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Wealth effects on consumption in financial crises: the case of Norway

Abstract:
A dynamic consumption function, where consumption in the long run is determined by households' disposable income and wealth, has been superior to the Euler equation in explaining the development of Norwegian aggregate consumption over several decades. This period covers the years of financial deregulation in the mid 1980s, the banking crisis around 1990 following the deregulation and the current international financial crisis. In the current version, long run consumption is homogeneous in income and wealth and there is also a significant effect from after-tax real interest rates. A change in the correlation pattern between real interest rates and wealth, which is related to a change in the monetary policy regime, is the reason why both variables need to be included in the long run relationship in order to explain the development over the past four years.

Keywords: financial crisis, consumption, wealth effects, interest rates, savings rate.

JEL classification: C51, C52, C53, E21

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1. Introduction

The present international financial crisis has given rise to considerable changes in the value of financial assets as well as in real estate prices in all developed economies. This has revitalized the debate about how – and by how much – revaluations of financial and tangible wealth affect the real economy. In particular, it is of interest to examine how the real value of household wealth affects aggregate consumer demand. The magnitude of this effect may shed light on the extent that changes in wealth lead to financial consolidation, i.e. increased household savings. The experience from the Norwegian banking crisis in 1990-92 is that a drastic fall in household wealth can lead to a significant rise in the savings rate.

The consumption function has been debated by economists for more than 70 years. The idea of such an empirical relationship goes back to Keynes (1936), who held that increased household income would cause consumption to rise. The work of Keynes was published at a time of mass unemployment which the reigning economic paradigm, classical economics, could not account for. The degree of acceptance of the concept reached a high in the 1950s and 1960s, as Keynesian models and policies then dominated.

There is more than a formal difference between Keynes and the classics: Clower (1965) showed that the Keynesian consumption function cannot arise in a Walrasian general equilibrium model. The reason for this is that incomes are defined in terms of quantities as well as prices, and quantity variables never appear explicitly in the market excess demand functions of Walras. It follows that an agent's spending decision is not conditioned on income, see Hoover (2009, p.22) for details. In the classical model world the consumers rule the day. They own the production sector and their supply of labor are fully employed at real wages provided by the market's invisible hand. Hence, the households get the economic activity level (as well as the consumption and the incomes) they want according to their own preferences. As Trygve Haavelmo has pointed out there is no lack of logical consistency in the classical model. What is lacking is relevance,

> *"..because it gave no realistic description neither of how production decisions take place in a modern capitalistic economy nor of how actual investment and saving decisions are made. The problem is that it is not the consumers who decide on production – if that were the case, there would no doubt always be full employment as most people would want high consumption."* (Haavelmo in Andvig (1979), my translation.)

Mainstream modern macroeconomics – with its emphasis on microfoundations that entail rational intertemporal-optimizing consumers – has inherited many of these stylized properties from classical economics. For example DSGE models, which have been the dominating modeling tool in academia and in central banks for a decade[1], build on a set of Euler equations which implies that the economy is always near equilibrium. Moreover, if exposed to shocks, such a model economy rapidly returns to equilibrium. The recent financial crisis – with falling production and rising unemployment across the world – shows that this is not a good model of the actual economy out there.

Also, the DSGE models play down the role of fiscal policy. Still, the models have important implications for economic policy. In the context of inflation targeting, the demand effects –via real income and wealth – of lowering the interest rate in order to bringing a low inflation up towards the target, are absent in models based on Euler equations. An assessment of the recent massive expansive policies in industrial economies – which have entailed fiscal as well as monetary stimuli – cannot be done within DSGE-type models, nor can they be rationalized within such models. This and many other shortcomings of DSGE models are examined by Muellbauer (2009), who also remarks that:

> *One of the eminent proponents of DSGE models with New Keynesian frictions, Jordi Galí admitted at the European Area Business Cycle Network conference in November 2008 that the New Keynesian DSGE models had little to say about the current crisis.*
> (Muellbauer 2009, p. 2)[2]

Since private consumption constitutes about half of the domestic absorption in modern economies, it is imperative for policy makers to have a firm grip on the driving forces behind aggregate consumption. Accordingly, to decide on policy actions to counteract the current crisis and to make sensible forecasts for the short and medium run, it is helpful to have access to a well-specified empirical macroeconometric model, see Bårdsen and Nymoen (2009) and Bårdsen *et al.* (2005). Such a model should capture the main driving forces behind the development of both the supply side and the main demand components in the economy.

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[1] There is a plethora of such models entertained by central banks: Federal Reserve Board (Edge *et al.* (2007), the European Central Bank (Smets and Wouters (2003) and Christoffel *et al.* (2008), Bank of Canada (Murchison and Rennison (2006)), Bank of England (Harrison *et al.* (2006), Bank of Finland (Kortelainen (2002) and Kilponen and Ripatti (2006), the Bank of Norway (Brubakk *et al.* 2006), Bank of Sweden (Adolfsson *et al.* (2007)), and the Bank of Spain (Andrés *et al.* (2006)).
[2] Interestingly, the first line of the abstract in Galí et al. (2007) – over 70 years after Keynes – states that "recent evidence suggests that consumption rises in response to an increase in government spending", before the authors carry on to extend the standard new Keynesian model to allow for the presence of rule-of-thumb consumers.

In this paper we first demonstrate the empirical relevance of a long run relationship between consumption, income and wealth, including the issue of stability across periods of financial deregulation, and we discuss its relevance for the aggregate consumption across countries in section 2. For Norway a conditional consumption function with household real income and wealth as long run determinants has – subject to minor updates and revisions – been able to explain the development in aggregate consumption over a period of 20 years.

In section 3 we estimate a version of this model. We compare it with two Euler equations in sections 4 and in the following section we compare it with an alternative equilibrium correcting consumption function without any wealth effects in the long run part of the model. We find that the baseline consumption function forecast encompasses the one without wealth effects as well as the Euler equations. In section 5 we also demonstrate how a changing pattern of correlation between the real after-tax interest rate and household wealth necessitates the inclusion of both in order to obtain a well-specified conditional macro consumption function. The importance of taking account of wealth effects in the long run consumption relationship is further analyzed in section 6 where we look at consequences for the savings rate. When wealth are allowed to play a role, it is shown that a fall in equity as well as in housing prices – as we saw after the collapse of Lehman Brothers in September 2008 – will lead to a financial consolidation in the households which manifests itself as a rise in the savings rate. Section 7 concludes.

2. A long run relationship between consumption, income and wealth

Several empirical investigations have come up with support for a long run relationship between consumption (C), income (Y), and wealth (W) based on aggregated national accounts data from different countries. By means of cointegration analysis of a well specified VAR (see Johansen, 1988), it is commonly found that the three variables are I(1) and that there exists one and only one cointegrating relationship between them on logarithmic form, henceforth c, y, and w respectively. Then it follows from Granger's Representation Theorem that there exists an equilibrium correction relationship between them, see Engle and Granger (1987). It does, however, not follow which of the variables that equilibrium corrects.

Lettau and Ludvigson (2001, 2004) find on US data that it is only wealth that plays this part. They interpret this finding in the light of a theoretical model where total wealth is split into two components, tangible and financial wealth on the one hand and human capital on the other. The latter is unobservable. As an empirical approximation human capital is set equal to the present value of all future income. This again is assumed proportional to the income level today. Hence, income can be interpreted as a proxy for human capital and the ratio between consumption and total wealth can be expressed as a log-linear relationship ($c - \beta_y y - \beta_w w$) where homogeneity is imposed, that is $\beta_y + \beta_w = 1$. Moreover, Lettau and Ludvigson (2001, 2004) call upon microfoundations by referring to Campbell and Mankiw (1989) who show that the log of the ratio between consumption and total wealth for a representative agent becomes the sum of rational predictions of the difference between the future return on total wealth and the future consumption growth. Lettau and Ludvigson (2001, 2004) argue that this expression is stationary, which amounts to cointegration between c, y and w. Rudd and Whelan (2006) have, however, based on a re-examination of the data cast doubt about the support for cointegration on statistical grounds.[3]

Similarly, Hamburg *et al.* (2008), using data from Germany for the period 1980(1)-2003(4), find that it is income, and only income, that equilibrium corrects to the long run equation. In other words, when consumption, income and wealth deviate from there common trend growth path, it is income that adjusts to re-establish the equilibrium. These results are consistent with the literature based on the Euler equation, which seeks to aggregate the optimal intertemporal consumption decision of a representative consumer characterised by rational expectations, cf. Hall (1978) and section 4 below.

Barrell and Davies (2007) have argued in favour of the conditional consumption function, *i.e.* that it is consumption that equilibrium corrects. They follow the tradition of empirical work based on broad implications of the life cycle model, whereby planned consumption is a function of both human and non-human wealth, see Deaton (1992).[4] They assess the impact of financial liberalisation on consumption for seven OECD countries – the United States, the UK, Germany, France, Japan, Canada and Sweden – utilising a dynamic equilibrium correction model with both tangible and financial wealth. At the same time they allow liberalisation to impact differentially on the determinants of consumption in the short and long run. For all countries, except Germany and partly Japan, estimates of the significance of leveraged dummies for liberalisation are found to be consistent with the prior

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[3] Rudd and Whelan (2006) also claim that the consumption variable adopted by Lettau and Ludvigson (2004) – *i.e.* consumption of non-durables excluding shoes and clothing – is inappropriate because the budget constraint relates to total consumption, and the excluded parts constitute an increasing share of real total consumption over time (cf. Figure 1 in Rudd and Whelan (2006)).

