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Consumption and population age structure

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Abstract In this paper, the effects on aggregate consumption of changes in the age distribution of the population are analysed empirically. Economic theories predict that age influences individuals' saving and consumption behaviour. Despite this, age structure effects are rarely controlled for in empirical consumption functions. Our findings suggest that they should. By analysing Norwegian quarterly time series data, we find that changes in the age distribution of the population have significant and life-cycle-consistent effects on aggregate consumption. Furthermore, controlling for age structure effects stabilises the other parameters of the consumption function and reveals significant real interest rate effects.

Keywords Consumption · Demography · Cointegration

Jel Classification C51 · C53 · E21

1 Introduction

A key question in economics is whether changes in the age structure of the population affect macroeconomic variables such as aggregate consumption and the savings rate. Economic theory suggests that the influence could be substantial. The life cycle model of Modigliani and his collaborators, for instance, predicts that individuals' consumption and saving behaviour are functions of their age; an individual borrows as young, saves as middle-aged and dissaves when old (see Modigliani and Brumberg 1954 and Ando and Modigliani 1963). Aggregating up, changes in the age distribution over time could hence induce variations in a nation's private saving rate.

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The link between the savings rate and demography, with particular focus on the issue whether the elderly dissave, has been widely explored in the empirical literature. However, no consensus about the impact of a changing age distribution, neither when it comes to the direction nor the size of it, has been reached. At least in part, the discrepancies stem from the use of different types of data and methods.

For example, there is a tendency that studies based on aggregate macroeconomic data report significant and often numerically important age structure effects on consumption and savings. Thus, these studies usually confirm the predictions of the life cycle model, finding that savings decrease or aggregate consumption rises when the share of elderly persons in the population increases. Among these are Attfield and Cannon (2003), Higgins (1998), Horioka (1997), and Masson et al. (1996). Fair and Dominguez (1991) specifically find that prime-age people consume less relative to their income than other age groups on US data. However, they did not find life-cycle-consistent effects on the aggregate personal savings rate. On the other hand, studies using household survey data often find no or only modest effects on savings of changes in the age distribution of the population, e.g., Parker (1999) and Bosworth et al. (1991).

There are several explanations for the discrepancies between macro and micro evidence of the importance of age effects. Weil (1994) shows that if interactions between generations, such as bequest, are important, one would not expect the estimates from the two types of studies to be the same. Miles (1999), among others, suggests that part of the discrepancy is due to the fact that household survey studies often use savings rates which overestimate the values of pension assets. Furthermore, Deaton and Paxson (2000) emphasise that household survey data are likely to suffer from sample selection biases as these are based upon households and not on individuals. By taking account of the latter two arguments when estimating the savings–age profile based on UK household survey data, Demery and Duck (2001) derive a savings–age profile which is much more consistent with the life cycle model.

In this paper, we test for age structure effects on aggregate consumption in Norway, by estimating a consumption function which takes account of changes in the age distribution of the population. The model is estimated on quarterly time series data over the period 1968(3)–2004(4). The motivation is twofold. First, it is of importance to investigate the impact of a changing age distribution on Norwegian consumption. Similar to many other developed countries, Norway experienced low birth rates in the interwar years and a baby boom in the first decades after World War II. The small cohorts born in the interwar years are currently in the retiring phase, while the baby boomers are turning middle-aged. If life cycle saving behaviour applies, the ongoing change in the age distribution will thus put a downward pressure on Norwegian consumption, other things equal. Hence, if an age structure effect on aggregate consumption can be affirmed empirically, it will be of interest for both medium-term activity control and for a longer-term analysis of private saving and wealth. Using high-quality micro data, Halvorsen (2004) finds strong age effects on Norwegian household savings rates, while cohort effects are weak. This suggests that age structure may be an underlying structural aspect also of aggregate consumption, a possibility which we pursue in this paper.

A second motivation of the paper is methodological and is related to the modelling consequences of demographic changes. A demographic change of some magnitude is an example of a structural change, which can potentially overturn existing macroeconometric relationships and cause forecast failure. However, it can also lead to new knowledge, which in some instances can be accommodated by revision and extension of an existing model. Hence, the effects of structural changes are not always destructive. From this perspective, the paper investigates the possible incorporation of age structure in an existing consumption function. The incumbent model, specified by Brodin and Nymoen (1992), builds on cointegration between private consumption, income and wealth, a well-defined causal structure that has been stable over a 15-year period (see Eitrheim et al. 2002, for an analysis). As both theory and empirical macro studies suggest that changes in the age structure can affect aggregate consumption, it should hence be tested whether the changes in the age structure of the population is an omitted variable in the consumption function. Another variable that may influence consumption in the long run is changes in the real interest rate. Hence, we also test whether real interest rates should be included in the consumption function.

To preview our main findings, we find significant and numerically important age structure effects on Norwegian aggregate consumption. The effects are consistent with the life cycle model; consumption falls when the share of middle-aged persons in the population increases, other things equal. An important result in the paper is that the other parameters of the consumption function become more stable when we control for age structure effects. In addition, controlling for the effects of a changing age structure reveal significant effects of changes in the real interest rate on aggregate consumption.

