empirical consumption functions, they do not necessarily imply a structural break, in say the relationship between consumption and income. For example, while a genuine structural break is sufficient to induce predictive failure it is not necessary in the following sense: If all the true behavioural equations relating to e.g. consumption, income and credit, remained unaltered but the behaviour of some exogenous variables changed, then all misspecified econometric approximations to the structural equations could manifest shifts (i.e. structural breaks (cf. Hendry (1979)).

Hence, the predictive failure in empirical consumption functions may disclose that the existing conditional models (2-1) and (2-2) are not structurally invariant to the changes in the marginal distributions of income and credit brought about by financial deregulation, Engle et.al. (1983). In other words, both income and credit may not be super exogenous variables in models (2-1) and (2-2). This fact does not imply that alternative conditional models may not be constructed which are structurally invariant to the events in the breakdown period.

One important distinction between the two interpretations, is that misspecification provides a parsimonious account of the observed predictive failure. This is true in the sense that the "genuine structural break" interpretation leads to a thorough reinvestigation and reformulation of the system, while the misspecification interpretation leads to the more limited task of reconstructing the conditional consumption function, assuming the underlying system to be stable. To us it seems worthwile to try out the parsimonious explanation of predictive failure before undertaking a major theoretical and empirical reformulation of consumption behaviour.

The structural invariance interpretation also indicates a simple criterion for successful reconstruction: A reformulated empirical consumption function should perform at least as well as the existing models on pre-1984 data, and significantly better in the "breakdown period". The rest of this paper discusses the evidence at the present stage of our research. We start by taking a closer look at the main properties of the data.

3. Consumption and income data. Temporal properties

Figure 3.1 shows private consumption (c), consumption of non-durables (c_n) and the disposable income of households (y). The data are measured in log scale.[2] All three series show persistent growth, and have clear seasonal patterns: Consumption peaks in the fourth quarter and income in the third quarter of the year.

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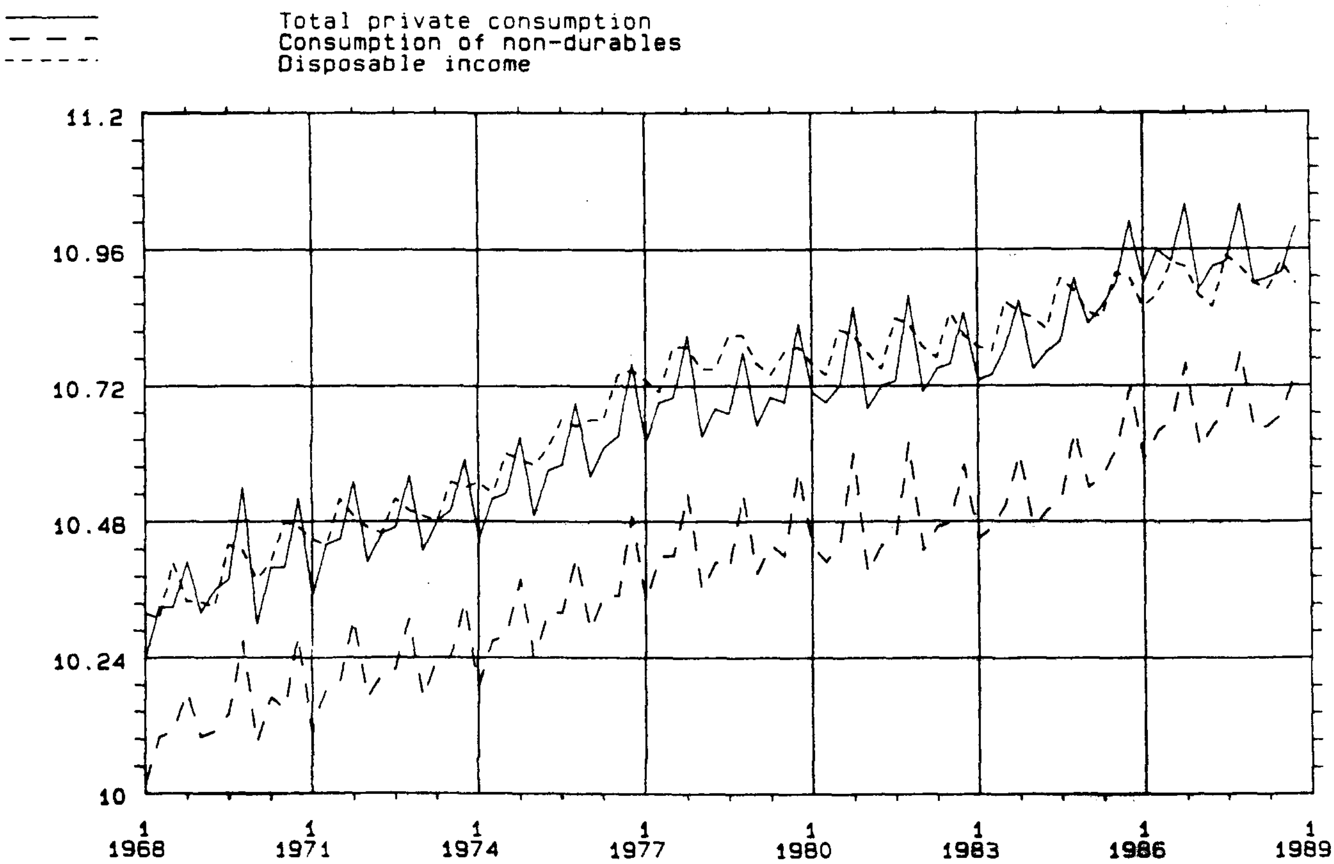
2) The basic consumption and income data are in millions kroner, in fixed 1983 prices.

In Norway, previous studies using quarterly data have aimed at modelling total consumption. Both economic theory and also the empirical literature emphasises the relationship between non-durables consumption and income. In principle therefore the breakdown may be explained by aggregation bias. Inspection of the data does not give much support to this explanation of the predictive failure. The growth rates of total consumption behave in much the same way as the growth in the consumption of durables. It is true that the growth in non-durables consumption falls somewhat short of growth in total consumption in 1985, but it is not likely that the gross predictive failures may be attributed to this alone. In the following we concentrate on total consumption.

Figure 3.1
Log of total private consumption, non-durables and disposable income



Next, we investigate the temporal properties of the consumption and income data. The recent literature shows that care must be taken when applying conventional tests such as the DW (Sargan and Bhargava (1983)) and the Dickey-Fuller test (cf. e.g. Dickey and Fuller (1981), and Fuller (1976)) to quarterly data: If there is seasonality in the data, these tests may be inappropriate. Typically, the tests do not allow for seasonality in the system under the null of non-stationarity. If the seasonality is deterministic, this problem may be tackled by introducing seasonal dummies in e.g. the augmented Dickey-Fuller test. There is still a problem finding the appropriate distribution for this modified Dickey-Fuller test, but intuitively the usual critical values may be used if one allows for somewhat "fatter tails".

Another aspect of seasonality is the possibility of unit roots at the seasonal frequencies instead of, or in addition to, unit roots at the zero frequency. In the case of seasonal unit roots, we say that the data are seasonally integrated, reservering the term integrated series for variables which are integrated at the zero or long run frequency only. Since the conventional tests do not consider seasonal integration, they may well be inappropriate for quarterly data.

A recent paper by Hylleberg, Engle, Granger and Yoo (HEGY) (1988) develops integration test which introduce explicit treatment of the seasonality aspects of quarterly data. HEGY takes into account that seasonality may be due to stationary stochastic processes and deterministic factors, as well as unit roots at the seasonal frequencies.

In table 3.1 we apply the test procedure proposed by HEGY to the consumption and income data, as well as to the consumption-income ratio. Let us assume that a time series X_t has the following generating equation:

(3-1) $\phi(L)X_t = \mu_t + \epsilon_t$

where μ_t consists of deterministic terms and ϵ_t is white noise. HEGY shows that if $\phi(B)$ is rational, this polynominal may be written in the following way:

$$(3\text{-}2) \qquad \begin{aligned}\phi(L) &= \Pi_1[L(1 + L + L^2 + L^3)] \\ &+ \Pi_2[- L(1- L + L^2 - L^3)] \\ &+ (\Pi_4 + \Pi_3 L)(-L(1 - L^2)) \\ &+ \phi^*(L)\ (1 - L^4)\end{aligned}$$

where $\phi^*(L)$ is a real residual polynominal. If $\phi(1) = 0$ there is a long run unit root. From (3-2) we see that $\Pi_1 = 0$ in this case. Similarely, $\Pi_2 = 0$ corresponds to a root at minus one or the 1/2 frequency (biannual cycle). A root at the seasonal frequency (annual cycle) corresponds to $\phi(\pm\ i)=0$, which implies $\Pi_3 = \Pi_4 = 0$.

$$(3\text{-}3) \qquad \begin{aligned}Y_{4t} &= \Pi_1 Y_{1t-1} + \Pi_2 Y_{2t-1} + \Pi_3 Y_{3t-2} + \Pi_4 Y_{3t-1} \\ &- \tilde{\phi}(L)\ Y_{4t-1} + \mu_t + \epsilon_t\end{aligned}$$

with $\phi^*(L) = 1 - \tilde{\phi}(L)$ and:

$$(3\text{-}4) \quad \begin{aligned}&a)\ Y_{1t} = (1 + L + L^2 + L^3)X_t \\ &b)\ Y_{2t} = -\ (1 - L + L^2 - L^3)X_t \\ &c)\ Y_{3t} = -\ (1 - L^2)X_t \\ &d)\ Y_{4t} = (1 - L^4)X_t\end{aligned}$$

$Y_{it}(i=1,2,3,4)$ are filtered in order to isolate the different forms of non-stationarity. Y_{1t} is a moving average of X_t and removes unit roots on the seasonal frequencies and leaves us with a possible unit root on the zero frequency. Y_{2t} removes the zero and the annual frequency, while Y_{3t} removes the zero and the biannual frequencies. Y_{4t} removes all unit roots. In (3-1) the $\hat{\Pi}_1$ will for example tend to zero if we have a unit root on the zero frequency. This follows from the fact that the left hand side variable Y_{4t} is stationary. Since Y_{1t}, Y_{2t}, Y_{3t} and Y_{4t} contains possible unit roots at different frequencies, they will be orthogonal, so we do not have to specify a priori which unit roots that we assume are present in the data.

The tests are based on estimation of (3-3) with different assumptions of the stochastic and deterministic augmentation, i.e. the $\tilde{\phi}(L)Y_{4t-1}$

+ μ_t term. For a given augmentation, we first test for $\Pi_4 = 0$ by consulting the relevant "t" distribution (cf. HEGY op.cit.) If $\Pi_4 = 0$ is sustained, we compute "t-values" for Π_1, Π_2 and Π_3, conditional on $\Pi_4 = 0$. Each "t" is a (one sided) test of the null of roots at the zero, 1/2 and 1/4 (3/4) frequency respectively. Alternatively, we may test for roots at the seasonal frequencies by computing an "F-test" of the joint hypothesis $\Pi_3 = \Pi_4 = 0$.