[4] The authors refer to Ludvigson and Steidel (1999) who also study wealth effects in a quarterly log-linear long run US consumption relationship. Ludvigson and Steidel (1999) find a common trend and also statistically significant wealth and income effects.

view that removal of liquidity constraints reduces the response of consumption to real disposable income and boosts the wealth effects.

These issues are also the focus of the empirical work based on UK and United States data by John Muellbauer and his collaborators,[5] who are using a different functional form derived in Muellbauer and Lattimore (1995). The long run is derived as an approximation to the rational expectations permanent income hypothesis with habits in which wealth enters as a levels ratio to income in an otherwise linear relationship in the logs of consumption and income. The model is augmented with a credit conditions index (CCI) which is allowed to interact with the income and wealth terms both in the short and long run. In principle this enables the investigator to identify the effect of credit liberalisation from the income and wealth effects *per se*. The CCI for the UK in particular seems to adequately represent the effects of financial liberalisation in the 1980s, see Fernandez-Curogedo and Muellbauer (2006). The estimated consumption model with CCI encompasses the standard model without effects of financial liberalisation. The results of Muellbauer and his co-authors show that omitting the effects of credit conditions exaggerates the wealth effect and leads to underestimation of the income effect.

For Norway Brodin and Nymoen (1992) were the first to provide a model for total consumption in which cointegration analysis established the log-linear long-run relationship,

$$c = \text{constant} + 0.56\,y + 0.23\,w, \tag{1}$$

where log of total consumption (c), household real disposable income (y) and household real wealth (w) are non-stationary and integrated variables. Moreover, tests allow income and wealth to be regarded as weakly exogenous for the cointegration parameters, see Johansen (1992), and hence it is consumption that equilibrium corrects. Finally, estimation of marginal models for income and wealth showed evidence of structural breaks. The joint occurrence of a stable conditional model and unstable marginal models for the conditioning variables is evidence of within sample invariance for the coefficients of the conditional model and hence super exogenous conditioning variables (*i.e.* income and wealth), cf. Engle *et al.* (1983).[6] This finding entails that the Lucas critique – for example due to changing income expectations in Hall's model – has low power empirically in Brodin and Nymoen (1992).

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[5] See Muellbauer (2008) for a survey and Aron *et al.* (2008, 2009) for the UK and for the United States, respectively.

[6] The results of invariance have been corroborated by Jansen and Teräsvirta (1996) using an alternative method based on smooth transition models.

Brodin and Nymoen (1992) estimated their model on quarterly data 1968(3) – 1989(4), see Table 1. It is worth noting that frequent re-estimation of the model, extending the data gradually over a period of 15 years, never revealed any breakdown in the long run relationship.

Table 1. Survey of estimated long run coefficients in the Norwegian consumption function*

Reference	Sample	Income elasticity	Wealth elasticity	Semi-elasticity of after tax real interest rate	Semi-elasticity of age variable	Speed of adjustment	Residual st.error (per cent)
Brodin and Nymoen (1992)	1968(3)-1989(4)	0.56 (0.03)	0.27 (0.02)			-0.71 (0.08)	1.33
Eitrheim *et al.* (2002)	1968(3)-1998(4)	0.65 (0.17)	0.23 (0.07)			-0.34 (0.08)	1.55
Erlandsen and Nymoen (2008)	1968(3)-2004(4)	0.66 (0.03)	0.17 (0.02)	-0.42 (0.19)	-0.31 (0.08)	-0.47 (0.07)	1.43

* standard errors in parenthesis.

A more complete re-analysis of the long run relationship in the consumption function has been carried out in Eitrheim *et. al.* (2002) and in Erlandsen and Nymoen (2008). Eitrheim *et. al.* (2002) reconfirm the main results of Brodin and Nymoen (1992) on a sample from 1968(3) to 1998(4), that is with 36 additional quarterly observations. As is seen from Table 1, the income elasticity increases from 0.56 to 0.65, whereas the elasticity for wealth is reduced from 0.27 to 0.23. It is also shown that FIML estimation of a simultaneous structural system – with private consumption, income, wealth and housing prices as endogenous variables – yields results for the consumption function (on equilibrium correction form) which are very close to the single equation OLS estimates. This suggests that simultaneity bias is not an issue in the case of the Norwegian consumption function.

Moreover, Brodin and Nymoen's consumption function is also found to be stable across the financial deregulation – in contrast to the findings of the UK studies cited above. Up to the mid-1980s it was the consensus view that the aggregate consumption in Norway could be well explained by real disposable income both in the short and long run. The advent of financial deregulation changed all that. The pre-existing consumption functions failed to forecast and failed to explain *ex post* the consumption boom that followed in 1985-87 and the subsequent trough in the years after the boom leading up to the crisis in the Norwegian banking sector (1990-92).

The *ex post* forecast failures of models based on a constant equilibrium ratio between consumption and income in the long run offered an opportunity to test a conditional consumption function (CF) against an Euler equation (EE) on pre-deregulation data, as was done in Eitrheim *et al.* (2002). They found that while the conditional consumption function encompasses the Euler equation on a sample from 1968(2) to 1984(4), both models fail to forecast the annual consumption growth in the next years. In the paper the theoretical properties of the forecasts are derived for both models. Assuming that the EE is the true model and that the CF is a mis-specified model, it is shown that both sets of forecasts are immune to a break - that is a shift in the equilibrium savings rate – that occurs after the forecast has been made. Moreover, a failure in “before break” CF-forecasts is only logically possible if the consumption function is the true model within sample. Hence the observed forecast failure of the CF is corroborating evidence in favour of the conditional consumption function for the period before the break occurred. However, a re-specified consumption function that introduced wealth as a new variable in the long run was successful in accounting for the breakdown *ex* post.

When Eitrheim *et al.* (2002) carry out the cointegration analysis of the VAR recursively, the statistical support for one and only one cointegration vector between consumption, income and wealth appears to become gradually weaker over time. Erlandsen and Nymoen (2008) shed light on this issue using a data set which is extended up to and including 2004(4). They show that the statistical support for the long run relationship is re-established if they allow for the real after-tax interest rate and an age composition variable to be included in the cointegration space. There is a large literature internationally showing that the age composition of the population in a country can be of importance for the aggregate consumption[7]. Likewise, a negative interest rate effect is plausible as a substitution effect since increased real interest rate makes consumption today more expensive relative to consumption tomorrow.

A restriction of homogeneity in income and wealth in the long run is rejected in all of the Norwegian studies referred to above. The lack of homogeneity rules out a steady state growth path where consumption, income and wealth grow proportionally. A possible explanation might be that the accumulation equation, $\Delta W_t = Y_t - C_t$, does not apply in the data because of revaluations of both tangible and financial wealth.[8]

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[7] See *e.g.* Fair and Dominguez (1991), Higgins (1998), and Horioka (1997), who – much in line with the predictions of the life-cycle hypothesis (Modigliani and Brumberg, 1954) – find that savings decrease and aggregate consumption rises when the share of elderly people increases in a country’s population.

[8] In the case of Brodin and Nymoen (1992) it may also be due to a narrower wealth concept as only liquid financial assets were included in the net financial wealth measure.

3. A consumption function for Norway with long run wealth effects

In this section we report results from a re-analysis of the consumption function in Erlandsen and Nymoen (2008) − Model A in the sequel − based on extended data. In doing so we have followed procedures similar to those adopted in that study.[9] The key consumption concept to be explained is total consumption exclusive of health services and services from housing.[10] We have carried out a Johansen-analysis of a VAR consisting of the consumption variable (c), real disposable after-tax income, exclusive of equity income, (y) and real net wealth (w), all transformed into logs. In addition we condition the analysis on two exogenous variables, the age composition variable AGE and real after-tax interest rate, RR,[11] both of which may enter the cointegrating space. Tests show that c and y are I(1), w is I(1) with a deterministic trend and AGE is I(0). RR is also deemed to be I(0) when we allow for deterministic shifts in the mean. [12]

A VAR with 5 lags in c, y and w, and one lag in AGE and RR, yields a well specified model according to the diagnostic tests reported in Table 2. We have followed the recommendations in Harbo *et al.* (1998) and included a deterministic trend which is restricted to enter the cointegration space. In addition the model contains restricted deterministic terms: a constant, seasonal dummies (CS1, CS2, CS3) to take account of seasonal adjustment and a dummy variable CPSTOP which picks up the effect of a wage price freeze in 1978-79. Table 2 reports trace-tests for the rank of the system. The tests support the hypothesis of only one cointegrating vector in the system.

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[9] The estimation is carried out with the programs in Oxmetrics 5, see Doornik and Hendry (2006a,b) and Oxmetrics 6, see Doornik and Hendry (2009) .