While the intellectual rationale for controlling for age structure effects in a consumption function is clear enough, there is no clear-cut operational route. Different candidates for operational definitions of age structure changes have been suggested in the literature, and in Section 2 we motivate our preferred measure. Section 3 explains how the age structure variable, together with a real interest rate variable, represent an extension of the existing Norwegian consumption function. Section 4 presents the econometric evidence for a long run consumption function, which includes an age structure variable, and derives the corresponding error correction model. Section 5 concludes.

2 Age structure effects on consumption

Different approaches have been used in the empirical literature to test for age structure effects on aggregate consumption. The life cycle model suggests that the marginal propensity to consume (MPC) vary over the life cycle alongside changes in an individual's preferences, needs and income. On this basis, one would want to estimate an empirical relationship between aggregate consumption, income and wealth, where the MPC is allowed to change with the individuals' age when testing for age structure effects. However, disaggregate age group data for macroeconomic variables are rarely available, and hence when using this type of data one is usually not able to test for age-sensitive MPCs. It is thus more common in empirical macro studies to test for age structure effects on the average propensity to consume. This is typically done by including either one or several age structure variables in a regression model. If the variable(s) is(are) significantly different from zero, the changes in the age structure of the population is found to have effect on aggregate consumption.

Among the age structure variables, the 'dependency ratio', defined as the number of children and retired persons to those of working age, is often used to represent changes in the age structure; Leff (1969), Masson et al. (1996) and Horioka (1997) are among those using this variable. In this paper, we use a somewhat different age structure variable, namely, the number of persons in the 'prime-saver' age group to the rest of the adult population, used by McMillan and Baesel (1990) for a different purpose, namely, to identify age structure effects on US real interest rates, income, inflation and unemployment.

Similar to the 'dependency ratio', our age structure variable is an operational life cycle saving measure. As argued by McMillan and Baesel (1990), the ratio of 'prime-savers-aged' persons to the rest of the adult population may be a closer approximation to the life cycle ideal than the 'dependency ratio'. The 'prime savers' are assumed to be of middle age and have relatively high earnings at the same time as the size of their households are small; hence; their needs are smaller than when they were younger, and they are prone to start saving to secure quality of life after retirement. The middle-aged persons may therefore have a lower propensity to consume than both those who are younger and older.

The results in Erlandsen (2003) indicate that it is the age group of 50- to 66-year-old persons which, among the adult population, has the smallest average propensity to consume. This finding is consistent with the findings of Attfield and Cannon (2003) on UK data and also that of Fair and Dominguez (1991) who report results on US data showing that the lowest propensity applies to age groups that are 10 to 15 years younger. On this background, we have chosen to denote the 50- to 66-year-old Norwegians as the prime savers or the middle-aged, and we include the following age structure variable in the econometric model below:

$$AGE_t = \frac{(Population\ 50-\ 66\ years\ old)_t}{(Population\ 20-49\ years\ old\ and\ 67+\ years\ old)_t}. \tag{1}$$

The development of AGE over the sample period 1968(3)–2004(4) is shown in Fig. 1. The graph shows that the share of prime savers in the Norwegian population was declining from the mid-1970s until the beginning of the 1990s. From then on, AGE has been increasing, as the baby boom generation is now turning middle-aged. Given the prediction that middle-aged persons save more than the rest of the population, AGE is expected to enter the consumption function with a negative coefficient.[1]

3 Extending the consumption function

Up until the beginning of the 1980s, Norwegian private consumption was, according to the consensus view, well represented by a (log-)linear model between

[1]We acknowledge that any one-dimensional measure of age structure can be contested and the alternatives considered. However, the results do not depend critically on how age structure is represented, and in Erlandsen (2003), we include the 0–19 age group in the numerator.

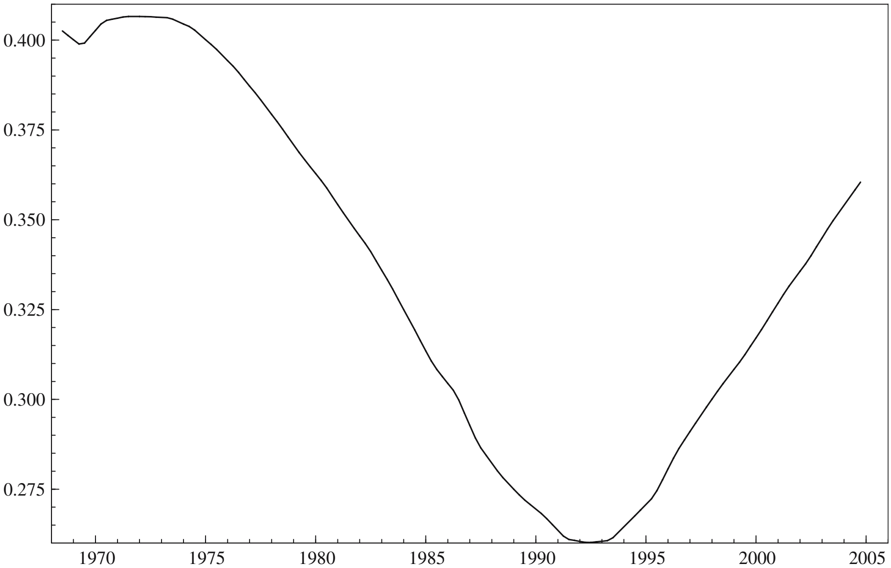

Fig. 1 The age structure variable AGE over the period 1968(3)–2004(4)