Looking at table 3.1, we find that for all three series the "F-test" reject integration at the seasonal frequencies if an intercept and seasonal dummies are included. The significant F is clearly due to the Y_{3t-2} term, since the "t_4" of $\Pi_4 = 0$ is not rejected.

Looking at the "t"-tests for $\Pi_3 = 0$ (j=1,2,3) conditional on $\Pi_4 = 0$, we find that integration at the annual frequency is rejected for the cases of no augmentation and an intercept term. Again this is true for all three series. At the biannual frequency all the three series reject integration for the I, SD case, while consumption is $I_{½}(1)$ for the I, SD, TR case. On the other hand, income and the consumption income ratio turns out to be $I_{½}(0)$ for the I, SD, TR case. This is against intuition since theoretically the sum of $I_{1/2}(0)$ and $I_{1/2}(1)$ is itself $I_{1/2}(1)$. Turning finally to the zero (long run) frequency, both consumption and income are integrated. Furthermore they are not cointegrated with cointegration parameter equal to one, since c-y is $I_0(1)$.

Table 3.2 gives the results of the HEGY-tests for the period 1969 1 - 1983 4. This is the period before the financial deregulations and the breakdown of the consumption functions. Exclusion of this volatile period does not change the result that consumption and income are not cointegrated with cointegration para-meter equal to one at the zero frequency.

Table 3.1 HEGY tests for unit roots in the log of private consumption (c), households' disposable income (y) and the difference (c-y). 1969 1 - 1988 4

"F" and "t_4" : $Y_4 = \Pi_1 Y_{1-1} + \Pi_2 Y_{2-1} + \Pi_3 Y_{3-2} + \Pi_4 Y_{3-1}$ + augmentation + residual

"t_1", "t_2", "t_3": $Y_4 = \Pi_1 Y_{1-1} + \Pi_2 Y_{2-1} + \Pi_3 Y_{3-2}$ + augmentation + residual

Variable	Augmentation[1)]	"t_1" Π_1(zero frequency)	"t_2" Π_2(biannual)	"t_3" Π_3(annual)	"F" $\Pi_3 \cup \Pi_4$	"t_4" Π_4
c	-	3.13	-0.95	-0.70	0.26	-0.21
	I	-1.42	-0.92	-0.65	0.22	-0.18
	I, SD	-1.52	-2.57*	-6.49*	20.73*	-0.06
	I, SD, TR	-2.28	-2.80	-6.84*	23.04*	0.01
y	-	2.15	-1.82	-1.44	1.38	-0.72
	I	-2.45	-1.81	-1.44	1.27	-0.70
	I, SD	-2.57	-5.42*	-3.03*	5.87*	-1.54
	I, SD, TR	-1.24	-5.43*	-2.90*	5.46	-1.54
c-y	-	-0.95	-0.73	-1.15	0.65	0.10
	I	-0.93	-0.73	-1.15	0.64	0.11
	I, SD	-2.04	-3.01*	-4.86*	13.07*	-1.47
	I, SD, TR	-2.34	-3.03*	-4.85*	12.93*	-1.41

I: intercept, SD: seasonal dummies, TR: trend

An * means significant at the 5 percent level (cf. HEGY (1988)).

1) In addition to deterministic terms, the augmentation contains significant lags of Y_4

Table 3.2 HEGY tests for "unit roots", 1969 1 - 1983 4

"F" and "t_4" : $Y_4 = \Pi_1 Y_{1-1} + \Pi_2 Y_{2-1} + \Pi_3 Y_{3-2} + \Pi_4 Y_{3-1}$ + augmentation + residual

"t_1", "t_2", "t_3": $Y_4 = \Pi_1 Y_{1-1} + \Pi_2 Y_{2-1} + \Pi_3 Y_{3-2}$ + augmentation + residual

Variable	Augmentation[1)]	"t_1"	"t_2"	"t_3"	"F"	"t_4"
		Π_1(zero frequency)	Π_2(biannual)	Π_3(annual)	$\Pi_3 \cup \Pi_4$	Π_4
c	-	3.73	-0.91	-0.60	0.18	0.04
	I	-2.04	-0.85	-0.48	0.11	0.07
	I, SD	-1.85	-2.63*	-5.29*	14.13*	0.71
	I, SD, TR	-0.89	-2.66	-5.26*	13.91*	0.68
y	-	2.30	-1.97*	-0.87	-0.39	-0.39
	I	-1.84	-2.01*	-0.76	0.36	-0.40
	I, SD	-1.91	-5.38*	-3.20*	6.41*	-1.53
	I, SD, TR	-2.08	-5.61*	-3.01	5.56	-1.38
c-y	-	-0.01	-1.02	-0.84	0.48	0.50
	I	-1.24	-1.04	-0.86	0.52	0.55
	I, SD	-2.44	-2.71*	-5.42*	14.39*	-0.10
	I, SD, TR	-3.22	-2.67	-5.67*	15.80*	-0.23

* means significant at the 5 percent level.

I: Intercept, SD: Seasonal dummies, TR: Trend

1) In addition to deterministic terms, the augmentation contains significant lags of Y_4

4. Implications for dynamic modelling

During the last 10 years or so, the work of Davidson, Hendry, Srba and Yeo (DHSY (1978)) and Hendry and Ungern-Sternberg (HUS (1981)) for U.K. data, have led to the formulation of error-correction models of income and consumption in other European countries. (Rossi and Schiantarelli (1982), Heinesen (1987)). Given the considerable empirical success of these models, one could reasonably conjecture that the breakdown in the Norwegian consumption function is due to a failure to develop data-coherent error-correction formulations of income and consumption (See however Biørn and Olsen (1989) for a recent contribution). However, the integration tests in section 3 indicate that there are fundamental problems with e.g. a DHSY analogue for our data. This is because the consumption income ratio is estimated to be non-stationary, and hence plim $\hat{\mu} = 0$ in the error-correction model (4-1):

$$\text{(4-1)} \qquad \Delta c_t = \alpha + \underline{\beta}\Delta\underline{z}_t - \mu(c-y)_{t-1} + \text{seasonals} + e_t$$

$\underline{z}_t$ is a vector of (weakly) exogenous integrated stochastic regressors. e_t is a (designed) well-behaved residual.

Of course, the non-stationarity in $(c-y)_t$ may reflect measurement errors in income (capital gains etc.), as well as economic behaviour. In the following, we will not try to discriminate between the causes of non-stationarity in the savings ratio, and we do not correct the income series for measurement errors.

Another caveat is that $c-y \sim I_0(1)$ does not necessarily imply that $c-\gamma y \sim I_0(1)$, with γ different from unity. However, judging from cointegration regression (4-2), $\gamma = 1$ is indeed a reasonable approximation:

(4-2) $c = 1.03y + \text{constant} + \text{seasonals}$

1968 1 - 1988 4 $T = 84$ $R^2 = 0.96$ $DF = 2.98$ $ADF = -1.97$[3)]

With these temporal properties in mind, it is possible to proceed either with a model in differenced data only, or to formulate a more general error-correction model with additional levels variables. The first approach is optimal only if a cointegrating vector cannot be established , the second in contrast presumes that (at least) one co-integrating vector exists between the augmented set of level variables. Hence, according to the second approach we have:

$$(4\text{-}2) \qquad \Delta c_t = \alpha + \underline{\beta}\Delta \underline{z}_t - \mu(c-y)_{t-1} + \underline{\phi}\ \underline{x}_{t-1} + \text{seasonals} + e$$

instead of (4-1). $\underline{x}_t$ is a vector of $I_0(1)$ series. Theoretically, (4-2) is an error-correction model, corresponding to the situation in which both $c_t - y_t$ and $\underline{x}_t$ are integrated but cointegrated at the long-run frequency. In a stationary steady state:

$$(4\text{-}3) \qquad c - y = \alpha/\mu + \underline{\phi}\ /\mu\ \underline{x},$$

or for the savings ratio ($\bar{s}$):

$$(4\text{-}4) \qquad \bar{s} \approx - \frac{\alpha}{\mu} - \phi/\mu\underline{x}$$

Hence the elements in $\underline{x}$ may be interpreted to as the long-run determinants of the savings ratio.

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3) DF is the Dickey-Fuller test for the residuals from the co-integration regression (4-2). ADF is the augmented Dickey-Fuller test. (Four lags in the differenced residual). See e.g. Engle and Granger (1987), Engle and Yoo (1987).

5. The long-run relationship

In this section we concentrate on the long-run part of the model, i.e. the elements in the $\underline{x}$ vector in the error-correction formulation (4-2).

Wealth is a natural first choice for inclusion in the long-run part of the model on both empirical and theoretical grounds. If we make a number of assumptions, notably that there is a perfect credit market, that changes in asset prices are negligible, and that the individual has homothetic preferences, the traditional life-cycle theory predicts that the individual's real consumption will be proportional to his or her expected life-time resources, defined as the sum of the present real value of expected labour incomes and the real market value of net assets. The aggregate life-cycle consumption function is obtained by adding assumptions such as constant distribution of expected income and net assets among individuals and constant age distribution.

In (5-1) real consumption (C_t) is a function of current real income (Y_t) and real value of wealth (W_t).

$$(5\text{-}1) \qquad C_t = BY_t^{\alpha}W_t^{\beta} \qquad 0 < \alpha,\ \beta < 1$$

where B is a scaling factor.

(5-1) is consistent with a life-cycle framework when we assume that expected discounted future income is proportional to current income (see e.g. Ando and Modigliani (1963)).

Alternatively (5-1) could be rationalized by adopting a 'liqudity' theory of consumption, i.e. consumers are constrainded by their initial assets, see (Tobin(1972)).

If some of the assumptions underlying the simple life-cycle model are relaxed, additional explanatory variables become relevant. According to the life-cycle hypothesis an individual's marginal propensity to consume varies with age. Demographic variables may be introduced if there are changes in the age distribution. A recent study by Heller (1988) suggests that in seven OECD-countries the rise in the share of elderly in the population has contributed to reduced aggregate saving.

It is well known that the theoretical sign effect of interest rates on consumption is ambigious. Empirically, Boskin (1978) find a positive interest elasticity for U.S. consumption. Dicks (1988) also find a positive interest elasticity for U.K. consumption. A different interpretation of the inclusion of interest rates, is that nominal or real interest rates, act as proxies for credit constraints.

Imperfections in the capital and money market may cause liquidity constrained consumers. Norwegian households have experienced various degrees of credit rationing in the sample period. Even after the deregulation of the credit market in 1983, credit rationing may persist: Credit rationing may be the equilibrium outcome in markets with imperfect information. The market value of housing properties could be especially important for liquidity constrained households. Although the aggregate personal sector cannot easily convert their capital gains by selling, a price increase may provide extra borrowing power. The credit expansion variable of the consumption function in KVARTS (2-1) may act as a proxy for the effect of credit rationing.

Since we do not have data of expected discounted lifetime resources, empirical relationships of consumer behaviour often introduce variables which try to capture the degree of uncertainty, e.g. the level of unemployment and changes in the aggregate price level.

Tax structures may influence consumer behaviour because after-tax rates of return tend to differ from before-tax rates. This fact will discriminate future relative to current income. Whether such a distortion is important depends on the size of the tax wedge (the difference before and after tax rates of return) and the elasticity of consumption with respect to after tax rates of return.