[10] The previous studies − Brodin and Nymoen (1992), Eitrheim *et al.* (2002) and Erlandsen and Nymoen (2008) − analyzed total consumption. Our choice corresponds to the endogenous component of consumption in the Statistics Norway quarterly forecast model KVARTS, see Eika and Moum (2005) for a description. Private consumption of health services is exogenous to the model because it is largely determined by public policy decisions as almost all private health expenses are refunded by the government. Housing consumption is determined as a more or less fixed share of the housing capital, which corresponds to the way this component of private consumption is calculated in the national accounts. The private consumption variable we consider covers approximately 75 per cent of total private consumption.

[11] Following Erlandsen and Nymoen (2008) the real interest rate interact with a step dummy which takes on the value one as of 1984(1) – when direct regulations on the interest rates was lifted – and zero before that date. The age composition variable gives the number of persons between 50 and 66 years old as a fraction of the adult population over 20 years old.

[12] Details about the tests are available upon request from the author (eja@ssb.no or tskrogh@gmail.com). The test are carried out for the full sample, with the exception of RR, where we look at the period after deregulation of the credit markets (1984(2) − 2007(4)). Impulse dummies are used to control for short spells with extraordinarily high or low interest rates in 1997, 1998, 2003 and 2004, see footnote 3 in Erlandsen and Nymoen (2008).

Table 2. Johansen tests of cointegration based a conditional VAR in consumption, income and wealth

Eigenvalue	Hypotheses on rank		Trace tests (λ_{trace})	
λ_i	H_0	H_1	λ_{trace}	5 % critical level*
0.158	$r = 0$	$r \geq 1$	59.81	57.32
0.118	$r \leq 1$	$r \geq 2$	33.60	35.96
0.090	$r \leq 2$	$r = 3$	14.49	18.16

VAR system of order 5, 1970(3)-2008(2), endogenous variables (logs) c, y and w, exogenous variables AGE, RR and trend; deterministic variables, Constant, CPSTOP, CS1, CS2, CS3.

Tests of the VAR(5) system **)

Vector AR(1-5) test : $F(45,333) = 1.24$ [p-value = 0.15]

Vector normality test : $\chi^2(6) = 4.20$ [p-value = 0.65]

Vector heteroskedasticity test : $F(216,524) = 0.84$ [p-value = 0.94]

*) Critical values are from Table 13 in Doornik (2003) – for the case of two exogenous variables in the system

**) See Doornik and Hendry (2006b).

Table 3. Tests of overidentifying restrictions for the cointegration vector in a cointegrated VAR (r=1) − derived from the VAR(5) system in Table 2. 1970(3)-2008(2).

i) Model without testable restrictions (note: $\beta_c = 1$ is an identifying restriction)

$c + \beta_y y + \beta_w w + \beta_{AGE} AGE + \beta_{RR} RR + \gamma$ Trend; no restriction on α_c α_y α_w

Log L =1166.25

ii) Model: $\gamma = 0$

$$c - \underset{(0.05)}{0.79}\, y - \underset{(0.03)}{0.18}\, w + \underset{(0.12)}{0.07}\, AGE + \underset{(0.30)}{0.70}\, RR; \quad \alpha_c = -\underset{(0.10)}{0.48}.\ \alpha_y = \underset{(0.10)}{0.02},\ \alpha_w = -\underset{(0.13)}{0.18}$$

Log L = 1165.67 $\chi^2(1) = 1.15$ [p-value = 0.28]

iii) Model: $\gamma = 0$, $\beta_{AGE} = 0$

$$c - \underset{(0.05)}{0.78}\, y - \underset{(0.03)}{0.18}\, w + \underset{(0.30)}{0.58}\, RR; \quad \alpha_c = -\underset{(0.10)}{0.48},\ \alpha_y = \underset{(0.09)}{0.01},\ \alpha_w = -\underset{(0.12)}{0.16}$$

Log L = 1165.53 $\chi^2(2) = 1.43$ [p-value = 0.49] $\chi^2(1) = 0.28$ [p-value = 0.59]

iv) Model: $\gamma = 0$, $\beta_{AGE} = 0$, $\alpha_y = 0$, $\alpha_w = 0$

$$c - \underset{(0.05)}{0.74}\, y - \underset{(0.03)}{0.20}\, w + \underset{(0.20)}{0.49}\, RR; \quad \alpha_c = -\underset{(0.10)}{0.47}$$

Log L = 1164.93 $\chi^2(4) = 2.62$ [p-value = 0.62] $\chi^2(2) = 1.20$ [p-value = 0.55]

v) Model: $\gamma = 0$, $\beta_{AGE} = 0$, $\alpha_y = 0$, $\alpha_w = 0$, $\beta_y + \beta_w = -1$

$$c - \underset{(-)}{0.85}\, y - \underset{(0.02)}{0.15}\, w + \underset{(0.22)}{0.71}\, RR; \quad \alpha_c = -\underset{(0.08)}{0.38}$$

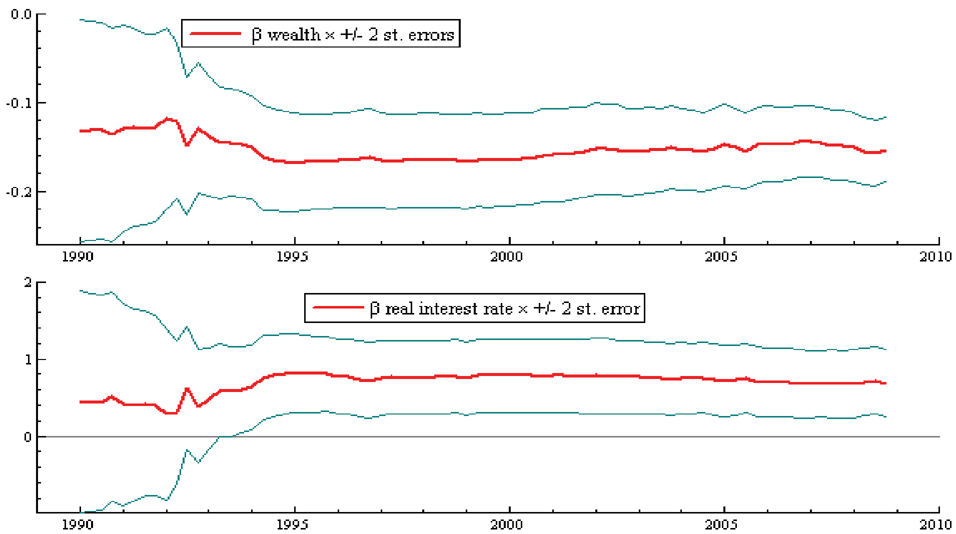
Log L = 1163.37 $\chi^2(5) = 5.75$ [p-value= 0.33] $\chi^2(1) = 3.13$ [p-value = 0.08]

When we condition on the rank being one we can test further restrictions on the coefficients in the cointegration vector. Table 3 shows that the deterministic trend is insignificant in the cointegrated VAR (p-value=0.28). Likewise, we can drop the AGE variable (p-value = 0.59). More importantly, both y and w can be considered weakly exogenous with respect to the long run parameters in the cointegrating vector (p-value = 0.55 for the joint test), and the data support a restriction of homogeneity in income and wealth, albeit marginally (p-value = 0.08), which means that the long run relationship can be written as:

2) $$ecm_{At} = c_t - 0.85\, y_t - 0.15\, w_t + 0.7\, RR_t$$

Figure 1 demonstrates that the estimated long run coefficients are recursively stable.

Figure 1. Consumption function A with wealth effect in the long run: Recursive long run coefficients estimates for wealth and real interest rate, see Table 3 v), where $\beta_y + \beta_w = -1$. 1990(1)–2008(2)



Next, we have estimated the conditional consumption function on a sample 1971(1) – 2008(4), general to specific, starting with a general specification with four lags in Δc, Δy, Δw, and ΔRR. We have included a double set of seasonal dummies as of 2002(1) in order to pick up the effects of a change in the way the quarterly income statistics are constructed. In Table 4 we have estimated the model both with the equilibrium-correction term as a variable (Model A1) and the level variables specified separately (Model A2). Both models appear to be well specified according to diagnostic tests, and from Figures 12 and 13 of Appendix C it can be seen that the coefficients are recursively stable. The table also includes a Model A3, which is identical to Model A1, apart from being estimated on the sample 1986(3) – 2008(4) in order to facilitate a comparison with the alternative consumption functions in sections 4 and 5 below, which are estimated with data from this time period. The coefficients remain stable, but exhibit smaller t-values. Model A1 is re-estimated on a sample ending in 2005(4) and Figure 2 shows one-step-ahead forecasts for the growth in C (Δc_t). The forecast tests reported support parameter stability and the model tracks the development over those 12 quarters satisfactorily. Repeating the experiment with Model A3 yields conditional forecasts that are almost as good as for Model A1 – the p-value for the forecast $\chi^2(12)$- test drops from 0,99 to 0,92, partly because of a smaller standard error of regression.