total private consumption expenditures and real disposable income in the household sector. The speed of adjustment was estimated to be quick, and hence on annual data a static relationship worked well. This changed however in the aftermath of the credit liberalisation at the beginning of the 1980s, when aggregate consumption rose sharply relative to income. Similar developments took place in, e.g., the UK and other Scandinavian countries as well (see Muellbauer and Murphy 1990; Berg 1994 and Lehmussaari 1990). The existing empirical consumption functions subsequently broke down. The respecification of the Norwegian consumption function by including a broad measure of household wealth succeeded, however, in accounting for the breakdown ex post (see Brodin and Nymoen 1992). In more detail, the results of Brodin and Nymoen (1992), based on quarterly data from 1968(1)–1989(4), can be summarised in four points:

1. Cointegration: The log-linear relationship between the three variables C (total private consumption), Y (real disposable income) and W (net household wealth in real terms)

$$\log C_t = Constant + \beta_1 \log Y_t + \beta_2 \log W_t, \tag{2}$$

   constituted a cointegrating relationship. In the equation, subscript t denotes the time period, while β_1 and β_2 denote the cointegration parameters for Y and W, respectively. The "Appendix" contains detailed definitions of the variables C, Y and W.
2. Weak exogeneity: Income and wealth were weakly exogenous to the cointegration parameters.
3. Invariance: Estimation of the marginal models for income and wealth showed evidence of structural breaks. The joint occurrence of a stable conditional model (the consumption function) and unstable marginal models for the conditioning

variables is evidence of within-sample invariance of the coefficients of the conditional model and hence super exogenous conditioning variables (income and wealth). The invariance is confirmed by Jansen and Teräsvirta (1996) who use an alternative method based on smooth transition models.

4. Speed of adjustment is relatively fast, both with respect to revaluation of wealth and changes in income.

Eitrheim et al. (2002) showed that these features of the Brodin and Nymoen (1992) model remain more or less unchanged when the consumption function is estimated on the extended sample 1968(3)–1998(4); consumption, income and wealth were seen to cointegrate; the latter two variables were found to be weakly exogenous with respect to the cointegrating parameters and the recursive estimates of the cointegration coefficients for income and wealth were stable over the sample period. Furthermore, the adjustment speed towards equilibrium was still relatively fast, although somewhat lower than that reported in Brodin and Nymoen (1992).

However, at the turn of the century, the performance of the consumption function of Eitrheim et al. (2002) deteriorates somewhat. To see this, consider Table 1 where we show the results of estimating Eq. (2) over the sample of Eitrheim et al. (2002), as well as the results for an extended sample ending in 2004(4). In both cases, the same estimation method as in Eitrheim et al. (2002) has been used.

The results for the 1968(3)–1998(4) sample comes very close to the results in their Table 6. Eitrheim et al. (2002) estimate the income elasticity to be 0.65 and the wealth elasticity to be 0.23 . The small changes shown in the table can be attributed to revisions of the historical data (national accounts revisions and adjustment of the income series, cf. Section 4 and the "Appendix"). However, the sample extension to 2004(4) affects the estimated coefficients markedly, and the standard errors of the elasticities are larger than for the Eitrheim et al. (2002) sample. One interpretation of this development is that Eq. (2) subsumes an age variable in the constant term, which has been undetected empirically because the samples used so far have been dominated by colinearity with the trend in income and wealth. As shown in Fig. 1, the age composition of the Norwegian population has behaved like a trend over large parts of the sample period. This can have hidden age composition effects in the consumption function over a long period of time, making the function stable as long as the age structure development have co-moved with either income or wealth (or with a linear combination of the two).

Another variable that may influence consumption in the long run is the real interest rate. The income effect of changes in interest rates is already included in the Brodin and Nymoen (1992) consumption function via the income variable.

Table 1 Long-run consumption functions estimated on two different sub-samples

Sample period	Estimated long-run consumption functions
1968(3)–1998(4)	$\log \hat{C}_t = \underset{(0.03)}{0.68} \log Y_t + \underset{(0.02)}{0.22} \log W_t$
1968(3)–2004(4)	$\log \hat{C}_t = \underset{(0.05)}{0.75} \log Y_t + \underset{(0.03)}{0.15} \log W_t$

Johansen (1995) estimation method. Standard errors are in parentheses.