Institutional factors such as the level of the social security benefits may also have effects on aggregate consumption. Feldstein (1974) found that social security benefits depress personal saving by 30-50 percent.

The theoretical consepts of income and wealth differ from the calculations in the National Accounts. This introduces measurement errors in the consumption function. For example, capital gains and losses are not included in the National Account's concept of household income. For instance if inflation is persistent, there will be losses and gains on monetary assets and liabilities which is not included in the national account's calculations of household income. This phenomenon has led to the concept of inflation-adjusted income, see HUS(1981). Another alternative is to include the changes in the price level as a proxy for the gains and losses due to inflation. Another source of measurement error is the lack of reliable estimates for the real value of household's total wealth. One alternative is to use proxies for wealth, such as liquid assets (see HUS(1981)).

Lack of a consistent estimate of total household wealth is also a problem for our study. For the components of financial wealth we have imperfect information on households stock of bonds and company shares at real market values, and equity in life assurance and pension funds. Our set of financial data therefore includes the money stock, liquid assets and liabilities. For physical assets there are no consistent data series on the market value of the housing stock. We have constructed a market price for houses by linking different data sources, see Brodin (1989).

In face of these data limitations, our next question is how to implement wealth effects in the consumption function. Our (narrow) measure of real wealth is:

$$W^{*} = (M + DEP - LI + HO \cdot PH)/PC$$

where:

W^{*} - Real value of wealth.

M - Households' share of the money stock.

DEP - Deposits in banks and other financial institutions.

LI - Liabilities, loans and mortgages to banks and other financial institutions.

HO - Real value of housing stock.

L - M + DEP, Liquid assets.

PH - Housing prices.

PC - Implicit deflator of total expenditure.

Alternatively, the wealth variable can be replaced by a marginal model of wealth holding. One could argue that the difficulties in wealth measurement are fundamental, and that one is better served by conditioning on the determinants of wealth, rather than a measure of wealth. However, in Norway, there is little consensus about the likely determinants of wealth holding. Among other things, there is considerable uncertainty about the behavioural effects of the financial deregulation in the early eighties. In particular, it is likely that this "change of regime" will affect any (marginal) model of households' wealth. On the other hand, a consumption function which is conditional on wealth variables, may be structurally invariant to credit market deregulation.

Note that a relevant restriction on (5-1) is the homogeneity constraint:

(5-2) $$\beta = 1 - \alpha$$

which implies that a one percentage increase in income and wealth increase consumption by one percent. (5-2) is also required to ensure long run consistency of C_t and W_t for a given development of Y_t. To show this we follow Molana (1987) and use the first order Taylor approximation:

(5-3) $$\exp(z) = \exp(\bar{z}) + x\exp(\bar{z})$$

where $z=\ln(Z)$, $\bar{z}=\ln(\bar{Z})$ and $x=z-\bar{z}$. Using (5-3) first for $\ln(Y/W)$ and then for $\ln(C/W)$ we obtain:

(5-4) $$\ln(Y/W) = H(1-h)+H\ln(Y/W)$$

(5-5) $$\ln(C/W) = D(1-d)+D\ln(C/W)$$

In (5-4), $h=\ln(H)$, where H is the steady state value of Y/W. Similarily in (5-7) $d=\ln(D)$, D is the steady state value of C/W. In principle, W_t follows the definition equation:

(5-6) $$\Delta W_t = Y_t - C_t$$

which is conveniently rewritten:

(5-7) $\Delta W_t/W_t = Y_t/W_t - C_t/W_t$

Using (5-4) and (5-5) and setting $\Delta W_t/W_t = \Delta W_t/W_{t-1} = \Delta \ln(W_t)$, we obtain:

(5-8) $\Delta \ln(W_t) = H(1-h) - D(1-d) + H\log(Y/W)_t - D\log(C/W)_t$

In steady-state: $\ln(Y/W)=h$ and $\ln(C/W)=d$, and so the accounting equation (5-6) implies that in a steady-state, the growth rate of wealth depends on the steady state ratios D and H only. Looking back at (5-1), it is immediately clear that this places a restriction on the consumption function, namely equation (5-2). If (5-2) holds, we can write:

(5-9) $D = BH^{\alpha}$

which implies:

(5-10) $\Delta \ln(Wt) = H(1-h) - bH^{\alpha}(1-\ln(B)-\alpha h) + Hh - Bh^{\alpha}(\ln B+\alpha h)$

so that the growth in wealth is again only a function of the steady state ratios H and D. Unless (5-2) is imposed, this is not the case.

Even though our data of consumption, income and wealth do not satisfy the budget constraint in (5-6), we think the consistency requirement (5-3) deserves attention in applied work.

Our next step is to establish a cointegrating equation. We start by investigating the temporal properties of some of the variables discussed above. Table 5.1 shows that, once we include seasonals, none of the variables appear to be non-stationary at the sesonal frequencies. At the zero frequency, the tests indicate that both wealth and liquid assets are trend-stationary. But inspection shows that this is wholly dependent on inclusion of the 1988 observations in the sample.

The nominal interest rate (R) is stationary on the long run frequency without any deterministic part, while the real interest rate (RR) is stationary both in the case without augmentation, and in the I,SD,TR case. If the interest rates are stationary they can clearly not explain the long run behaviour of consumption. With these temporal properties in mind we move on to tests of cointegration.

Table 5.1 HEGY-tests for "unit roots", 1969 1 - 1988 4

"F" and "t_i" (i=1,2,3,4) $Y_4 = \Pi_1 Y_{1-1} + \Pi_2 Y_{2-1} + \Pi_3 Y_{3-2} + \Pi_4 Y_{3-1}$ + augmentation + residual

"t_1" "t_2", "t_3" : $Y_4 = \Pi_1 Y_{1-1} + \Pi_2 Y_{2-1} + \Pi_2 Y_{3-2}$, only if "$t_4$" insignificant. + augmentation + residual

Variable	Augment. 1)	Π_1 "t_1"	Π_2 "t_2"	Π_3 "t_3"	$\Pi_3 \cup \Pi_4$ "F"	Π_4 "t_4"
log W	-	0.79	-1.27	-1.21	16.96*	-5.65*
	I	-1.76	-1.16	-1.15	15.13*	-5.25*
	I, SD	-1.52	-4.98*	-6.49*	70.20*	-6.05*
	I, SD, TR	-5.88*	-4.81*	-8.03*	78.05*	-5.74*
log L	-	1.85	-0.67	-0.80	4.34	-2.82*
	I	-0.84	-0.64	-0.82	4.20	-2.77*
	I, SD	-0.89	-6.02*	-5.28*	42.32*	-6.42*
	I, SD, TR	-3.53*	-5.91*	-5.86*	43.51*	-5.99*
log CR	-	1.76	-2.40	-3.83*	9.34*	-1.88
	I	0.84	-2.43	-3.77*	9.26*	-1.95
	I, SD	0.75	-3.27*	-4.53*	14.66*	-2.66*
	I, SD, TR	-1.96	-3.07*	-4.65*	13.79*	-2.20*
log R	-	4.21*	-3.08*	-3.27*	22.63*	-5.51*
	I	0.53	-3.09*	-3.33*	22.51*	-5.44*
	I, SD	0.54	-3.20*	-3.32	22.60*	-5.45*
	I, SD, TR	-2.92	-4.10*	-4.30*	29.73 *	-5.72*
RR	-	-2.00*	-5.15*	-5.26*	53.15*	-7.61*
	I	-2.05	-5.12*	-5.24*	52.71*	-7.58*
	I, SD	-2.02	-5.03*	-5.16*	51.36*	-7.48*
	I, SD, TR	-4.02*	-4.49*	-5.73*	48.67*	-7.02*

* means significant at the 5 % nivå.

I: Intercept, SD: Seasonal dummies, TR: Trend

1) In addition to determministic terms, the augmentation contains significant lags of Y4

In table 5.2 we have listed some potential cointegrating equations, together with conventional cointegration statistics. These test statistics assume that we have no non-stationary seasonal components in the data, which on balance seems realistic against the background of table 3.1 and 5.1

Table 5.2 Cointegrating equations, 1968 (1) - 1988 (4)

Dependent variable: $\ln C_t$

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Coefficient estimates								
Const	-0.34	1.13	-	0.52	1.16	0.86	-	0.64
$\ln Y_t$	1.03	0.52	0.71	0.54	0.53	0.64	0.78	0.34
$\ln W_t$		0.31	0.24	0.34	0.30			
$\ln L_t$								0.54
$\ln R_t$				-0.05				
RR_t					0.00			
$\ln CR_t$						0.24	0.19	
Diagnostic statistics								
σ	0.066	0.052	0.054	0.052	0.053	0.055	0.056	0.054
CRDW	2.33	2.94	2.83	2.92	2.96	2.95	2.90	2.55
DF	-10.87	-15.77	-14.28	-15.54	-15.96	-15.94	-14.98	-12.20
ADF(4)	-2.14	-3.66	-3.32[1)]	-3.16	-3.57	-2.08	-2.80	-2.59

1) Insignificant lags excluded.

The main results of this table is that the only equations with significant cointegration statistics are the ones in columns 2, 3, 5 and maybe 4. Consumption, income and wealth appear to co-integrate. It is clear from column 2 and 3 that the income and wealth elasticites are sensitive for the inclusion of an intercept. Especially the income elasticity of column 2 seems rather low.

From a theoretical point of view we would like consumption to be homogeneous of degree one in income and wealth. The sum of the income and wealth elasticity should therefore approximate one. This is the consistency condition (5-2) imposed by the budget constraint. From table 5.2 we can see that the elasticities of column 3 is close to one. However, closer inspection of the recursively estimated coefficients reveals that the elasticities of column 2 are the most stable coefficients. Figure 5.1-5.3 shows the recursively estimated coefficients of column 2.