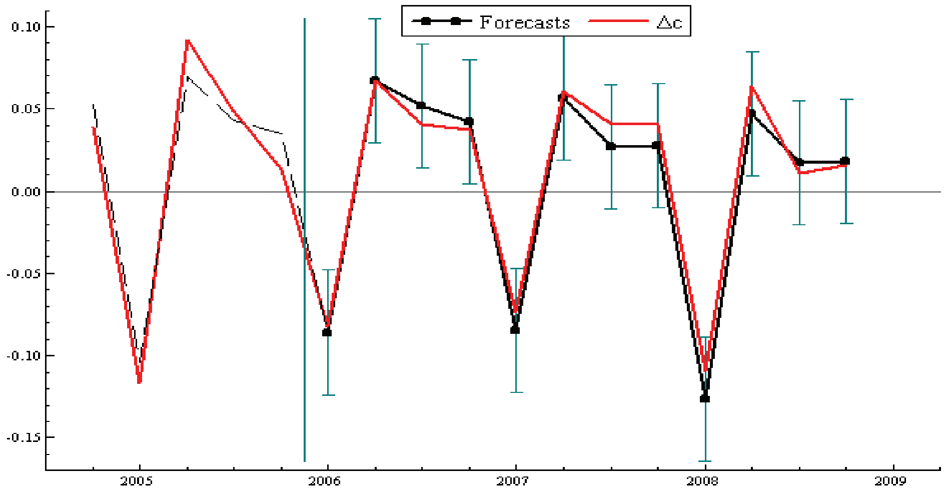
Table 4. Consumption function A for Δc_t where $ecm_{At} = c_{t-1} - 0.85\, y_{t-1} - 0.15\, w_t + 0.7\, RR_{t-1}$
Dependent variable: Δc_t

Model (Sample)	A1 (1971(1) – 2008(4))		A2 (1971(1) – 2008(4))		A3 (1986(3) – 2008(4))	
	Coeff.	t-value	Coeff.	t-value	Coeff.	t-value
Δc_{t-4}	0.25	4.14	0.25	4.23	0.23	2.78
Constant	-0.22	-6.15	0.10	0.66	-0.27	-5.92
Δy_t	0.24	3.14	0.24	3.12	0.23	1.81
Δy_{t-1}	-0.28	-3.52	-0.26	-3.27	-0.34	-2.19
Δy_{t-2}	-0.29	-3.91	-0.29	-3.89	-0.26	-1.95
Δw_{t-1}	0.25	4.23	0.24	4.04	0.12	1.64
ecm_{At}	-0.51	-8.30			-0.59	-7.55
c_{t-1}			-0.55	-8.55		
y_{t-1}			0.41	7.77		
w_t			0.10	5.62		
RR_{t-1}			-0.28	-2.70		
dummies						
SER (st. error regression)	1.82 %		1.81 %		1.66 %	
Tests*:	statistics	Value [p-value]	statistics	Value [p-value]	statistics	Value [p-value]
AR 1-5	F(5,133)	1.61[0.16]	F(5,130)	1.62[0.16]	F(5,71)	0.79[0.56]
ARCH(4)	F(4,130)	0.69[0.60]	F(4,127)	0.65[0.63]	F(4,68)	0.22[0.92]
Normality	$\chi^2(2)$	0.09[0.96]	$\chi^2(2)$	0.91[0.63]	$\chi^2(2)$	0.77[0.68]
Heteroskedacity	F(19,118)	1.07[0.39]	F(25,109)	0.98[0.49]	F(19,56)	0.49[0.95]
Reset	F(1,137)	0.41[0.52]	F(1,134)	0.02[0.89]	F(1,75)	0.62[0.43]

*) See Doornik and Hendry (2006a)

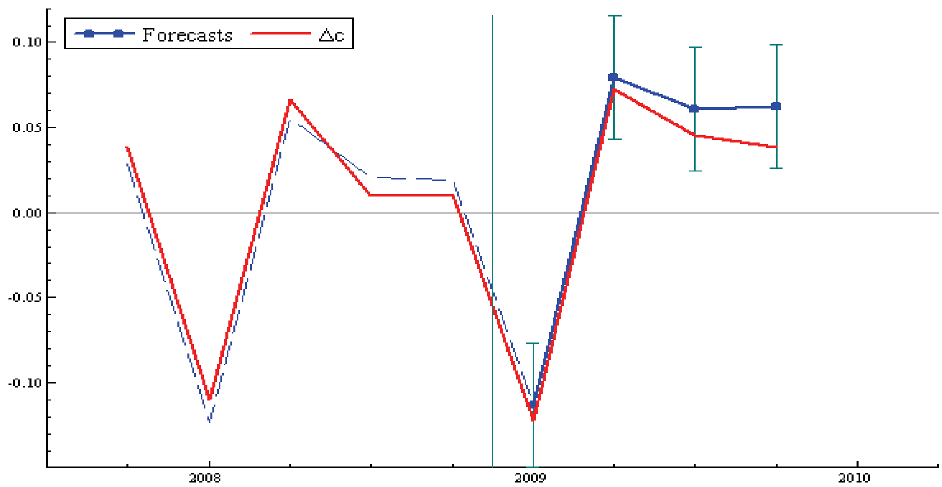
Preliminary data for 2009 allow evaluation of *ex ante* forecasts with the conditional consumption function. Norway was the first country in Europe to recover through 2009 from the slump following the impulses of the international financial crises. The Norwegian Government countered these impulses by increasing annual public spending 3 percentage points above the trend growth (of about 3 per cent) in 2009 and the central bank signal interest rate was cut from 5.75 per cent in October 2008 to 1.25 per cent in June 2009. Together this brought about an increase in the households' real disposable after-tax income of around 4 percent in 2009 despite a GDP growth of -1.5 per cent. Also the household wealth recovered early in 2009, see footnote 15 below. Figure 3 shows the one step ahead forecasts for the four quarters of 2009 for the consumption function with wealth effects (A1). The change in consumption is slightly overpredicted in the first two quarters and more so in the last two quarters of 2009, even though the model passes the forecast tests. The overpredictions suggest that there may have been precautionary savings due to the uncertainty perceived by the public, which is not fully captured by the current specification of the model.

Figure 2. The consumption function with wealth effects in the long run: Model A1 of Table 4. One-step-ahead forecasts for Δc_t, 2006(1) – 2008(4), estimation period 1971(1) – 2005(4).



Forecast-tests: Forecast $\chi^2(12) = 3.59$ [p- value = 0.99], Chow: $F(12,126) = 0.24$ [p-value = 0.99], CUSUM $t(11) = 0.70$ [p- value = 0.50], see Doornik and Hendry (2006a)

Figure 3. *Ex ante* forecasts with the consumption function with wealth effects in the long run: Model A1 of Table 4. One-step-ahead forecasts for Δc_t, 2009(1) – 2009(4), estimation period 1971(1) – 2008(4)



Forecast-tests: Forecast $\chi^2(4) = 2.82$ [p-value = 0.59], Chow: $F(4,138) = 0.56$ [p-value = 0.70], see Doornik and Hendry (2009).

4. Comparison with consumption models based on Euler equations

We shall first compare the consumption function of the previous section with two canonical specifications where the macro consumption is based on a model of a representative agent who optimizes consumption over time according to Hall (1978). If there is no credit rationing and the agent has rational expectations about future income, the growth in consumption is independent of current income as only surprise, i.e. unforeseen, income matters. If the real interest rate is constant, consumption becomes a random walk

$$(3) \qquad \log (C_t) = \text{constant} + \log (C_{t-1}) + \varepsilon_t ,$$

where ε_t is a white noise error term. Deviations from trend growth in consumption are unpredictable according to (3). If, however, the real interest rate varies over time, theory predicts that the consumption growth will vary accordingly.

As a reference point we have estimated the model E1, which is the same as (3) augmented with dummies to pick up seasonal variation in the (unadjusted) data. From Table 5 it is seen that the fit deteriorates compared to the models of Table 4 as measured by the standard errors of regression. If we add the after tax real interest rate (RR) or lagged level of consumption, the latter to represent habit formation, these variables are insignificant both separately and jointly. However, we notice that the model E1 suffers from autocorrelation in the residuals. Adding to the equation lagged dependent variables up to fourth order does not remove this autocorrelation, but in this case the after tax real interest rate and lagged consumption come out nearly significant and the fit is improved by 14 % (model E2). We note that the real interest rate has an unexpected, positive sign. Recursive estimation shows that the coefficient estimates for both models are stable, except for the last two quarters of 2008 for model E2. Despite large forecast errors, as is seen from Figure 4 for the random walk model E1, forecast tests based on one step ahead forecasts for the period 2006(1) – 2008(4) do not reject the two models in Table 5. This is of course largely due to the sizable standard errors of regression, which have a marked influence on the outcome of these tests.[13]

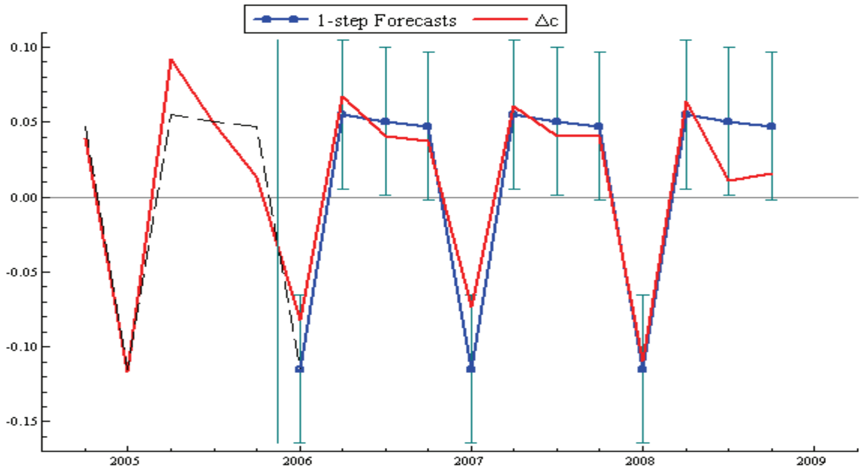
[13] Model E2 corresponds to the consumption Euler equation in the current version of the Norges Bank's DSGE model NEMO, as described in Brubakk and Sveen (2009). A difference is that in NEMO consumption today depends not only on today's interest rate level, but also on expectations about future interest rates, i.e. the entire interest rate path.

