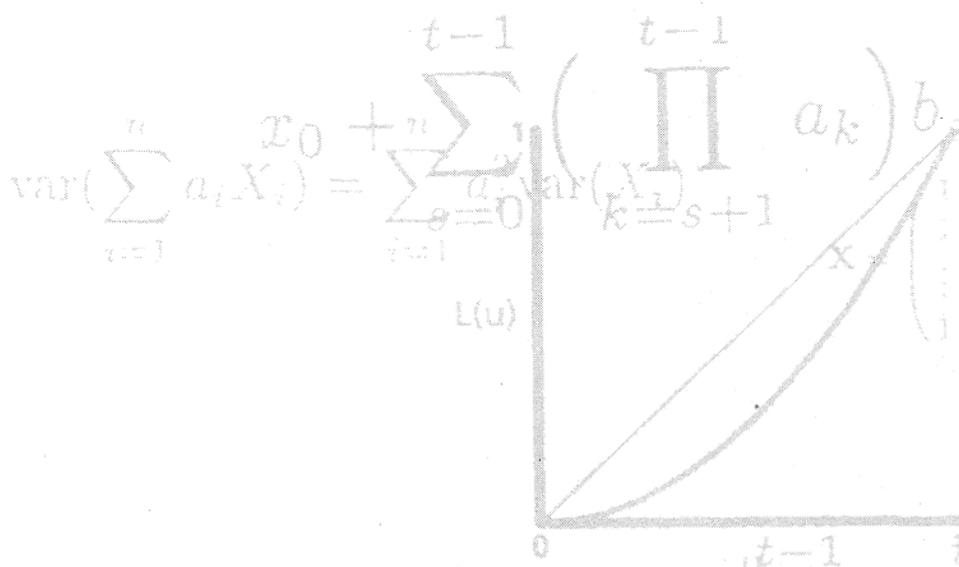Statistics Norway
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Financial Deregulation and Consumer Behavior: the Norwegian Experience

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Financial Deregulation and Consumer Behavior: the Norwegian Experience

Abstract:
The present paper uses the model by Campbell and Mankiw (1991) to examine the Norwegian consumer behavior and the role of the financial deregulation during the 1980s. For quarterly data on non-durables and services, we estimate the fraction of current income consumers to be in the range of 37% and 75% before the financial deregulation. This evidence indicates a substantial departure from the rational, forward-looking behavior, and there is thus reason to believe that liquidity constraints did bind the Norwegian consumer behavior until the mid 1980s. Our results further suggest that this evidence has disappeared after the financial deregulation in that the estimated fraction of current income consumers is essentially zero after 1985. This finding is so much more remarkable in that hardly any other aggregate time-series data set, from any country, conforms this closely with the forward-looking hypothesis.

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1. Introduction

During the 1970s and 1980s a number of European countries took steps to liberalize their financial markets. Norway, which used to have one of the most heavily regulated systems, completed its main deregulation effort in the mid—1980s. The purpose of the liberalization was to increase market efficiency by improving incentives and removing restrictions on optimization behavior. For example, credit policy had imposed quantity restrictions on the lending by banks and other financial institutions. As a consequence, households expecting future income increases might be liquidity constrained when attempting to borrow against this expectation to finance current consumption[1].

Liquidity constraints — whether imposed by the market or by government decree — have been cited frequently as sources of rejection of the forward-looking hypothesis of consumer behavior — also referred to as the permanent income or life cycle theory (Hayashi 1985, Zeldes 1989, Wilcox 1989, Deaton 1991). To the extent that households have been kept from acting optimally by government-imposed liquidity constraints, financial deregulation thus should allow them to behave in closer accordance with this hypothesis. In this paper, we investigate whether financial deregulation has brought the behavior of Norwegian households closer in line with the forward looking hypothesis. Our tools are derived from the procedure used by Campbell and Mankiw (1991) and others, which permits estimation of a parameter that can be interpreted as the percentage of the population that does not follow the forward looking hypothesis, but instead consume their disposable income as they receive it. We then test the hypothesis that this percentage has declined after the financial deregulation in Norway. Campbell and Mankiw (op.cit.) also look for changes over time as a result of general improvements in financial markets, but fail to find significant effects (except for the United Kingdom, where the change was in the wrong direction).

As argued by Wilcox (op.cit.) and others, liquidity constraints may be imposed by the market as well as government regulations. For example, lending criteria based on payment-to-income ratios make borrowing against future income difficult. Thus, it would not be surprising to find that a substantial portion of the population was liquidity constrained both before and after the deregulation. However, if the proportion of liquidity-constrained households did not decrease, that would indicate that the government-imposed liquidity constraints were not seriously binding.