However, there may also be substitution effects from interest rate changes; an increase in real interest rates makes consumption today more expensive relative to tomorrow's consumption; hence, consumption is expected to decline. From 1984(1), the after-tax real interest rate variable, RR_t, is calculated as

$$RR_t = RLB_t(1 - \tau_t) - \Delta_4 \log CPI_t, \tag{3}$$

where RLB_t =average nominal interest rate on bank loans, τ_t =marginal income tax rate for households and $\Delta_4 \log CPI_t$ =change in annual consumer price inflation. RR_t is set to zero before 1984, because of strict credit market restrictions which prevented the interest rate movements to have significant effects on saving. Norwegian credit markets were gradually de-regulated from 1984.

We next attempt to establish an extended consumption function, where the effects of changes in the age structure and real interest rates are accounted for. Because of the focus on demography, we choose to express the extended consumption function in per adult capita terms, where aggregate consumption, income and wealth variables are divided by the adult population, defined to be those 20 years or older. However, let it be said that the results obtained do not depend in any substantive way on the per capita formulation.

In the outset, we assume that both the age structure variable and real interest rates are weakly exogenous and test the hypothesis that the following relationship constitutes a cointegrating relationship:

$$\begin{aligned} &\log c_t = \beta_1 \log y_t + \beta_2 \log w_t - \beta_3 AGE_t - \beta_4 RR_t + Constant, \\ &\quad \beta_j > 0, \mathrm{j} = 1, 2, \text{ and } \beta_\mathrm{i} \geq 0, \mathrm{i} = 3, 4. \end{aligned} \tag{4}$$

Small letters denote the corresponding variable in per adult capita terms ($c_t = C_t/N_t$, etc., where N_t is the population of age 20 or older) in Eq. (4).

In the extended sample period, i.e. from 1999(1)–2004(4), temporary changes in the Norwegian tax system have led to large changes in dividends paid to private households from 1 year to another. The propensity to consume dividend payments is likely to be lower than other sources of disposable income, and hence we use a disposable income series which is adjusted for dividends. The disposable income series is adjusted for dividends from 1978(1) onwards. We would have ideally adjusted the disposable income series back to 1968(3) but the lack of data makes this difficult to do. However, as dividend payments are low in the end of the 1970s, we do not believe this to influence the results. In fact, this measure was already used in Table 1, and we continue to use the adjusted income series as our primary income measure in this paper. All main results are, however, robust to the choice of income measure.[2]

[2] Interested readers can also consult the working paper version of Erlandsen and Nymoen (2004), where the unadjusted disposable income series is used throughout (see http://www.oekonomi.uio.no/memo/memopdf/memo2704.pdf).

4 Cointegration and exogeneity

The underlying modelling assumption used by the previous studies is that (log of) C_t, Y_t and W_t are integrated of degree one, $I(1)$ in a common notation. In this subsection, we test the hypothesis that the extended relationship in Eq. (4) is a cointegrating relationship. Unless the cointegration claimed by the earlier studies is spurious, finding Eq. (4) to be cointegrating then logically requires that AGE_t and RR_t are integrated of degree zero, $I(0)$. Another interpretation, which can also be reconciled with the existing evidence, is that the two variables are $I(1)$ but cointegrated. We build on the first interpretation for the following reasons. First, being a rate, AGE_t has finite variance by construction; although, from Fig. 1, the characteristic root(s) of AGE will be close to unity empirically, at least on samples ending before 1990. Second, the $I(0)$ interpretation does not rule out non-stationarity in the form of deterministic shifts in mean, which is particularly relevant for the real interest rate. The nominal interest rate for example has been affected by the following regime changes: first, as noted above, administered rates until 1984, then interest rate determination in the domestic money market and in the few periods of currency crises in the market for foreign exchange, and since 2001 as an instrument in an inflation targeting regime. Other shocks to the real interest rates come from currency devaluation in 1986 and lower-than-anticipated inflation in 2003 and 2004.

The null hypothesis of a unit-root in AGE_t is formally rejected by a fifth-order augmented Dickey–Fuller (ADF) test. For the sample 1968(3)–2004(4), the Dickey–Fuller statistic is -3.64, which is significant at the 1% level (constant term is included). In the case of RR_t, it is not straightforward to establish the statistical adequacy of the model underlying the ADF, which is important (see Andreou and Spanos 2003). This is mostly due to the several regime changes noted above. When the impact of discernible regime shifts are taken care of, statistical adequacy is achieved and there is support for rejection of the unit root hypothesis also for RR_t.[3]

To test for cointegration, we apply the Johansen methodology (Johansen 1988, 1995). In line with earlier studies of Brodin and Nymoen (1992) and Eitrheim et al. (2002), a vector autoregressive (VAR) model of the log of the three variables, c_t, y_t and w_t, with five lags is used to obtain Gaussian residuals. To take account of the two new variables introduced above, we condition on AGE_{t-1} and RR_{t-1}, which are regarded as weakly exogenous variables in the system. For the purposes of testing for the rank, the two variables are taken as $I(1)$. In accordance with the suggestions of Harbo et al. (1998) for partial systems, a deterministic trend is also included in the system. In addition, a constant term, centred seasonal dummies ($CS1$, $CS2$ and $CS3$), and the two dummy variables VAT and $STOP$, which capture the effects of the introduction of the VAT in 1970 and of a wage and price freeze in 1978–1979, respectively, enter the system unrestricted.