Table 5. Consumption models based on Euler equations. Dependent variable: Δc_t

Model (Sample)	E1. (1986(3) – 2008(4))		E2. (1986(3) – 2008(4))	
	Coefficient	t-value	Coefficient	t-value
Constant	0.09	15.1	-0.69	-1.55
Δc_{t-1}			-0.49	-4.27
Δc_{t-2}			-0.33	-2.67
Δc_{t-3}			-0.12	-0.93
Δc_{t-4}			0.11	2.92
RR_t			0.29	1.94
c_{t-1}			0.04	1.78
dummies				
SER (st. error regression)	2.40 %		2.06 %	
Tests*:	statistics	Value [p-value]	statistics	Value [p-value]
AR 1-5	F(5,77)	4.01[0.003**]	F(5,71)	2.88[0.02*]
ARCH(4)	F(4,72)	0.91[0.46]	F(4,82)	0.87[0.49]
Normality	$\chi^2(2)$	3.03[0.22]	$\chi^2(2)$	1.27[0.53]
Heteroskedacity		n.a.	F(19,70)	1.05[0.42]
Reset		n.a.	F(2,74)	0.34[0.56]

*) See Doornik and Hendry (2009)

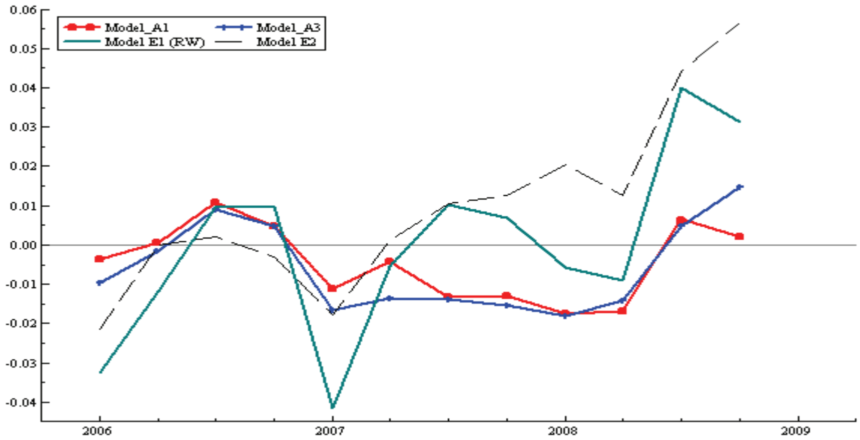
Figure 4. The random walk consumption model: Model E1 of Table 5. One-step-ahead forecasts for Δc_t, 2006(1) – 2008(4), estimation period 1986(3) – 2005(4)



Forecast-tests: Forecast $\chi^2(12) = 9.77$ [p-value = 0.64], Chow: F(12,70) = 0.56 [p-value = 0.86], CUSUM t(11) = -0.18 [p-value = 0.86] see Doornik and Hendry (2009)

We can however compare the forecast errors of the consumption function A1 in Table 4 with those from Table 5 directly. In order to control for different sample length we have chosen also to include Model A3 from Table 4. We see from Figure 5 that the random walk E1 has difficulties in tracking the strong consumption growth in early 2007 and it overpredicts the consumption growth in 2008(3) and 2008(4) when actual consumption dropped as the financial crisis hit Norway at full force, see Figure 9 (right panel) below. The consumption functions of type A avoid big errors even though it misses out some of the strong consumption growth in 2007 and in the beginning of 2008.

Figure 5. Comparison of forecast errors (forecast – actual change, one-step-ahead) for Δc_t, 2006(1) − 2008(4), for the consumption function estimated over two different sample lengths (models A1 and A3 of Table 4) and the Euler equations (model E1 and E2 of Table 5). Model A1 is estimated on the sample 1971(1)−2005(4), whereas models A3, E1and E2 are estimated on the sample 1986(3) −2005(4)



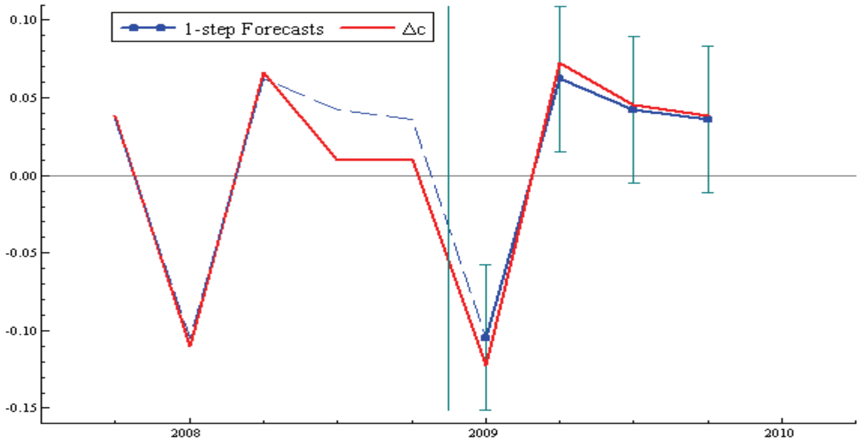
We have carried out forecast encompassing tests for the forecast errors from the consumption functions A1 and A3 against the forecast errors of the Euler equations E1 and E2 in Table 6. The test consists in running a static regression of the forecast errors of Model 1 on the difference between the forecast errors of an alternative Model j and Model 1. Under H0 that Model j does not have explanatory power in excess of what is already explained by Model 1, the regression coefficient is zero, see Bårdsen *et al.* (2005, p.178) for details. From the table it is seen that model A encompasses the alternative specification clearly while the opposite is not the case. Also, the model based on the longer sample period has the best forecasting performance, quite in line with the visual impression in Figure 5.

Table 6. Encompassing tests of the forecast errors of model i and model j, where i=E1,E2 and j= A1, A3. 2006(1) - 2008(4). The p-values are calculated for the coefficients of the forecast error differences

Version of model (j)	Model E1 vs. Model j	Model j vs. Model E1
Model A1	0.0000**	0.91
Model A3	0.002**	0.56
Version of model (j)	Model E2 vs. Model j	Model j vs. Model E2
Model A1	0.0000**	0.38
Model A3	0.0003**	0.88

If we extend the testing window to include also data for 2009, the outcome of the forecast encompassing tests remains the same. But the random walk model E1 now forecasts 2009(3) and 2009(4) better than model A1, according to the ex ante forecasts shown in figure 6 (to be compared with figure 3 of section 2). This is so because the quarterly consumption growth − corrected for seasonal differences − is back on the quarterly trend growth in Norway in those two quarters.

Figure 6. ***Ex ante*** **forecasts with the random walk model: Model E1 of Table 5. One-step-ahead forecasts for Δc_t, 2009(1) – 2009(4), estimation period 1986(3) – 2008(4).**



Forecast-tests: Forecast $\chi^2(4) = 0.80$ [p-value = 0.94], Chow: $F(4,137) = 0.18$ [p-value = 0.95], see Doornik and Hendry (2009)

One possible caveat against the above comparisons between the Euler equations and the consumption functions of section 3 could be that the proprietors of Euler based models almost surely would adopt seasonally adjusted data. In Appendix B we have redone the above analysis with seasonally adjusted data and we show that all main conclusions are robust to making such transformations of the data.

5. Does wealth matter?

In the following we shall also compare the empirical properties of consumption function A of Section 2, where consumption in the long run is determined by income, wealth and the after-tax real interest rate with an alternative without explicit long run wealth effects, in the sequel called model B.[14] According to consumption function B the long run is given by:

(4) $$ecm_{Bt} = c_t - 0.98\, y_t - 0.02\, ya_t + 1.7\, RR_t,$$

Equation (4) is homogeneous in the total real after-tax disposable income of the households, where we distinguish between equity incomes (ya) and non-equity incomes (y). The effect of equity incomes on long run consumption is almost negligible, there is no wealth effect and the semi-elasticity of real after-tax interest rates is 2.5 times the corresponding estimate in (2) based on consumption function A. An ADF test of the residuals from (4) supports stationarity with a Dickey-Fuller t-statistics of -4.1.

[14] Incidentally, both alternatives have been used in the Statistics Norway quarterly model KVARTS, where the consumption function with wealth effects (A) was substituted for the one without (B) in February 2009.