Figure 5.1 Recursively estimated coefficient of constant (cf. column 2, Table 5.2)

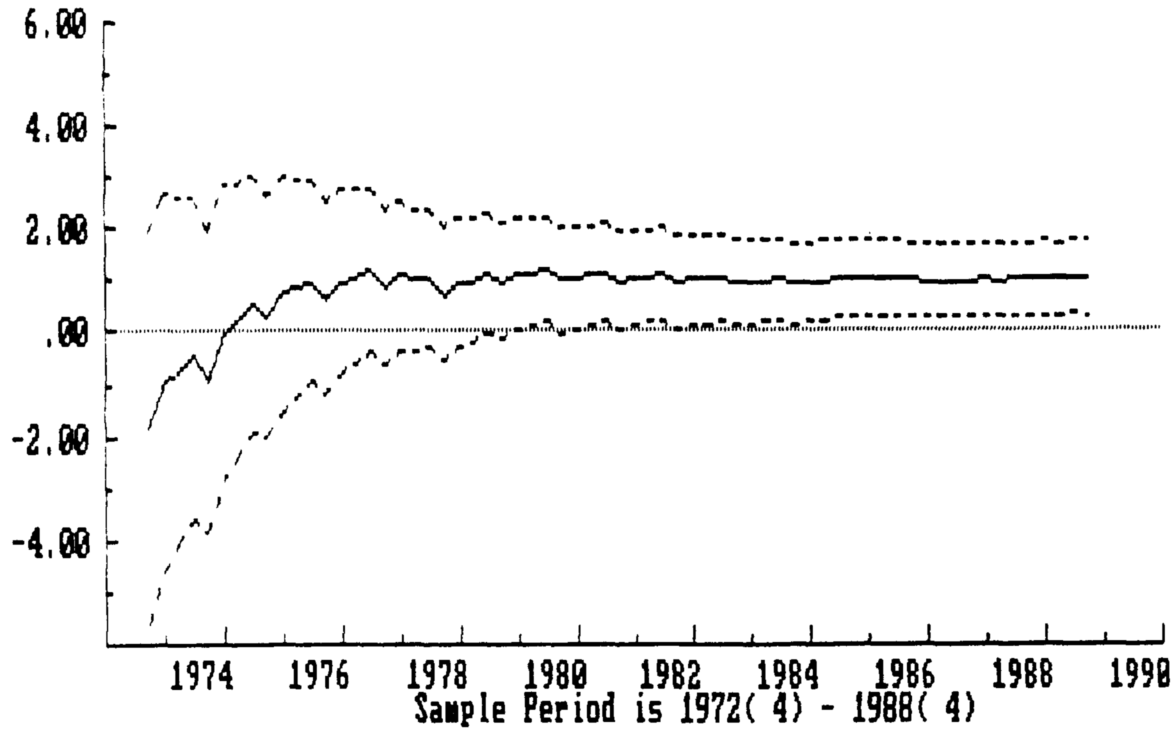


Figure 5.2 Recursively estimated coefficient of $\ln Y_t$ (cf. column 2, Table 5.2)

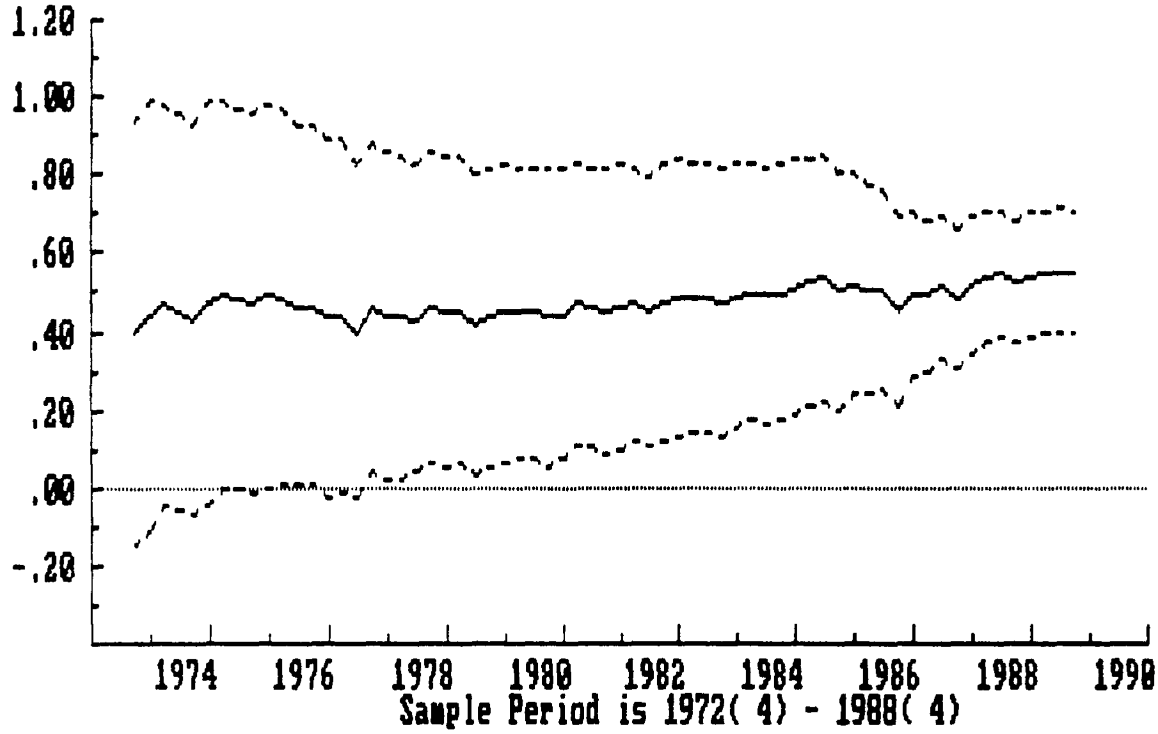
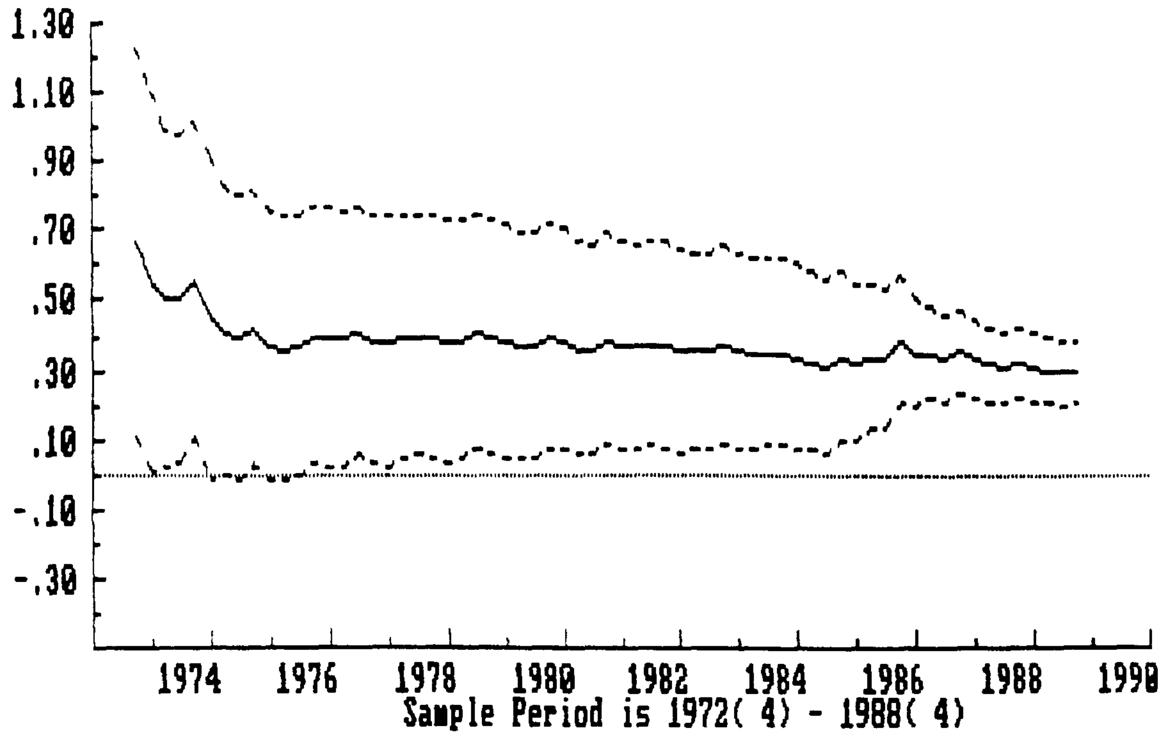


Figure 5.3 Recursively estimated coefficient of $\ln W_t$ (cf. column 2, Table 5.2)



From table 5.2, neither the nominal nor the real interest rates appear to be important long-run determinants of the the consumption funtion. As discussed above, this may be explained with the stationarity of the data. Columns 6 and 7 substitute the wealth variable by a measure of credit expansion. This is in line with the consumption function of the KVARTS model. However, the ADF statistic is insignificant and we cannot conclude that income and credit expansion are sufficient for explaining long-run consumption. We have mentioned that there are major problems in constructing a wealth measure. Another alternative is to use wealth proxies such as liquid assets, see Zellner et al (1965) and HUS (1981). This variable could also be given the interpretation of a pure liqudity effect, see Pissarides (1978). Columns 8 shows that consumption, income and liquid assets do not appear to co-integrate.

Eventhough the estimates from the static cointegrating regression is (super) consistent, left out short run dynamics may still cause substantial short run bias (Banerjee et al. (1986)). To supplement the static regression results, we therefore calculate the long-run coefficient from a dynamic model. A third alternative is to use band spectrum regression (Engle (1974)). Using spectral regression, we can remove the high frequency components altogether, and estimate the transformed equation for the relevant long run frequencies.

In table 5.3 we give the results of the three different methods for the equations in column 2 and 3 of table (5.2). For the dynamic model we have used an autoregressive distributed lag model with 6 lags in consumption, income and wealth. We have not tried to design a parsimonious model of the dynamic model. This will be done in section 6. If the restrictions of the parsimonious model are valid, the long-run coefficients should be fairly similar. In the frequency band regression we have defined the high frequencies as the band between 0 and 0.125, which implies that cycles of less than 8 quarters are high frequency cycles. The third column in table 5.2 excludes narrow bands around the seasonal fre-quencies 0.25 and 0.5. This regression presents a method for avoiding the seasonality issue. We note that all the three methods give quite similar results of the income and wealth elasticities.

Table 5.3 Long-run parameters, dependent variable: ln C_t

	Static[1]		Dynamic[2]		Frequency band regression: Seasonal frequencies excluded[3]		Only low frequencies[4]	
Const	-	1.30	-	1.30 (0.18)	-	1.13	-	1.20
ln Y_t	0.71	0.52	0.77 (0.09)	0.55 (0.04)	0.72	0.53	0.73	0.52
ln W_t	0.24	0.30	0.19 (0.08)	0.28 (0.02)	0.22	0.29	0.22	0.31
D1		-0.12		-0.09 (0.04)				
D2		-0.08		-0.07 (0.03)				
D3		-0.10		-0.07 (0.03)				

1) Static regression

2) Autoregressive distributed lag regression of order 6,6,6.

3) The frequencies [0.23,0.27] and [0.46,0.50] have been excluded.

4) The frequencies [0, 0.125] have been included.

Notes: Standard errors in parentheses.

In table 5.2 we have implicitly assumed that the cointegrating vector is unique. However, with three I(1) variables there may be 2 cointegrating vectors (Engle and Granger (1987)). In this case it is not clear which vector e.g. OLS is actually picking up. Also, the 2-stage procedure suggested by Engle and Granger (1987) is not fully efficient since the error-correction formulation in the second stage utilize the residuals from only one stationary linear combination. A recent paper by Johansen (1988) suggests a maximum likelihood estimation procedure which offers solution to this problem.

To outline the Johansen procedure, let X_t denote the nx1 vector of variables of interest and write the Vector Autoregressive Representation (VAR) in the following interim multiplier form:

$$(5\text{-}11) \qquad \Delta X_t = \sum_{i=1}^{k-1} \Pi_i \Delta X_{t-i} - \Pi_k X_{t-k} + V_t \text{ where } V_t \sim IID(0,\Omega).$$

The matrix Π_k is nxn and has rank $p \leq n$ in general , so that p is the number of cointegrating vectors.