[3] The sample is 1984(2)–2004(4), i.e. the period after credit marked liberalization. Dummies are used to mop up the impact of devaluation in 1986 and of short periods with unusually high or low interest rates in 1997, 1998, 2003 and 2004. The value of the Dickey–Fuller, though it can only serve as a guideline, is −2.94. The conventional 5% critical value is −2.9. Moreover, the recursive plot of the Dickey–Fuller statistic shows that it is decreasing steadily over the whole sample period.

Table 2 Johansen tests of cointegration

Eigenvalues	Hypotheses (of rank) and trace test			
λ_i	H_0	H_1	λ_{trace}	5% critical value[1]
0.17	$r = 0$	$r \geq 1$	61.03	57.32
0.14	$r \leq 1$	$r \geq 2$	33.29	35.96
0.08	$r \leq 2$	$r \geq 3$	11.63	18.16

VAR system of order, 5; range, 1968(3)–2004(4); endogenous variables, (log of) $c\ y\ w$; exogenous variables, $AGE\ RR\ \ Trend$; deterministic variables, $VAT\ \ STOP\ \ Const\ \ CS1\ \ CS2\ \ CS3$.

[1] The critical values are taken from Table 13 in Doornik (2003), allowing for two exogenous variables in the system.

Trace test statistics for the sample period 1968(3)–2004(4) are reported in Table 2. Although the critical values take into account that the system is partial, the correct distribution of the test statistics is unknown as it will also depend on the dummy variables which we have included in the system. That said, it is encouraging that the trace test statistics λ_{trace} in the table, for the null hypothesis of no cointegration $(r = 0)$, is above the 5% critical value. The statistic testing the null hypothesis $r \leq 1$ is below the relevant 5% critical value. Hence, the formal tests in Table 2 support at most one cointegrating vector.

As argued by Hendry and Juselius (2001), the decision about cointegration rank should be guided by a broader set of considerations besides formal statistical tests. For one thing, the interpretability of the cointegration vectors, also relative to pre-existing evidence, is of importance in econometrics. As we show below, choosing $r = 1$ gives a model with good properties in that respect. Moreover, recursive graphs of the trace statistic provides a useful additional informal test. If at least one eigenvalue is non-zero, the trace statistics should have a tendency to trend upward as the sample is extended. Along the same line, according to Hansen and Johansen (1999), the sup eigenvalue can be used as a formal test of cointegration.

Figure 2 shows that the recursive trace test statistics for the extended model are markedly more stable than for the Eitrheim et al. (2002) model, i.e., without AGE and RR . The sequence of trace statistics for the Eitrheim et al. (2002) relationship falls visibly around the time when the change in correlation between y, w and AGE and RR is beginning to make its mark. Though informal, the sequence of trace test statistics in Fig. 2 therefore lends support to the $r = 1$ decision.

The deterministic trend is insignificant in the $I(0)$ system, and it is therefore excluded from the VAR model in the following analysis. Conditioning on $r = 1$, we now turn to test for weak exogeneity of y_t and w_t in the system. The χ^2 -distributed statistic for testing the restrictions that the equilibrium correction coefficients of income and wealth are jointly zero has a p-value of 0.62. Hence, the joint hypothesis is acceptable statistically speaking, and y_t and w_t are thus considered as weakly exogenous for the cointegrating parameter. Equation (4) can therefore be regarded as a long run consumption function. The estimated cointegration parameters of the system are given in Eq. (5), with standard errors in parentheses.

$$\log c_t = \underset{(0.03)}{0.66} \log y_t + \underset{(0.02)}{0.17} \log w_t - \underset{(0.08)}{0.31} AGE_t - \underset{(0.19)}{0.42} RR_t + Constant. \tag{5}$$

A comparison with Table 1 shows that the income and wealth elasticities in Eq. (5) are very close to those of the Eitrheim et al. (2002) model, on their sample. The two new coefficients in the consumption function are both significant at a 1% level and have, as expected, negative signs. The model predicts that a unit increase in AGE reduces consumption by 0.31%. If income is unaffected, the savings rate is predicted to increase from 5 to 5.3%. As we have seen, AGE increased by several units in the course of only a few years in the 1990s. According to our results, this demographic change has significantly affected the savings rates. Consumption is similarly expected to decrease by 0.42% when after-tax real interest rates increases with one percentage point, ceteris paribus.