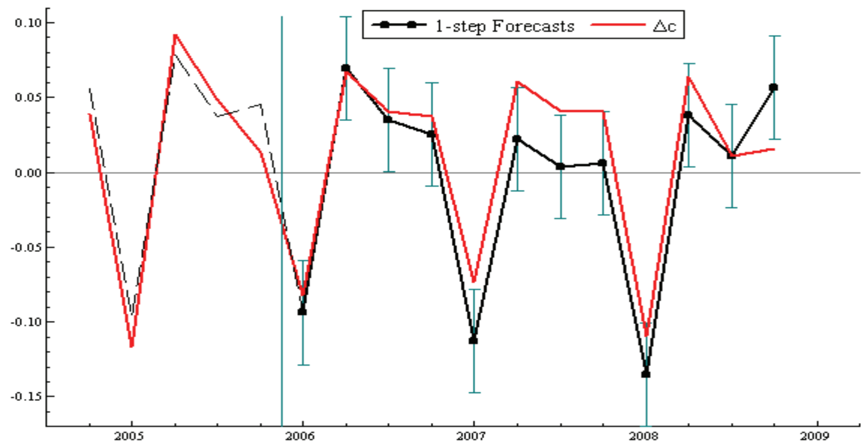
Table 7. Consumption function B without explicit wealth effects in the long run, where $ecm_{Bt} = c_{t-2} - y_{t-2} - 0.02\,(ya - y)_{t-2} + 1.7\,RR_{t-1}$. Dependent variable: Δc_t

Model (Sample)	B. (1986(3) – 2008(4))	
	Coefficient	t-value
Δc_{t-1}	-0.70	-8.33
Δc_{t-2}	-0.47	-5.77
Δc_{t-3}	-0.24	-2.90
Constant	0.06	4.92
Δw_{t-1}	0.30	4.10
ecm_{Bt}	-0.23	-3.97
dummies		
SER (st. error regression)	1.82 %	
Tests*:	statistics	Value [p-value]
AR 1-5	F(5,75)	0.42[0.83]
ARCH(4)	F(4,72)	0.85[0.50]
Normality	$\chi^2(2)$	1.36[0.50]
Heteroskedacity	F(14,65)	0.62[0.83]
Reset	F(1,79)	0.34[0.56]

*) See Doornik and Hendry (2006a)

As we did with model A, we have modelled B general to specific, starting with four lags in Δc, Δy, Δya, and Δw. The resulting model B reported in Table 7 contains only own dynamics and lagged wealth growth in the short run, and the model exhibits a poorer fit than the does the model A3 on the same sample period (1986(3) – 2008(4)).[15]

Figure 7. The consumption function without wealth effects in the long run: Model B of Table 5. One-step-ahead forecasts for Δc_t, 2006(1) – 2008(4), estimation period 1986(3) – 2005(4). Forecast-tests:



Forecast $\chi^2(12) = 27.80$ [p-value = 0.0059**], Chow: $F(12,67) = 1.54$ [p-value = 0.13],
CUSUM $t(11) = 1.70$ [p-value = 0.12], see Doornik and Hendry (2006a)

[15] We have also estimated the model with an information set which includes lags in ΔRR. This improves the fit slightly, but yields an implausible positive estimate for the short run effect of a change in the real interest rate.

Model B shows signs of instability over the last 12 quarters of the observation period, cf. Figure 14 of Appendix C. This leads to several one-step-ahead forecasts outside of the ± 2 standard error confidence band when we re-estimate this model with a sample ending in 2005(4), see Figure 7.

Again we can compare directly the forecast errors of the consumption functions A1 and A3 in Table 4 with those from the alternative B in Table 7. We see from Figure 8 that the consumption function B understates the strong consumption growth in 2006(3) – 2008(2) and fails in the opposite direction in 2008(4) in the quarter following the fallout of Lehman Brothers in September 2008. The consumption functions of type A avoid big errors even though it misses out the strength of the consumption growth in 2007 and into the beginning of 2008. Model A1 forecasts the drop in consumption growth in 2008(4) spot on.

Figure 8. Comparison of forecast errors (forecast – actual change, one-step-ahead) for Δc_t, 2006(1) - 2008(4), for the consumption function with wealth effects estimated over two different sample lengths (models A1 and A3 of Table 4) and the consumption function without wealth effects (model B of Table 5). Model A1 is estimated on the sample 1971(1) - 2005(4), whereas models A3 and B are estimated on the sample 1986(3) - 2005(4)

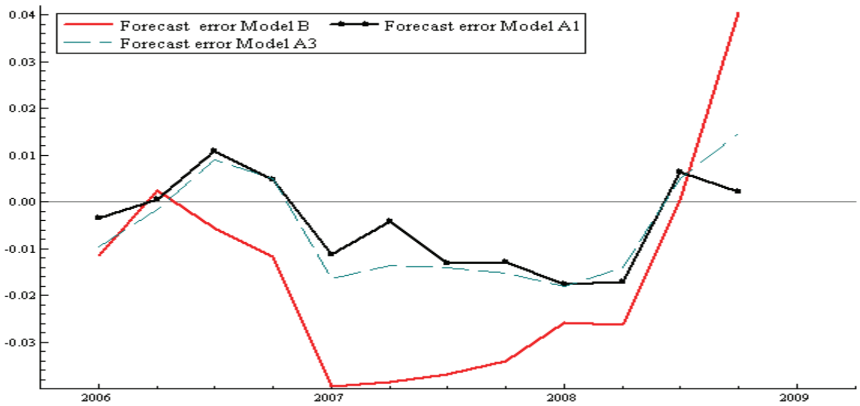


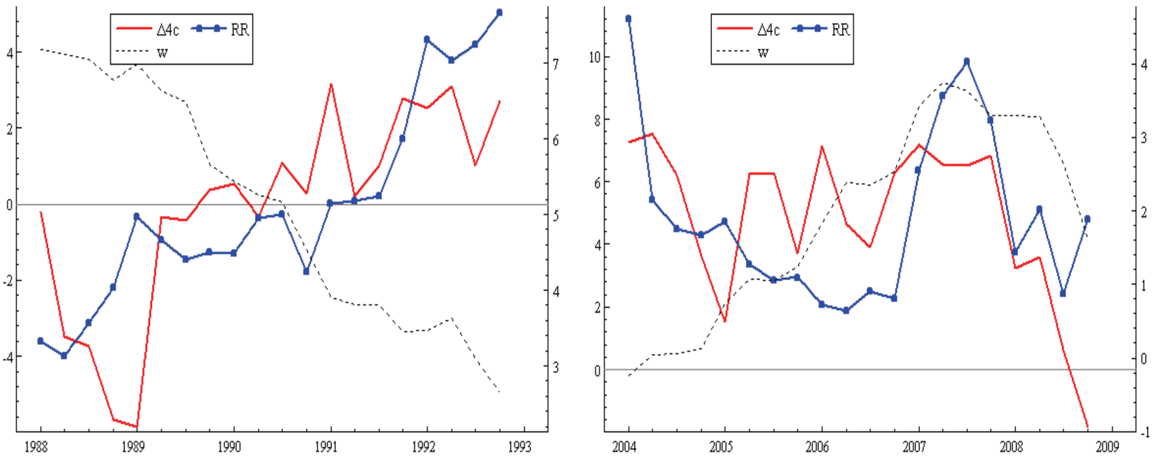
Table 8. Encompassing tests of the forecast errors of model B and model j, where j= A1, A3. 2006(1) - 2008(4). The p-values are calculated for the coefficients of the forecast error differences

Version of the new model (j)	Model B vs. Model j	Model j vs. Model B
Model A1	0.0002**	0.64
Model A3	0.016*	0.87

As in the preceding section we have carried out forecast encompassing tests for the forecast errors from the competing consumption functions in Table 8. From the table it is seen that model A encompasses the alternative specification clearly while the opposite is not the case. Also, the model based on the longer sample period has the best forecasting performance, quite in line with the visual impression in Figure 8.

That said, the two alternative models (with and without a long run wealth effects) seem to explain equally well what happened to private consumption in Norway before and under the banking crisis in the early 1990s. As a matter of fact there is little difference in their explanatory power up to and including 2005(4). In the following three years the model with long run wealth effects outperforms the competitor in forecasting, as we have seen above. Figure 9 gives us a simple explanation why including wealth becomes so important. In the period 1988(1) – 1992(4) the correlation between w and RR was equal to −0.85 whilst it was +0.76 in the period 2006(1) – 2008(4). Looking at the period 1986(3) – 2005(4) the correlation was −0.54. Since the estimated effects of the two variables have opposite signs it follows that at most one of the models will be able to explain the development *after* 2006 in a satisfactory way. This is a class-room example of how a changing correlation pattern can reveal that an important variable has been omitted in a model specification.

Figure 9. Annualised consumption growth ($\Delta_4 c$, left axis) is plotted against real interest rate (RR, right axis) and wealth (w, not on the axis) for the periods 1988(1)−1992(4), left panel, and 2004(1)−2008(4), right panel



Interestingly, this change in correlation between wealth and the real interest rate reflects a change in the monetary policy system, see Bjørnstad and Jansen (2007). Norway had an exchange rate target during the banking crisis 20 years ago and the procyclical interest rate setting was a consequence of the central bank defending the exchange rate against devaluation pressure. With the adoption of an inflation target for the central bank as of 2001(2), the scene was set for a countercyclical real interest rate when the financial crisis arrived.