Now we define two matrices α, β both of which are nxp such that

$$\Pi_k = \alpha \beta'$$

and so the rows of β form the p distinct cointegrating vectors.

The likelihood ratio test statistic for the hypothesis that there are at most p cointegrating vectors is

$$-2 \ln Q = -T \sum_{i=p+1}^{N} \ln(1 - \hat{\lambda}_i)$$

where $\hat{\lambda}_{p+1} \ldots \hat{\lambda}_N$ are the N-p smallest squared canonical correlations between the X_{t-k} variables and the stationary ΔX_t variables, corrected for lags in ΔX_t.

Johansen and Juselius (1989) show that care must be taken when introducing a constant and dummies into this framework. For example, if entered unrestricted, the constant will act as a linear trend and hence form a non-stationary part of the model. Alternatively, the constant may enter restricted, as a part of the stationary long run model. Different sets of critical values apply to the two ways of introducing the constant. From our experience with this data set it is most reasonable to include the constant in the equilibrium relationship. Hence, we use the critial values in table D.3 in Johansen and Juselius (1989)[4].

Table 5.4 Johansen tests for the number of cointegrating vectors, 1968 1 - 1988 4, k = 5

Case 1	$\ln C_t$, $\ln Y_t$, intercept + seasonal dummies	
	$-2 \ln Q$	Critical values
$p = 0$	18.85	20.17
$p \leq 1$	5.71	9.09
Case 2	$\ln C_t$, $\ln Y_t$, $\ln W_t$	
	$-2 \ln Q$	Critical values
$p = 0$	25.80	23.80
$p \leq 1$	11.48	12.00
$p \leq 2$	0.05	4.20
Case 3	$\ln C_t$, $\ln Y_t$, $\ln W_t$, intercept, seasonal dummies	
	$-2\ln Q$	Critical values
$p = 0$	43.25	35.07
$p \leq 1$	14.30	20.17
$p \leq 2$	2.99	9.09

----

4) The Johansen procedure is part of the TEST macro system developed in TROLL by Kjell Bleivik, Bank of Norway.

Table 5.5 Eigenvalues and eigenvectors

Case 1 ln C_t, ln Y_t, intercept + seasonal dummies

Eigenvalues (0.15, 0.07, 0.00)

Eigenvector
$$\left[\begin{array}{rrr} -7.19 & -25.15 & -7.19 \\ 2.55 & 28.19 & 2.55 \\ \hdashline 50.82 & -32.96 & 50.85 \end{array}\right]$$

Case 2 ln C_t, ln Y_t, ln W_t

Eigenvalues (0.17, 0.13, 0.00)

Eigenvector
$$\left[\begin{array}{rrr} 27.01 & -45.17 & -4.65 \\ -12.80 & 37.67 & -2.40 \\ -11.82 & 6.24 & 6.04 \end{array}\right]$$

Case 3 ln C_t, ln Y_t, ln W_t, intercept, seasonal dummies

Eigenvalues (0.31, 0.13, 0.04, 0.00)

Eigenvector
$$\left[\begin{array}{rrrr} 93.24 & 36.60 & 17.26 & -10.51 \\ -50.34 & -21.69 & 6.74 & -0.26 \\ -25.66 & -12.72 & -13.47 & 8.01 \\ \hdashline -130.92 & 4.00 & -84.74 & 14.68 \end{array}\right]$$

The Johansen tests in table 5.4 seem to confirm our previous findings that consumption and income do not co-integrate (case 1). On the other hand consumption, income and wealth appear to cointegrate (case 2 and 3). We also note that we cannot reject the null-hypothesis that the cointegrating vector is unique both in case 2 and in case 3. In table 5.5 we find that the VAR-model with the highest eigenvalue is the model in case 3. If we normalize column 1 in table 5.5 which corresponds to the eigenvector with the highest eigenvalue, we get (1.00, 0.47, 0.44) and (1.00, 0.54, 0.28, 1.40), where the last row in case 3 is the intercept. By comparing these normalized eigenvectors with the long-run parameters of table 5.3, we find that in case 3 the long-run parameters are very close to the ones in table 5.5, while they are quite different in case 2. The bottom line is that static and dynamic regression and the Johansen procedure, give very similar results, provided we include an intercept and the three seasonal dummies.

6. Dynamic econometric modelling

The previous section showed that income and wealth are the long-run determinants of consumption. From the representation theorem in Engle and Granger (1987) we know that a cointegrating vector can be represented by an error-correction model (ECM) which is not liable to the problems of 'spurious regression'. There are two different approaches for estimating ECM. One alternative is to use the two-stage procedure of Engle and Granger. This procedure is quite straightforward. In the first stage we test the cointegration hypothesis, cf section 5. If cointegration is confirmed we use the residuals from the cointegrating equation in place of the levels term in the ECM. Engle and Granger show that with this procedure the true parameter values converge at a faster speed than standard econometric estimates. The other alternative is to include the levels of the cointegrating variables in the ECM. As already stated the cointegrating parameter vector may be seriously biased in small samples.

Table 6.1 Dynamic models

Estimated period: 1968 2 to 1983 4. Post sample period: 1984 1 to 1988 4.
Dependent variable $\Delta_4 c$

	(1)	(2)	(3)	(4)	(5)
Constant	0.03	0.07	1.23	-	0.02
	(3.08)	(6.58)	(5.97)		(2.96)
$\Delta_4 c_{t-1}$	0.22	0.17	0.15	-	0.25
	(2.31)	(2.30)	(1.99)		(2.66)
$\Delta_4 y_t$	0.41	0.36	0.32	0.61	0.51
	(4.03)	(4.54)	(3.41)	(5.67)	(4.85)
$\Delta_4 w_t$	0.25	0.19	0.14	0.27	-
	(2.09)	(2.00)	(1.43)	(2.28)	
$z1_{t-4}$	-0.40	-	-	-	-
	(-3.96)				
$z2_{t-4}$	-	-0.82	-	-	-
		(-7.45)			
$(c-y)_{t-4}$	-	-	-	-0.67	-0.41
				(-5.87)	(-3.98)
$(w-y)_{t-4}$	-	-	-	0.01	-
				(3.34)	
c_{t-4}	-	-	-0.86	-	-
			(-7.78)		
y_{t-4}	-	-	0.46	-	-
			(3.98)		
w_{t-4}	-	-	0.24	-	-
			(2.79)		
q_{1t}	-0.04	-0.10	-0.11	-0.09	-0.04
	(-2.78)	(-6.29)	(-6.61)	(-5.30)	(-2.75)
q_{2t}	-0.03	-0.07	-0.07	-0.05	-0.03
	(-2.85)	(-6.29)	(-6.30)	(-4.56)	(-2.65)
q_{3t}	-0.04	-0.09	-0.09	-0.09	-0.05
	(-3.16)	(-6.52)	(-6.42)	(-5.46)	(-3.25)
Vat_t	0.07	0.06	0.06	0.09	0.07
	(4.68)	(4.77)	(4.73)	(3.68)	(4.55)
$d78_t$	-0.04	-0.03	-0.03	-0.06	-0.05
	(-1.79)	(-2.09)	(-2.01)	(-2.60)	(-2.17)
Diagnostics:					
σ	0.019	0.015	0.015	0.020	0.020
DW	1.70	2.00	2.10	1.74	1.70
$X^2(20)$	27.20	58.80*	76.20*	59.80*	64.40*
$F_{CH}(2,T1-k)$	1.14	1.67	1.31	1.44	2.16*
$t(19)$	-1.93	-1.11	-0.50	3.30*	2.27*
$F_{AR}(1,T1-k-1)$	1.50	0.01	0.24	0.74	2.22
$F_{AR}(5,T1-k-5)$	2.55	0.43	0.23	1.45	1.49
$F_{ARCH}(1,T1-k-1)$	0.00	0.35	1.38	2.51	0.57
$F_{ARCH}(5,T1-k-5)$	0.26	0.50	0.46	2.64	0.65
$X_N^2(2)$	0.53	0.40	4.39	0.95	3.96

* - significant at the 5 percent level
Z1 - Residual from the co-integration equation without intercept
Z2 - Residual from the co-integration equation with intercept

Statistics:

T_1, T_2, k = The number of observations, inside (T_1) and outside (T_2) sample, the number of explanatory variables (k).

σ = The percentage standard error of the regression

DW = Durbin-Watson statistic.

R^2 = Multiple correlation coefficient

$F_{ARCH}(j, T_1-k-j)$ = F-form test of LM-test of ARCH heteroscedasticity of order j, Engle (1982).

$F_{AR}(j, T_1-k-j)$ = F-form of the j'th order autocorrelation, starting at lag 1, Harvey (1981), Kiviet (1986).

$X^2{}_N(2)$ = Chi-square test of normality, Jarque and Bera (1980).

$X^2(T_2)$ = Post sample goodness of fit test based on the one step ahead predition residuals. (cf Hendry (1979)).

$F_{CH}(T_2, T_1-k)$ = F-form of the Chow-test of parameter stability. Chow (1960).

$t(T_2-1)$ = Test of zero forecast innovation mean (cf Hendry (1989)).

Table 6.1 shows different error correction models, using fourth differences and the level terms lagged four quarters. The estimation in column 1 and 2 is based on the two-stage procedure by Engle and Granger using the residuals from the cointegrating equations in column 2 and 3 in table 5.2. Additional I(0) variables such as changes in interest rates, consumer prices, or unemployment did not contribute to the explanation of the short run dynamics of the models.

The models in the table are restrictive since the elasticities of the various wealth components are not necessarily equal. This could also be relevant for the various income components. However, experiments with disaggregated income and wealth components were not successful. At the bottom of the table we show test statistics for residual autocorrelation, residual heteroscedasticity, various stability tests and finally tests of normality of the residuals.

As already stated column 1 uses the residuals from the cointegrating equation corresponding to column 3 in table 5.2. This is the equation with the most plausible elasticities form a theoretical point of view.

All the estimates are significant except for the dummy representing the wage and price freeze in 1978-1979. None of the misspesification test statistics are significant. Note that the stability tests have low values compared to the ones in table 2.1.

In column 2 we have used the residuals corresponding to column 2 in table 5.2. The standard error is reduced from 1.9% to 1.5%. The forecast X^2 is significant but this may only reflect better fit. The parameter estimates of the short run dynamics in the two models are quite similar. On the other hand, the error correction term of the model in column 2 is twice as big as the one in column 1. The speed of adjustment towards the long-run solution of the model is therefore much faster for the model in column 2.

The lagged levels of the long-run variables are included in the dynamic model in column 3. The implicit static long-run solution of this model is $c = 1.43 + 0.53y + 0.28w$ + sesonals. These coefficients are quite similar to that of the cointegration equation in, of column 2, table 5.2. However the parameter estimate of the annual change in wealth is no longer significant. Moreover the forecast X^2 is significant and slightly higher than in column 2.