The Engle–Granger method support cointegration: the Dickey–Fuller statistic (based on a ADF of order 4) is −10.2. The estimated cointegrating relationship is:

$$\log c_t = 0.57 \log y_t + 0.20 \log w_t - 0.34 AGE_t - 0.33 RR_t + Constant. \qquad (6)$$

Among the two new explanatory variables, it is the age structure variable which is most robust. Hence, it is necessary to include AGE in the model to obtain significant real interest rates effects, while AGE is still significant on a 10% level when RR is left out of the information set. However, both the impact and the significance of the age structure variable also increase when changes in real interest rates are controlled for. One interpretation of this interdependence is that the effects on aggregate consumption of changes in these two variables have counteracted each other. A glance at the development in AGE and RR over the sample period supports this view. With the exception of the last few years, AGE and RR have moved in opposite directions. Hence, as both variables enter the extended consumption function with negative signs, their separate effects on consumption may have been hidden. At the end of the 1990s, on the other hand, both variables

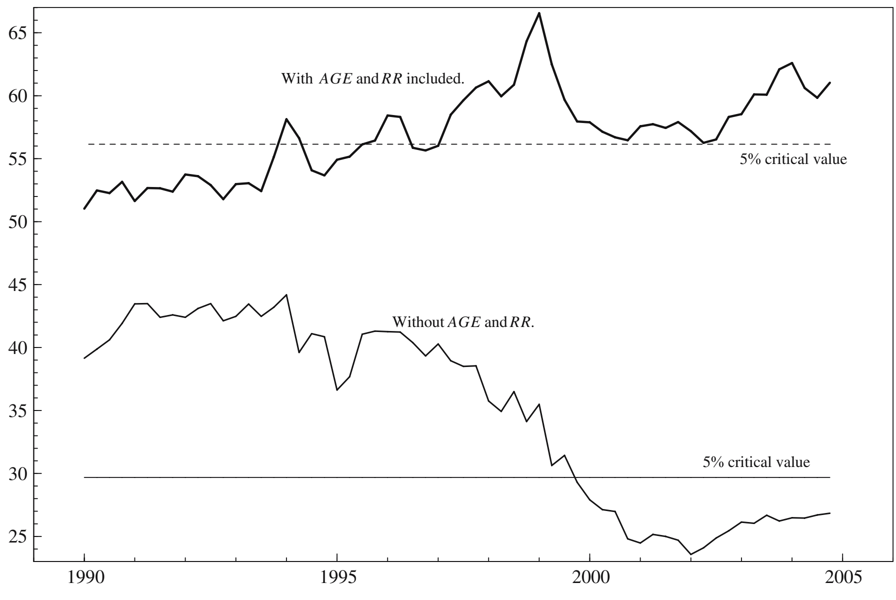

Fig. 2 Sequences of trace test statistics, with and without AGE and RR in the information set

moved in the same direction. It is when this period is included in the sample period that the evidence of cointegration between consumption, income and wealth become weaker.

The first row of Fig. 3 plots the recursive estimates of the income and wealth elasticities (β in the graphs) of Eq. (5) with ± 2 estimated standard errors (σ in the graphs). The plots show some instability in the period 1980–1989 (not significant though), but there is great stability over the last 15 years of data. The second row shows that the recursive plots of the two new coefficients in the consumption function are also relatively stable, being inside $\pm$ two standard errors for most of the period.

Having identified a revised cointegrating relationship, we next turn to the specification of an equilibrium correction model for $\Delta \log c_t$, conditional on the exogeneity of w_t and y_t, and with an equilibrium correction term as given by Eq. (5). Apart from those restrictions, the general single equation model which makes up the starting point of a general-to-specific search procedure contains the same lags and dummies that we used in the cointegration analysis. The preferred model is shown in Table 3. The equilibrium correction term, $EqCM_t$, is defined in accordance with Eq. (5) and is reproduced below the estimated equation in the table. Despite the new equilibrium correction term, 6 years of new data and data revisions, the equilibrium correction model is similar to the preferred equation in Eitrheim et al. (2002).

The bottom part of the table contains the following diagnostic tests; the multiple correlation coefficient (R^2), the residual standard error ($\hat{\sigma}$), the F_{Null} test of the null hypothesis of 'no relationship', the F distributed tests of residual autocorrelation ($F_{AR(1-5)}$) and autoregressive conditional heteroscedasticity ($F_{ARCH(1-4)}$). In addi-