6. Consequences for the savings rate

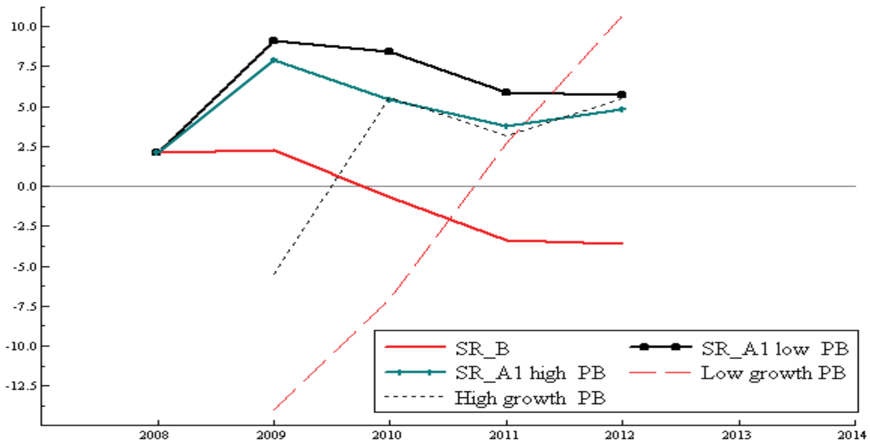
The long run relationships (2) and (4) for the competing consumption functions A and B both imply a constant savings rate in steady state. Model B without wealth effects on consumption in the long run has a constant ratio between income and consumption, provided that the consumption of health and housing services is proportional to the remainder, which the consumption measure adopted here. This constant ratio will however vary for different levels of the long run after-tax interest rate.

The consumption function A will – *ceteris paribus* – exhibit the same property along a growth path where income and wealth are growing proportionally. The savings rate is then:

(5) $\log((Y-C)/Y) = \text{constant} + 0.15 \log(Y/W)$

If real wealth is subject to a negative shock – as we observed in Norway throughout the period 1988(1)-1992(4) in Figure 9 (left panel) and in all quarters of 2008 in Figure 9 (right panel) – the response on the part of the consumers will be an increase in the savings rate when equation (5) applies.

Figure 10. Simulation results from the macroeconometric model KVARTS showing the savings rate (SR) under two alternative assumption about future housing prices (PB) for the consumption function with wealth effects in the long run (A1_high_PB, A1_low_PB) and for the consumption function without wealth effects in the long run (model B). The savings rate and annual growth in housing prices are percentages



Some simulations with the quarterly macroeconometric model KVARTS based on the information set underlying the forecasts for the Norwegian economy 2009(1)-2012(4) − as published in Statistics Norway's *Economic Survey for the year 2008* − may shed some light on these issues. Keeping other exogenous assumptions unchanged, we have examined for two model versions, which only differ with regard to the consumption function, the consequences for the savings rate under two scenarios for the housing prices in 2009 and 2010, one high and one low in 2009 and 2010, where the difference in the growth rate for housing prices is 10 percentage points each year.[16] We have in other words manipulated the housing prices exogenously and disregarded feedback effects, which would emerge if, say, the central bank reacts to changing housing prices in the interest rate setting. The feedback effects are judged to be small since the inflation targeting central bank focuses on a CPI measure with low pass-through of housing price changes, and we achieve in this way to focus on the partial wealth effects brought about by the housing prices in the consumption function. In Figure 10 we observe that there is no difference between the two housing price scenarios for the consumption function without wealth effects: the savings rate remains low and becomes negative due to low real interest rates and positive real income growth. For the new consumption function on the other hand we see a financial consolidation in the households through a considerable increase in the savings rate into and through 2009, and this increase is strengthened if we look at the low housing price scenario in 2009 and 2010. In 2011 the difference between the savings rates with and without wealth effects in the consumption function is 10 percentage points under the low housing price scenario and 8 percentage points for the high housing price scenario. These findings appear to be quite robust to alterations in other exogenous assumptions underlying the simulations.

7. Conclusions and further work

We have found that a consumption function including wealth effects in the short and long run explains the quarterly changes in consumption over the years 2006-2008 better than Euler based equations and also better than an alternative consumption function without a long run wealth effect. The wealth effects are shown to be strong enough to lift the savings rate considerably and to counteract the expansive effects of a low interest rate, which both has a positive direct effect on consumption and an indirect effect via a marked increase in real disposable income for the household sector, which has a large net interest bearing dept to other sectors.

One issue that has given rise to debate among the Norwegian economists is the measurement of housing prices in Norway before and under the credit market deregulation. No official housing price index exists for the years before 1992, and the housing prices adopted by Brodin and Nymoen (and

[16] At this point – with the advantage of hindsight – we can conclude that both scenarios are counterfactual as far as 2009 is concerned. Housing prices have risen more in 2009 than in the high scenario. Moreover, regarding financial wealth of the households, it is in both scenarios assumed that the index at the Oslo stock exchange flattens out in 2009(2) and increases at a moderate rate for the rest of the simulation period. As of the beginning of 2010 we know that the stock exchange has recovered much faster after a turning point at the beginning of 2009(2).

used by the subsequent studies reported here) were therefore extracted from several sources. It has been argued that their price index excludes a housing price hike in 1982 (due to the lifting of price regulations on small flats in urban regions) before the deregulation of the credit markets and that it likewise overstates the housing price increase shortly after financial liberalisation. Hence, it is possible that the housing price index used in the studies cited (and in this paper) actually picks up a combined effect – or, if you like, an interaction effect - of income and wealth and of credit market deregulation. Indeed, empirical work based on alternative housing price indices does suggest a break in consumption pattern in the wake of deregulation, much in line with the findings of Barrell and Davis (2007). [17]

17 In fact, the work plan at Statistics Norway includes ideas for constructing a credit conditions index (inspired by the work of John Muellbauer and co-authors cited in Section 1) in order to quantifying the separate effects of deregulation.

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Appendix A: Data definitions and sources

The data used in this study are collected from Statistics Norway's KVARTS database and Norges Bank's model database RIMINI (the 2005 version).

Symbol definitions and sources:

AGE – P50-66 / (P20-49 + P67up) , where P50-66 is population 50-66 years old, P20-49 is is population 20-49 years old and P67up is population more than 66 years old. Source: Statistics Norway.

C – Consumption expenditure in households and ideal organisations excluding expenditure on heath services and housing, fixed 2006 prices. Source: Statistics Norway.

CPI – Headline consumer price index (2006=1). Source: Statistics Norway.

PC – Price deflator for total consumption expenditure in households and ideal organisations (2006=1). Source: Statistics Norway.

RLB – Average interest rate on households' bank loans (Source: Norges Bank)

RR – Marginal after-tax real interest rate for households. 1970(1) – 2003(4), zero. 1984(1) – 2009(4), RLB ($1 - \tau$) + Δ_4 log CPI, where τ is marginal income tax rate for households. Source: Statistics Norway and Norges Bank.

CPSTOP – Income policy dummy; constructed to catch up the inflationary pressure that build up during the wage and price freeze in 1978. It takes non-zero values from 1979(1) to 1980(1) and zero elsewhere. See Brodin and Nymoen (1992) for details.

W – Real household wealth: nominal household wealth (financial and housing wealth) deflated by PC. Data for the period 1970(1) to 1992(3) for nominal household wealth is from Erlandsen and Nymoen (2008), source: Norges Bank, are chained in 1992(3) with data from the KVARTS database. (Source: Statistics Norway).

Y – Households' disposable income, excluding equity income; nominal income deflated by PC. (Source: Statistics Norway).

YA – Households' after-tax equity income; nominal income deflated by PC. (Source: Statistics Norway).

Appendix B: The forecast comparisons of section 4 with seasonally adjusted consumption data: the consumption function vs models based on Euler equations.

As we alluded to at the end of section 4, one possible objection against the comparisons we make between the Euler equations and the consumptions function is that the proprietors of Euler based models would be likely to use seasonally adjusted data. This makes a comparison more difficult as seasonal adjustment changes the series to be explained, and more so than the filter defined by deterministic dummy variables. That said, we have adopted a X12-ARIMA filter (with corrections for workdays and public holidays) to the C_t variable to obtain C_t*. This reduces the standard error of the regression to 0.016 for Model E1* in table 9 as compared to 0.024 for the case with deterministic dummies, cf. Model E1 in table 5. For Model E2 and Model E2* the corresponding reduction is from 0.021 to 0.015. The seasonal adjustment filter removes autocorrelation from the series and hence own dynamics is dropped in Model E2*.

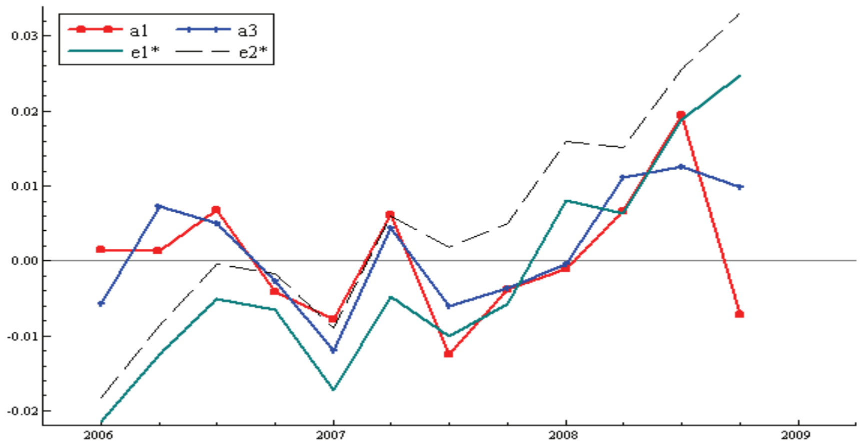
Table 9. Models based on Euler equations with seasonally adjusted consumption data. Dependent variable: Δc^*_t (s.a.)