If we impose the restriction $\alpha + \beta = 1$ of section 5, we obtain the results in column 4. In the terminology of HUS (1981) the levels term in column 4 correspond to error correction and integral control. The static long run solution of the model is $c = 0.98y + 0.02w$ + seasonals. Hence, the homogeneity restriction gives a large shift in the elasticities. Note also, that the t-test for zero forecast innovation mean is significant. The last test statistics indicates that the forecast errors have a systematic bias. Closer inspection of the forcast errors reveals a systematic underprediction of consumption in the post sample period.

An error correction model without wealth effects is shown in column 5, see e.g. DHSY (1978). From section 4 we recall that he HEGY tests showed that consumption and income are not cointegrated with cointegration parameter equal to one. Model misspesification is confirmed by the three stability tests although the LM-test for residual correlation and heteroscedasticity are insignificant.

In sum, both the cointegration results of the previous section and from table 6.1 support that inclusion of a wealth variable is essential in any serious attempt to reconstruct an empirical consumption function. However, we do not find support for imposing the homogeneity constraint leading to a HUS-specification of the wealth effect.

The models in columns 1-3 all satisfy the essential requirements for a reconstructed consumption function. They fit the data up to 1983 no worse than the defunct equations, (2 and 3 fit markedly better) and they all have insignificant Chow-tests over the "breakdown"-period.

The different models in table 6.1 may be regarded as parsimonious models of an autoregressive distributed lag (AD) model in consumption, income and wealth. However, we have not yet tested for the implicit zero and equality restrictions in table 6.1. In fact, starting out with a general AD, we ended up with the following equation, which we will hold as our preferred model.

Table 6.2 Final equation

Dependent variable $\Delta_1 c_t$
Estimation period: 1968 2 - 1983 4.
Post sample period: 1984 1 - 1988 4.

$R^2 = 0.97$ $\sigma = 0.015$ $DW = 1.85$

$F_{CH}(20,54) = 1.54$ $X^2(20) = 35,4^*$ $t(19) = -1.17$

$F_{AR}(1,53) = 0.69$ $F_{AR}(5,49) = 0.33$ $X_N^2(2) = 0.03$

$F_{ARCH}(1,52) = 1.84$ $F_{ARCH}(5,44) = 0.64$

* significant at the 5% level.

Variable	Coefficient	t-value
const	0.06769	7.07204
$\Delta_1 c_{t-1} - \Delta_1 c_{t-4}$	-0.19226	-2.78316
$\Delta_1 y_t$	0.35771	3.17146
$\Delta_1 w_t + \Delta_1 w_{t-2}$	0.19036	2.05184
$z2_{t-1}$	-0.59490	-4.78894
q_{1t}	-0.09984	-4.15653
q_{2t}	-0.09265	-11.25038
q_{3t}	-0.07167	-3.88822
Vat_t	0.06510	5.33970

x) z2 - Residual from the cointegration equation in table 5.2, column 2.

Figure 6.1 shows the observed and fitted values for this model in the post-sample period together with the forecast interval (two times the standard error). All the predicted values lie well inside this forecast interval.

The fitted values for the whole sample period are shown in figure 6.2 with estimated parameters from 1968 1 - 1983 4. The overall picture is that the model fits the data well both in the estimation period and in the post sample period. The model has some problems with the years around the wage and price freeze of 1978/1979. In spite of the inclusion of a dummy variable for this period, there are large residuals in this period. A substantial revision in the quarterly national accounts took place in 1978. This could be another explanation for the large residuals. The large residuals in 1988 may in the same way be explained by the wage freeze in 1988/89.

Figure 6.1 Actual and predicted values of the model in table 6.2

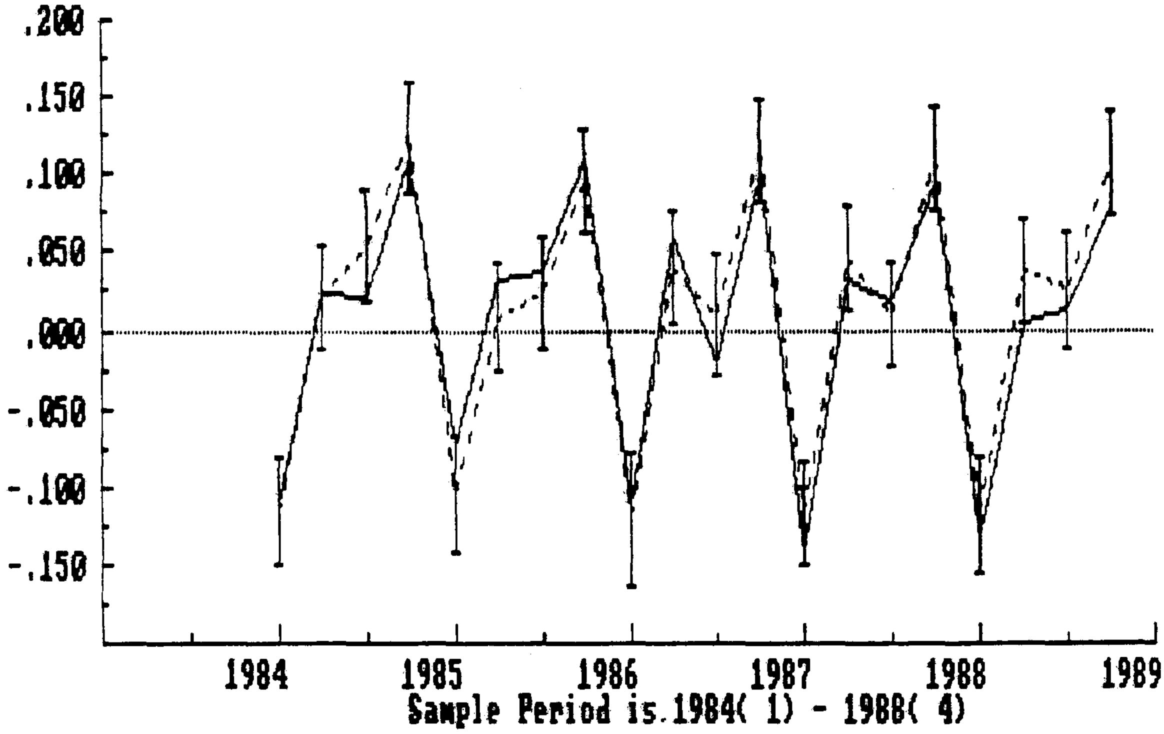
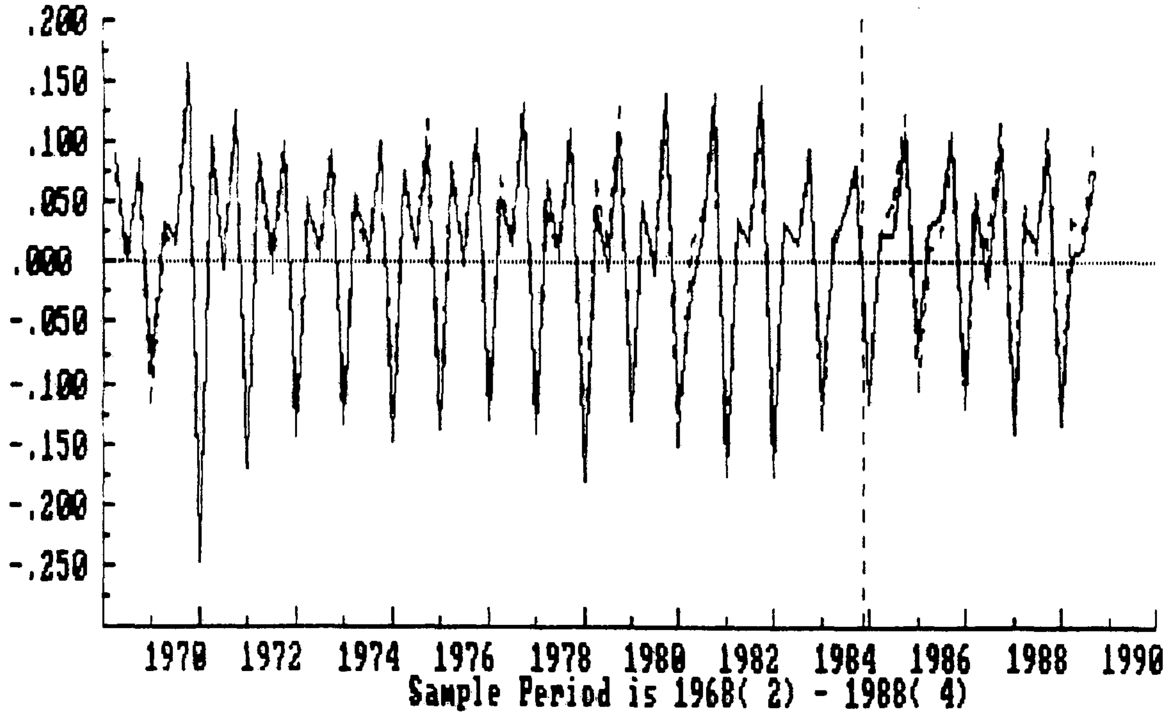


Figure 6.2 Fit of the model in table 6.2. Estimation period 1968 2 - 1983 4



The one-step residuals plotted in figure 6.3 together with their standard error, show that there is no increase in the residuals after 1980. In contrast to the defunct equations in the KVARTS and RIKMOD models, there is no increase in the residuals after the deregulation of the housing and credit market in 1983. In our view figure 6.2 shows that the model is structurally invariant with respect to the deregulations of the housing and credit market.

Additional evidence for structural invariance is provided by the recursively estimated parameters in figure 6.4 - 6.7. There is no evidence of permanent shifts in the parameters after 1980. However the estimated coefficient for the annual change in wealth shows a temporarily change around 1984/1985, but this is probably not significant.