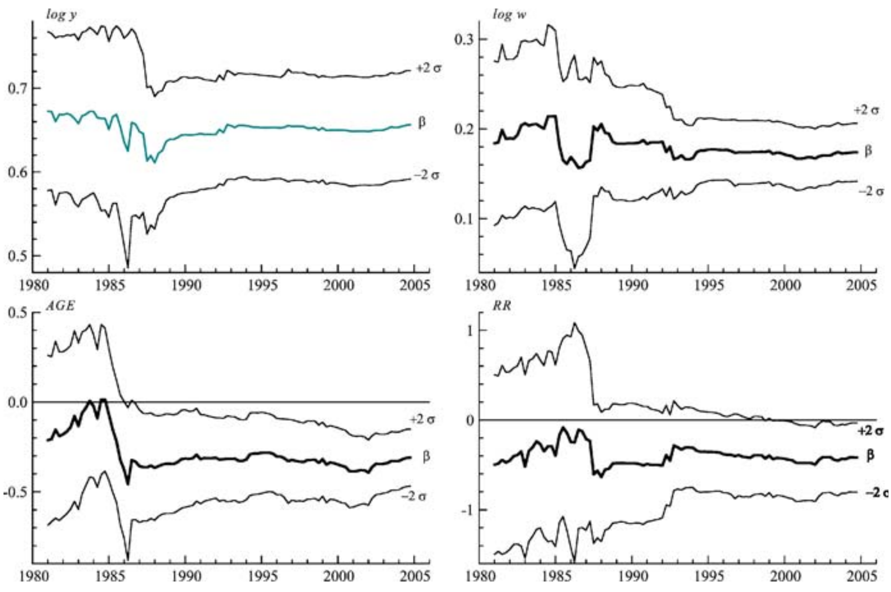

Fig. 3 Recursive estimates of the coefficients of the extended cointegrated consumption function (denoted β), with ± 2 standard errors (denoted σ). Full sample estimates correspond to results in Eq. (5)

Table 3 Equilibrium correction model of consumption

$$
\begin{aligned}
\Delta \log c_t = &- \underset{(0.06)}{0.14}\ \Delta \log c_{t-1} + \underset{(0.05)}{0.26}\ \Delta \log c_{t-4} + \underset{(0.03)}{0.18} \\
&+ \underset{(0.04)}{0.20}\ \Delta \log y_t + \underset{(0.04)}{0.21}\ \Delta \log w_{t-1} - \underset{(0.07)}{0.47}\ EqCM_{t-1} \\
&+ \underset{(0.07)}{0.18}\ STOP_t + \underset{(0.01)}{0.08}\ VAT_t \\
&- \underset{(0.01)}{0.07}\ CS1_t - \underset{(0.01)}{0.04}\ CS2_t - \underset{(0.01)}{0.03}\ CS3_t
\end{aligned}
$$

$$EqCM_t = \log c_t - 0.66 \log y_t - 0.17 \log w_t + 0.31 AGE_t + 0.42 RR_t$$

The sample is 1968(3) to 2004(4), 130 observations. Estimation is by OLS.

Standard errors are in parentheses below the parameter estimates.

$\hat{\sigma} = 1.43\%$	$R^2 = 0.96$	$F_{Null} = 322.6[0.00]$
$F_{AR}(1-5) = 0.23[0.95]$	$F_{ARCH(1-4)} = 1.30[0.27]$	$\chi^2_{normality} = 0.07[0.97]$

tion, we report the Doornik and Hansen (unpublished paper) chi-square test of residual non-normality ($\chi^2_{normality}$) (see Doornik and Hendry 1999). The numbers in brackets are p-values for the respective null hypotheses, implying that none of the diagnostic tests are significant.

Figure 4 shows the stability of the model over the period 1980(1)–2004(4). The first six graphs show the recursively estimated elasticities in the same order as in Table 3. The last three graphs show first the one-step residuals with ± 2 residual standard errors, $\pm 2se$ in the graph, the sequence of one-step Chow statistics scaled with their 1% critical levels and, finally, the recursive breakpoint Chow tests (also

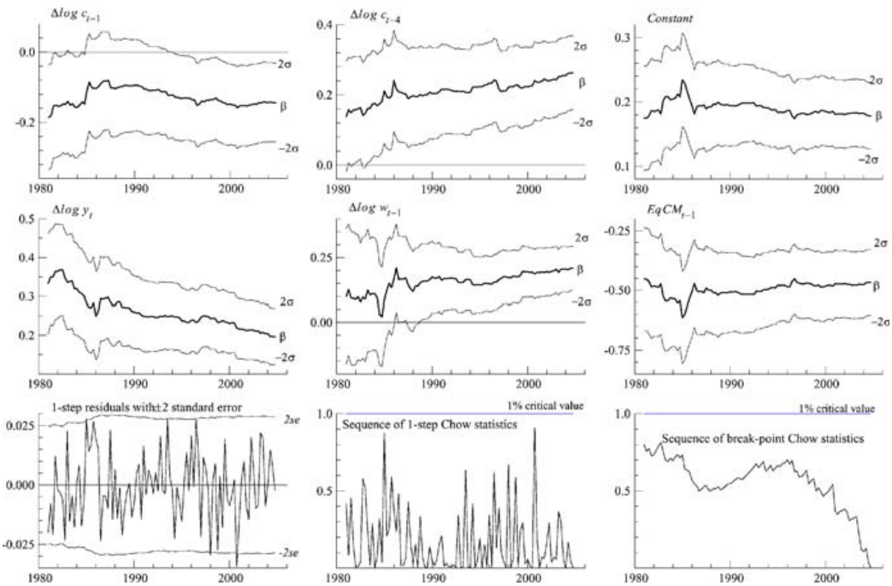

Fig. 4 Recursive OLS estimates of the equilibrium correction model of $\Delta \log c_t$

with the one-off 1% level indicated). All graphs show a high degree of stability. Note in particular the stability of both the equilibrium correction coefficient and the constant, implying that the mean of the long run cointegrating relationship is stable over the whole sample. Hence, the instability of the model of Eitrheim et al. (2002) can be attributed to the omission of AGE and RR, consistent with our hypothesis.