Model (Sample)	E1*. (1986(3) – 2008(4))		E2*. (1986(3) – 2008(4))	
	Coefficient	t-value	Coefficient	t-value
Constant	0.06	3.8	-0.20	-1.90
RR_t			0.16	1.48
c^*_{t-1}			0.02	1.95
SER (st. error regression)	1.56 %		1.54 %	
Tests*:	statistics	Value [p-value]	statistics	Value [p-value]
AR 1-5	F(5,84)	2.37[0.05*]	F(5,82)	1.88[0.11]
ARCH(4)	F(4,82)	1.37[0.25]	F(4,82)	1.34[0.49]
Normality	$\chi^2(2)$	5.20[0.07]	$\chi^2(2)$	2.92[0.23]
Heteroskedacity		n.a.	F(4,85)	1.10[0.36]
Reset		n.a.	F(2,84)	0.44[0.56]

*) See Doornik and Hendry (2009)

In order to compare the forecasts from the consumption function model we have seasonally adjusted the forecasts for the period 2006(1) – 2008(4) from Model A1 and Model A3 together with the original series for C_t and calculated the corresponding forecasts for the growth rate in C_t* with the "true" Δc^*_t . The corresponding forecast errors thus obtained are shown in figure 11. The root mean squared forecast errors are 0.82 %, 0.78 %, 1.36 %, and 1.52 % for A1, A3, E1* and E2* respectively.

Figure 11. Comparison of forecast errors (forecast – actual change, one-step-ahead) for Δc_t, 2006(1) −2008(4), for the consumption function estimated over two different sample lengths (models A1 and A3 of Table 4) and the Euler equations (model E1* and E2* of Table 5).



Model A1 is estimated on the sample 1971(1) - 2005(4), whereas models A3, E1*and E2* are estimated on the sample 1986(3) −2005(4). Models E1*and E2* are estimated on seasonally adjusted consumption data, whereas the forecasts from Models A1 and A3 are seasonally adjusted after the forecasts are made as explained in the text

Finally, we can do the same forecast encompassing tests as in Table 6 for the seasonally adjusted forecast errors from the consumption functions A1 and A3 against the forecast errors of the Euler equations E1* and E2*. Table 10 shows that models of type A encompass the alternative specifications of type E clearly while the opposite is not the case. However, it is no longer clear that the model based on the longer sample period has the best forecasting performance, as Model A3 performs better than Model A1 in all 4 encompassing tests. Extending the sample to include the four quarters of 2009 does not alter these results.

Table 10. Encompassing tests of the forecast errors of model i and model j, where i=E1*,E2* and j= A1, A3. 2006(1) - 2008(4). The p-values are calculated for the coefficients of the forecast error differences.

Version of model (j)	Model E1* vs. Model j	Model j vs. Model E1*
Model A1 s.a	0.0014**	0.36
Model A3 s.a.	0.0006**	0.46
Version of model (j)	Model E2* vs. Model j	Model j vs. Model E2*
Model A1 s.a.	0.0007**	0.31
Model A3 s.a.	0.0004**	0.59

Appendix C: Supplementary figures (referred to in the text).

Figure 12. The consumption function with wealth effects in the long run: Model A1 of Table 4. Recursive coefficients estimates 1990(1)−2008(4). The sample is 1971(1) − 2008(4). The graphs show: i) Δc_{t-4}, ii) constant, iii) Δy_t, iv) Δy_{t-1}, v) Δy_{t-2}, vi) Δw_{t-1}, vii) $ecm_{At} = c_{t-1} - 0.85\ y_{t-1} - 0.15\ w_t + 0.7\ RR_{t-1}$, viii) One-step-ahead Chow-test, see Doornik and Hendry (2006a)

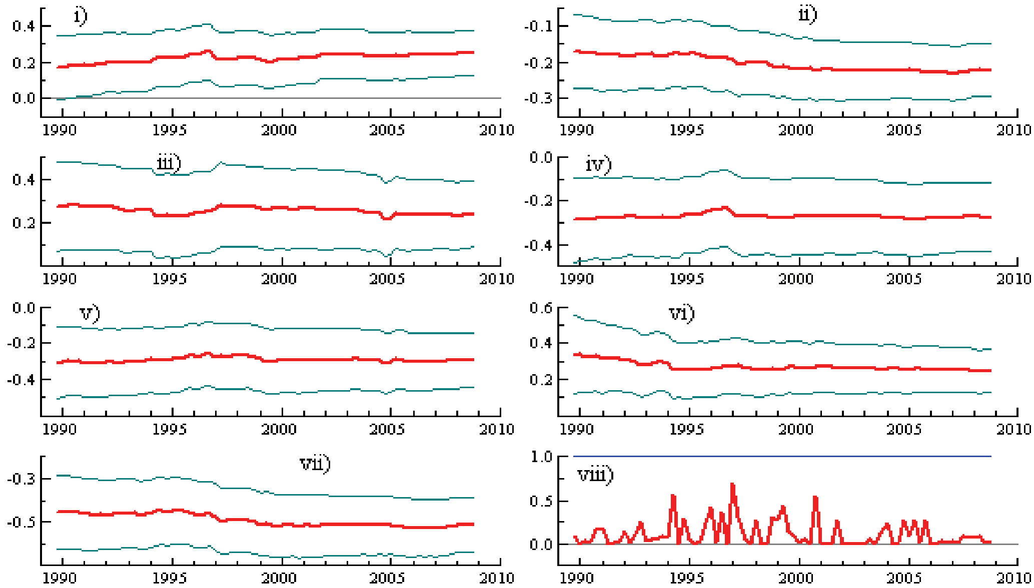


Figure 13. **Recursive long run coefficients estimates 1988(1) −2008(4) in the consumption function with wealth effects: Model A2 of tabell 4. The sample is 1971(1) − 2008(4). The graphs show: i) c_{t-1}, ii) y_{t-1}, iii) w_t, iv) RR_{t-1}, v) recursive residuals, vi) recursive forecast Chow-test, see Doornik and Hendry (2006a)**

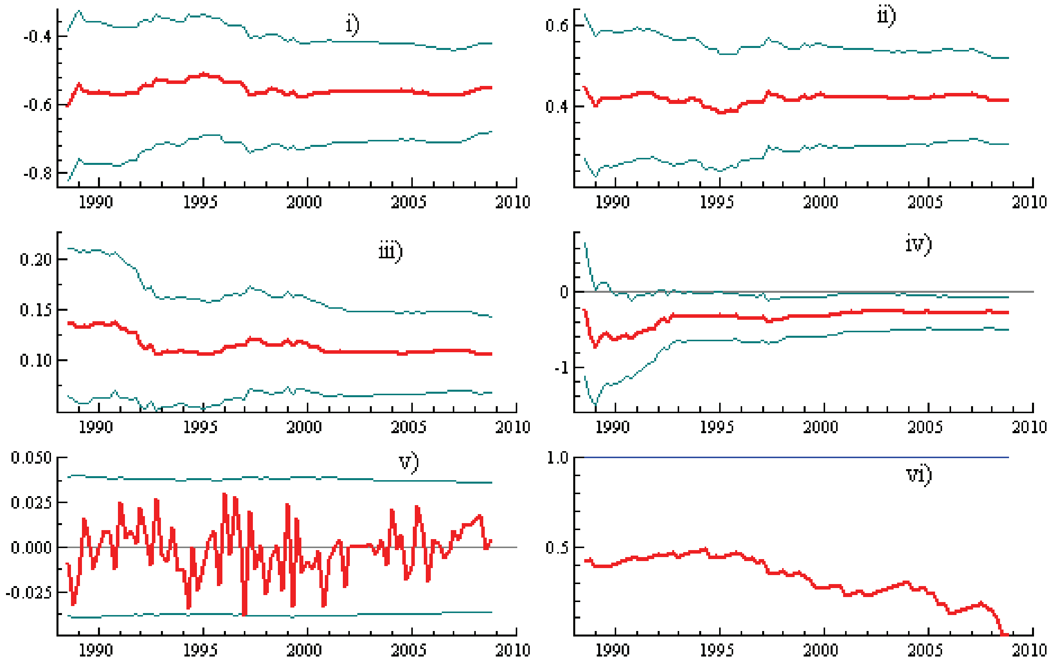


Figure 14. **The consumption function without wealth effect in the long run: Model B of Table 5. Recursive coefficient estimates 1995(1) −2008(4). Estimation period 1986(3) −2008(4). The graphs show: i) Δc_{t-1}, ii) Δc_{t-2}, iii) Δc_{t-3}, iv) constant, v) Δw_{t-1} vi) $ecm_{Bt} = c_{t-2} - y_{t-2} - 0.02\ (ya - y)_{t-2} + 1.7\ RR_{t-1}$, v) recursive residuals, vii) One-step-ahead Chow-test, see Doornik and Hendry (2006a)**