Figure 6.3 One step residuals of the model in table 6.2

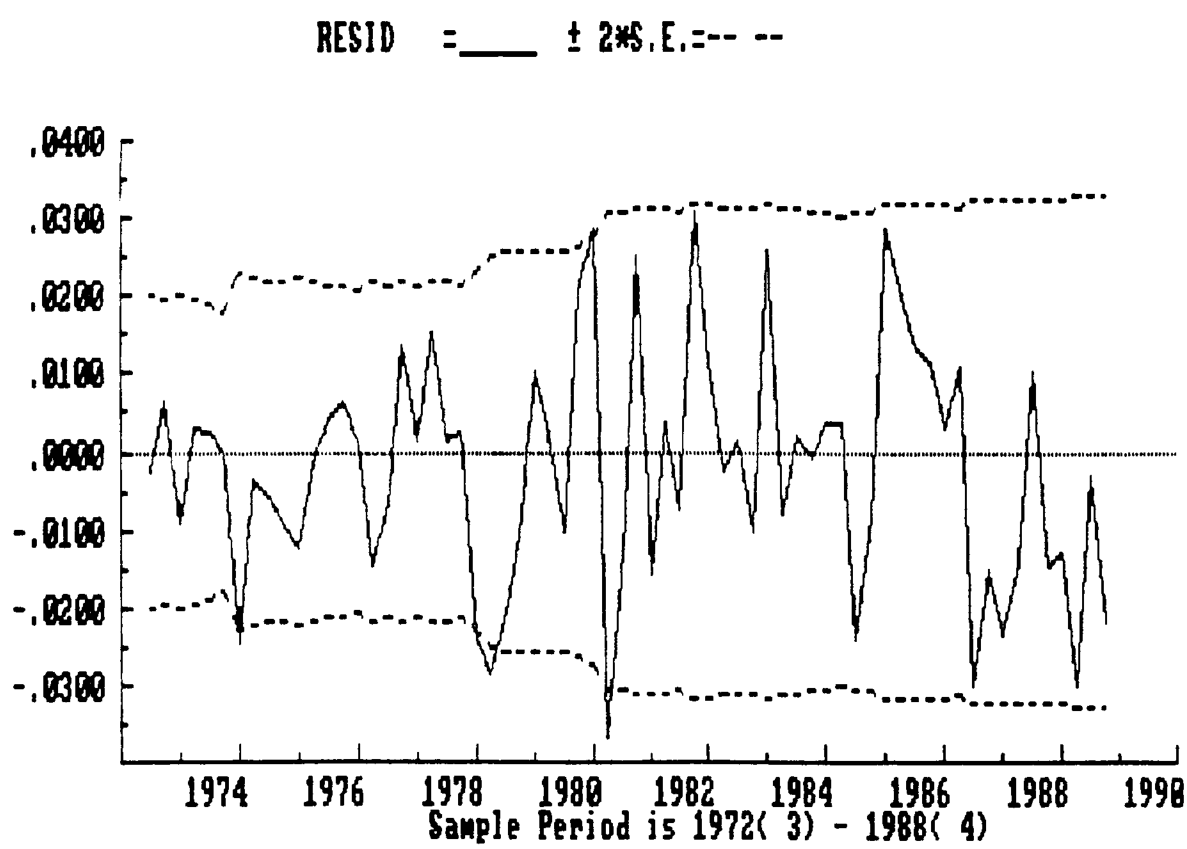


Figure 6.4 Recursively estimated coefficient of $\Delta_1 c_{t-1} - \Delta_1 c_{t-4}$ in table 6.2

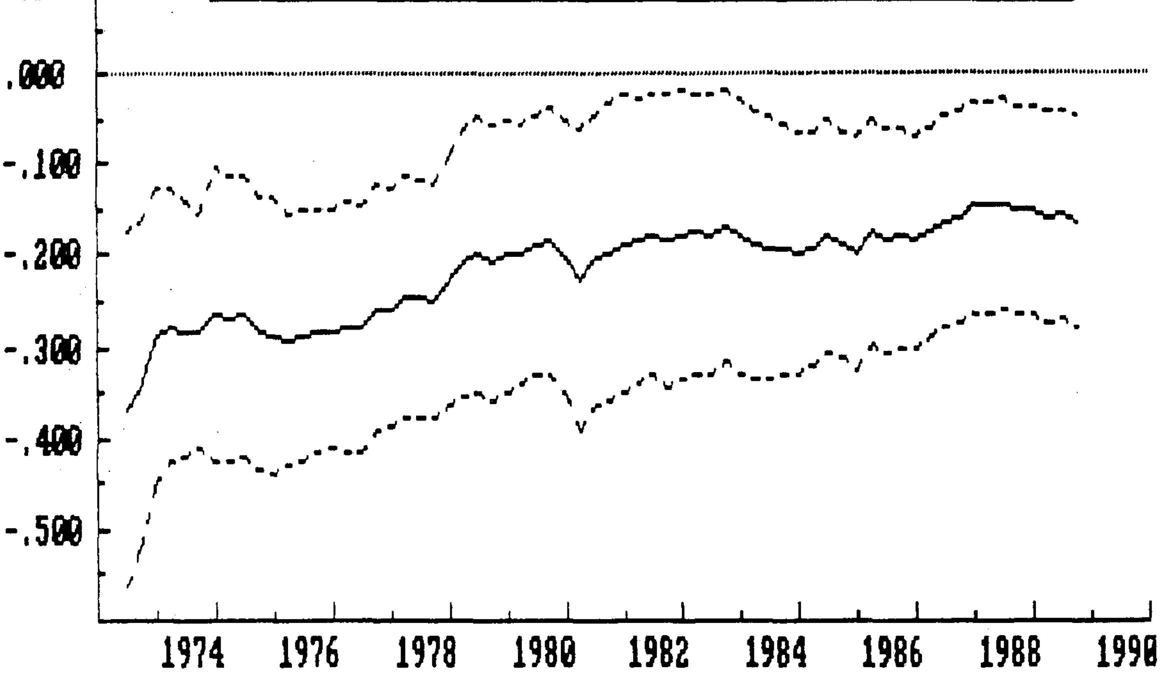


Figure 6.5 Recursively estimated coefficient of $\Delta_1 y_t$ in table 6.2

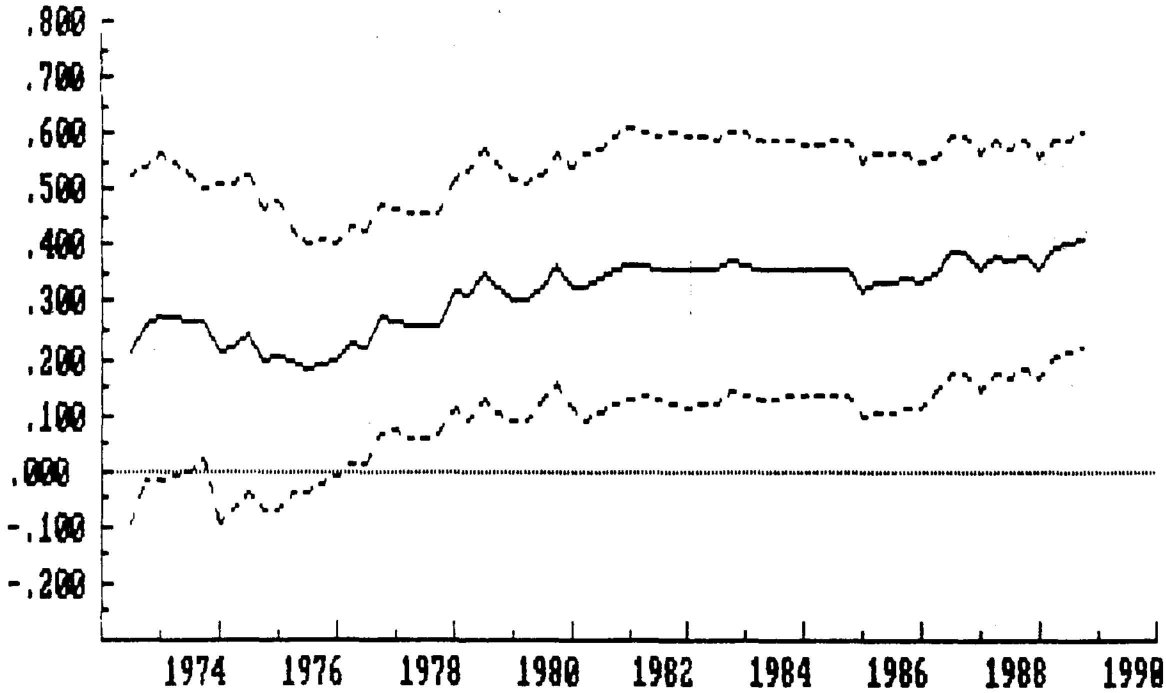


Figure 6.6 Recursive estimated coefficient of $\Delta_1 w_t + \Delta_1 w_{t-2}$ in table 6.2

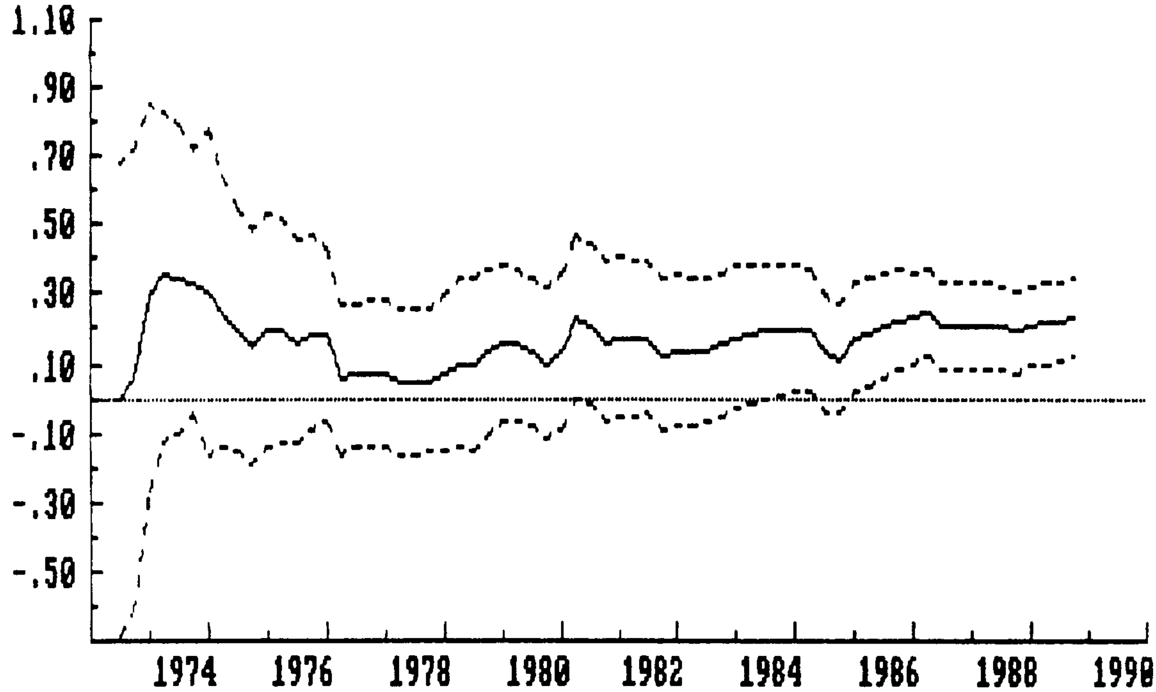
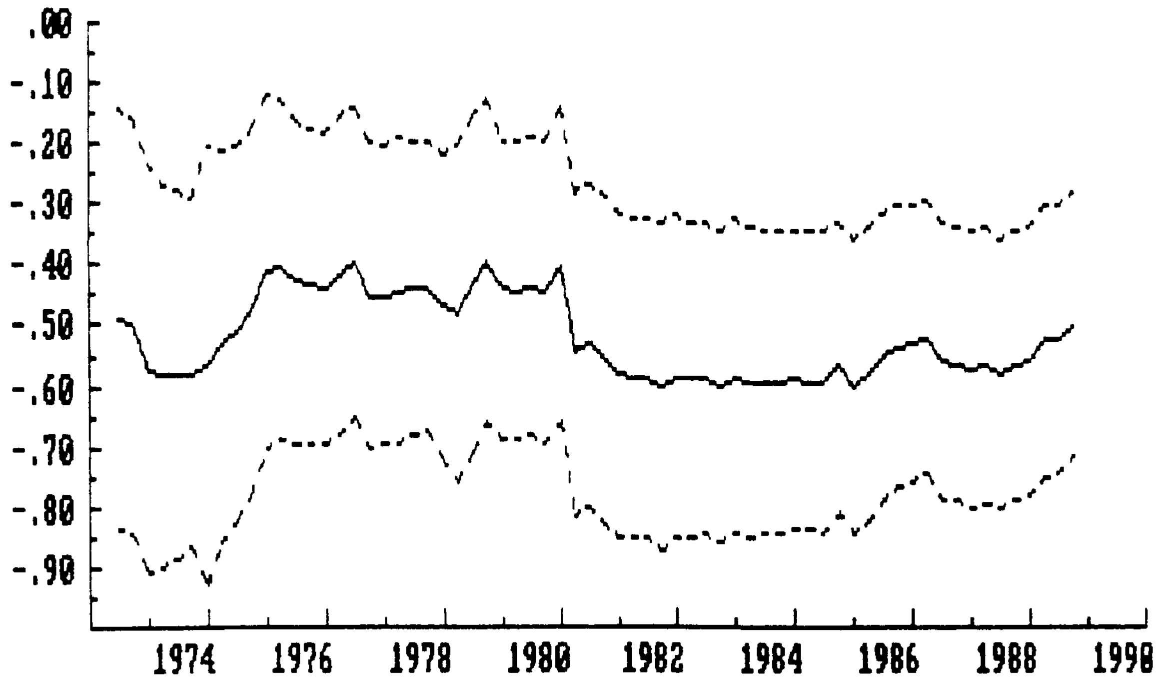