With one exception, all the results of this section are robust with respect to the choice of income measure: total disposable income or the adjusted series. The one exception is that, with unadjusted income, there is some evidence of endogenity of income with respect to the parameters of the cointegration relationship. However, when the right estimation methods are used both in the estimation of the long run consumption function and when estimating the equilibrium correction model, the results for the coefficient are, virtually speaking, unchanged.

5 Concluding remarks

In this paper, we have investigated the empirical relationship between aggregate consumption and the age structure of the population in Norway. The analysis is based on aggregate time series data, and age structure changes are represented by the ratio of number of middle-aged persons, defined as those between 50 and 66 years of age, to the rest of the adult population. Similar to what other studies using aggregate data report, we found that changes in the age structure of the population have significant and numerically important effects on Norwegian aggregate consumption. We more specifically found that aggregate consumption decreases when the share of the middle-aged persons in the population increases. Our results, hence, give support to the life cycle model. An important finding in the paper is that controlling for age structure effects stabilises the parameters of the consumption model when data up to year 2004 are included in the estimation sample. Although several studies have reported significant age structure effects on aggregate consumption in different countries, none has, to our knowledge, emphasised that a changing age distribution can play such a key role in the consumption function. Furthermore, controlling for age structure effects reveals significant real interest rate effects on aggregate consumption.

Our findings imply that the changes in the age distribution of the Norwegian population may have significantly affected domestic demand and households' savings. Our operational measure of age structure has increased by around ten units since the start of the 1990s. The predicted effect on the savings rate of such a change is in the range of $1.5-5$ percentage points. However, as it is quite possible that both income and wealth may have been directly influenced by such a marked shift in the age structure, at least through several years of adjustment, the dynamics of the savings rate may be more complicated than suggested by the results reported in this paper. More insight into the dynamics of the savings rate will require larger structural models though, which is an interesting area of future research. An issue not discussed in this paper is the role of demography in forecasting. In contrast to most of the economic explanatory variables, changes in the age structure of the population are known with approximate certainty in the short to medium run. Whether this property of the age structure variable or demography in general can be used to improve forecasts of consumption and gross domestic product growth also represents interesting topics for future research.

Although the analysis in this paper has been restricted to Norwegian data, there are good reasons to believe that the results also apply to other countries. Identifying age structure effects on consumption may be of increasing importance in many countries. Several developed countries experienced a baby boom in the first decades after World War II. The share of middle-aged prime savers relative to the rest of the population is, hence, also on the rise in these countries. If our findings on Norwegian data apply to these countries, the changing age distribution would then put a downward pressure on consumption in the Western world in the coming years; a pattern the empirical models then would not identify if age structure effects are not controlled for.

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Appendix

Data definitions

The data are taken from Norges Bank’s model database RIMINI (the spring 2005 version of the database) and Statistics Norway’s KVARTS database. In addition to variable descriptions, the table below contains the source for each of the variables.

Symbol and definitions and sources

AGE	$\frac{\text{Population } 50-66 \text{ years old}}{\text{Population } 20-49 \text{ years old} + \text{Population above } 66 \text{ years old}}$ (Statistics Norway).
C	Private consumption expenditure (incl. ideal organisations), fixed 2002 prices (Statistics Norway).
CPI	Consumer price index (2002=1) (Statistics Norway).
N	The population with age 20 or older (Statistics Norway).
PC	Price deflator for total private consumption expenditure (2002=1) (Statistics Norway).
RR	Marginal after-tax real interest rates for households. 1968(3)–1983(4), zero. 1984(1)–2000(4), $RLB(1-\tau)-\Delta_4 \log CPI_t$, where RLB =average interest rate on households’ bank loans and τ = marginal income tax rate for households (Statistics Norway and Norges Bank).
STOP	Income policy dummy; constructed for catching up the inflationary pressures which were being built up during the wage and price freeze in 1978. It takes non-zero values in the quarters from 1979(1) to 1980(1) and is zero elsewhere. See Brodin and Nymoen (1992) for details.
VAT	Dummy for the introduction of VAT. Takes the values 1 in 1969(4) and −1 in 1970(1).

W Real household wealth; nominal household wealth (financial and housing wealth) deflated by PC (Statistics Norway and Norges Bank).
Y Households' real disposable income; nominal disposable income deflated by PC (Statistics Norway).

As explained in the main text, the lower case letters c, y and w are used to denote per adult capita values of the corresponding variables, $c = C/N$, $y = Y/N$ and $w = W/N$.

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