Figure 6.7 Recursively estimated coefficient of $z2_{t-1}$ in table 6.2



The model in table 6.2 is:

(6-1) $$\Delta_1 c = \gamma_0 + \gamma_1(\Delta_1 c - \Delta_1 c_{t-4}) + \gamma_2 \Delta_1 y_t + \gamma_3(\Delta_1 w_t + \Delta_1 w_{t-2}) + \gamma_4(c - k - \alpha y - \beta w)_{t-1} + \text{seasonal dummies.}$$

The steady state solution to (6-1)is given by $\Delta_1 c = \Delta_1 y + \Delta_1 w = g$:

(6-2) $$C = \exp[(g(1 - \gamma_2 - 2\gamma_3) - k\gamma_4) / \gamma_4]Y^{\alpha}W^{\beta}$$

If the parameters in (6-2) is substituted with the estimated coefficients in table 6.2, and g=0.005, we find the following long-run solution:

(6-3) $$c = 0.99 + 0.52y + 0.31w$$

A recurrent issue in the estimation of consumption functions is the simultaneity problem in consumption, income and wealth. It should be noted that the consistency of Engle and Grangers two-step procedure implies that there is no large sample simultaneity bias in the estimation of the long run parameters. We therefore concentrate on the estimation of the short run effect of income. Since $\Delta_1 y$ is not lagged, we have assumed that $\Delta_1 y$ is weakly exogenous (Engle et al (1983)). An indirect test of the exogeneity assumption is to investigate parameter constancy. Large shifts in the parameters might be an indication of a non valid exogeneity assumption. A direct test of weak exogenity is to estimate the model by instrumental variables and compare those parameter estimates with those obtained from ordinary least squares. Table 6.3 gives the results of instrumental variable estimation of the model in Table 6.2. The estimated coefficients are negligibly different from those in Table 6.2.

Table 6.3 <u>Instrumental variable estimation</u>

Dependent variable $\Delta_1 c_t$

Endogenous variable $\Delta_1 y_t$, $(\Delta_1 w_t + \Delta_1 w_{t-2})$

Instrumental variables: $\Delta_1 c_{t-1}$ $\Delta_1 c_{t-2}$, $\Delta_1 c_{t-3}$, $\Delta_1 y_{-1}$, $\Delta_1 y_{t-4}$, $\Delta_1 w_{t-2}$, $\Delta_1 w_{t-3}$

Estimation period : 1969 1 to 1983 4

Post sample period : 1984 1 to 1988 4

$\sigma = 0.015$ DW = 1.82 $\chi^2_{AR}(1) = 0.49$ $\chi^2_{AR}(5) = 1.25$

$\chi^2_{ARCH}(1) = 1.76$ $\chi^2_{ARCH}(5) = 3.75$ $\chi^2_S(5) = 2.2$

$\chi^2(20) = 31$ $\chi^2_N\ (2) = 0.03$

*) significant at the 5% level.

Variable	Coefficient	t-value
const	.07301	4.9416
$\Delta_1 c_t - \Delta_1 c_{t-4}$	-.17074	-2.0074
$\Delta_1 y_t$	.43198	1.4502
$\Delta_1 w_t + \Delta_1 w_{t-2}$	.17025	1.3756
$z2_{t-1}$	-.57596	-4.3595
q_{1t}	-.11142	-3.4862
q_{2t}	-.09239	-9.6049
q_{3t}	-.08434	-2.1964
Vat_t	.06376	4.8424

<u>Statistics</u>:

T_1, T_2 = The number of observations, in and outside sample.
K = the number of parameters. $\chi^2_S(m)$ = Sargan's (1964) Chi-square test of the validity of m overidentifying instruments. $\chi^2_{AR}(j)$ - Chi-square test of j'th order autocorrelation. $\chi^2_{ARCH}(j)$ - Chi-square test of j'th order heteroscedasticity. For the other statistics cf Table 6.1.

Finally, we turn to the issue of encompassing (Hendry and Richard (1987)). The interesting question here is whether our model encompasses the earlier consumption functions using their data. The encompassing tests therefore only use data for the period 1969-83.

(6-4) $$\left(\frac{c}{y}\right)_t = \alpha_0 + \alpha_1\left[\sum_{i=0}^{3}(1-0.25\ i)\ \frac{\Delta y_{t-i}}{y_{t-i-1}}\right] + \alpha_2 0.25\left[\sum_{i=1}^{4}\left(\frac{c}{y}\right)_{t-i}\right] + \underline{\alpha}_3\underline{Q}_t$$

(6-5) $$\left(\frac{c}{y}\right)_t = \beta_0 + \beta_1(\Delta_1 c_t - \Delta_1 c_{t-1}) + (\beta_2 - 1)\Delta_1 y_t + \beta_3(\Delta_1 w + \Delta_1 w_{t-2}) + (\beta_4 + 1)c_{t-1} + \beta_5 y_{t-1} + \beta_6 w_{t-1} + \underline{\beta}_7\underline{Q}_t$$

where $\underline{Q}_t$ is a vector with seasonal dummies and a dummy for the introduction of the VAT in 1969.

(6-4) is a logaritmic version of the RIKMOD consumption function (2-2). This functional form does not fit any worse than the linear model. In (6-5) we have transformed our model in table 6.2 to give log of the consumtpion-income ratio as the dependent variable. Furthermore we have replaced the error-correction term in model 6-2 by the relevant level terms.

Table 6.4 Encompassing test, 1969 1 - 1983 4

Model 1 - (6-5)

Model 2 - (6-4).

$\sigma_1 = 0.0157$ $\sigma_2 = 0.0182$ $\sigma_{joint} = 0.0158$

Model 1 v.Model 2	Form	Test	Form	Model 2 v.Model 1
0.66	F(2,47)	Joint Model	F(6,47)	3.86
3.21	F(2,47)	Critical Values	F(6,47)	2.26

Table 6.4 shows the results for the encompassing F-test of Hendry and Richard op.cit. It is clear from table 6.4 that our preferred model encompasses the RIKMOD consumption function.

Hence, in the hypothetical situation that both models had been formulated on data up to 1983, one would have been better off choosing our specification. Surely this is a powerful property of a model explicitly designed to fit the data in the breakdown period, 1984-88. In particular the encompassing test show that we have not thrown away information in the pre-breakdown period. On the contrary our model accounts for more of the variation in the data than the contending model.

7. Conclusions

The empirical evidence in this paper may be summarized as follows. Section 2 shows that consumption and income are not cointegrated at the zero frequency. We interpret this as a a result of omitted variables in the cointegration equation. Section 4 summarizes some theoretical arguments for additional variables in the long-run equation, and we conclude that wealth is the most likely omitted variable. Section 5 shows empirically that wealth should be included in the cointegration equation.

Section 6 gives empirical results for dynamic models with estimated parameters that are relatively invariant to the changes in the credit and in the housing market. This shows that the predictive failure of existing consumption functions does not constitute conclusive evidence against a conditional modelling approach to private households consumption. This interpretation is strengthened by the fact that a model which does break down in the deregulation period is encompassed by our new specification, using data which does not include the breakdown period.

Appendix. Re-estimation of the KVARTS consumption function

The model is:

$$C \quad = \quad (\beta_0 + \sum_{i=0}^{7} \beta_i Y_{-i} + \sum_{i=0}^{3} \beta_i \left(\frac{\Delta CR}{P}\right)_{-i})\underline{Q}_i + \text{residual}$$

C = Total private consumption.

Y = Households' disposable income.

ΔCR = A credit expansion variable.

$\underline{Q}$ = Seasonal dummies etc.

P = The private consumption deflator.

In Jensen and Reymert (1984), the distributed lags are PDL's with a "tail" restriction for income and a "head" restriction for the credit variable. These restrictions are maintained in the re-estimation reported in Bowitz, Jensen and Knudsen (1987).
The basis for the stability tests in table 2.1 is a re-estimation for the period 1968 4 to 1983 4. For comparison with Bowitz et.al. (1987), the consumption and income data are in fixed 1980-prices, P = 1 in 1980.

The results are (t-values in brackets):

$$\begin{aligned} C = \{\text{constant} &+ \underset{(0.30)}{0.04}\, Y + \underset{(1.87)}{0.08}\, Y_{-1} + \underset{(8.94)}{0.12} Y_{-2} + \underset{(11.41)}{0.14}\, Y_{-3} \\ &+ \underset{(5.70)}{0.15}\, Y_{-4} + \underset{(4.28)}{0.14}\, Y_{-5} + \underset{(3.65)}{0.10}\, Y_{-6} + \underset{(3.20)}{0.06}\, Y_{-7} \\ &+ \underset{(3.17)}{0.20}\, \frac{\Delta CR}{P} + \underset{(3.39)}{0.28}\, \frac{\Delta CR_{-1}}{P_{-1}} + \underset{(3.43)}{0.26}\, \frac{\Delta CR_{-2}}{P_{-2}} + \underset{(1.27)}{0.11}\, \frac{\Delta CR_{-3}}{P_{-3}}\} \cdot \text{seasonals.} \end{aligned}$$

1968 4 - 1983 4.
$\hat{\sigma}\%$ = 1.92 DW = 1.73 T = 61 k = 9.

The results show some intriguing differences from those reported in Bowitz et.al. (1987). Firstly, the shape of the polynominal in income allocates less weight to the first lags than in the published equation. As a result the mean lag here is 3.77 and 2.61 originally. The sum of the lag coefficients above is 0.83, which is low compared to 0.91 in the published results.

The picture is reversed for the credit variable, the long coefficient is 0.85 compared to 0.346 in the original equation. The overall fit of the equation is comparable to the published version.

The estimation period in Bowitz et.al. is 1967 4 to 1984 4. Our results are not affected by including 1984 in the estimation period. But we were unable to include the four quarters at the start of the period.

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