Liquidity constraints are not the only conceivable reason for departures from forward looking behavior. Another alternative, somewhat erroneously referred to as myopia, arises if households constrain themselves from spending income until it arrives. Rather than devoting time and energy to optimization, people may find it convenient to organize their decisions around «mental accounts» as a guide to which money is to be spent on what and when (e.g. Sheffrin and Thaler 1988). If such behavior is the primary alternative to the forward looking hypothesis, there is no particular reason to expect financial deregulation to alter household saving behavior. Such a pattern of behavior by Norwegian consumers is suggested in a recent paper by Lønning and Mork (1994). It is worth noting as well that a previous study of Norwegian consumers by Mork and Smith (1989), based on panel data from well before the deregulation, fails to detect significant deviations from the forward looking hypothesis.

Steffensen (1989) has made a previous attempt to estimate a model similar to ours on annual Norwegian data. However, because his data series included very few observations after the deregulation, he had no opportunity to study the effects of deregulation. Other time-series

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[1] In fact, one of the authors of this paper once had the personal experience of having a loan application denied for the expressed reason that his bank had reached the government's credit ceiling for the relevant year.

studies of aggregate Norwegian consumption (c.f. Brodin 1988, Brodin and Nymoen 1988, and Brubakk 1994) have been based on a different theoretical framework and thus cannot be compared to our work. We note, however, that Cappelen (1991)[2] and others have made less rigorous claims to the effect that the unexpected movements in aggregate consumption in the period between 1984 and 1987 were due to the households' newfound opportunities to borrow.

Norwegian consumption expenditure did indeed increase much more steeply than disposable income between 1984 and 1987. However, a large portion of this increase came from the purchases of consumer durables. Unfortunately, durables purchases are much more difficult to analyze than spending on nondurable goods and services because the consumption of durable goods cannot be equated with expenditure. For this reason, and the fact that nondurable goods constitute 75% of total Norwegian consumption, we follow most of the previous studies in the consumption literature by analyzing nondurable goods and services only. We feel, however, that a study of consumer durables should be a natural next step.

Our paper is organized as follows: Sections 2 and 3 present our theoretical model and its empirical implementation, including data and estimation methods. Section 4 presents and interprets the results, and section 5 gives some conclusions.

2. Theory and Model

We begin this section by briefly outlining the linear-quadratic version of the forward looking hypothesis of consumer behavior with rational expectations used by Hall (1978) and Flavin (1981) as a point of departure for the presentation of the alternative model introduced by Campbell and Mankiw (1991).

The Forward Looking Hypothesis
The well known idea of the forward looking hypothesis is that the representative consumer bases his or her choice between consumption and saving not only on current income, but also on the prospect of future income. Formally, the consumer seeks to maximise the following intertemporal optimization problem under uncertainty:

$$(1) \quad \max U = E_t \sum_{i=0}^{\infty} (1+\theta)^{-i} u(C_{t+i})$$

subject to

$$(2) \quad A_{t+1} = (1+R)A_t + YL_t - C_t$$

and

$$(3) \quad \lim_{i \to \infty} E_t \left[A_{t+i} / (1+R)^i \right] = 0$$

In equation (1), C_t denotes consumption in period t, $U(\cdot)$ a time-additive utility function, θ the subjective discount rate, assumed constant, and E_t denotes the expectations conditional on the information available at time t. Hence, the consumer maximizes the discounted sum of expected

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[2] Cappelen, Å: *Kredittrasjonering og konsumentadferd*, in Husholdningenes sparing, Begrepsavklaring, dataproblemer og analyse (ed K. Moum), Reports 91/16, Statistics Norway.

future utility as of time t. This maximization is done subject to equation (2) and (3), which describe the evolution of non-human wealth A_t over time and the so-called No-Ponzi Game Condition, respectively. Here YL_t is labor income, and R is the constant interest rate. The latter constraint prevents the consumer from borrowing to finance an increase in consumption today and then borrowing forever to pay the interest on the debt.

The first order condition or the Euler equation for this intertemporal optimization problem is given by

(4) $U'(C_t) = (1 + \theta)^{-1}(1 + R)E_t[U'(C_{t+1})]$.

Following Hall (op.cit.), by assuming quadratic utility and $\theta = R$, then

(5) $C_t = E_t C_{t+1}$.

Equation (5) implies that consumption follows a random walk. Flavin (op.cit.) elaborates an expression for the permanent income consumption function by using the result in equation (5) together with the constraints in equation (2) and (3), to obtain

(6) $C_t = RA_t + R(1 + R)^{-1} \sum_{i=0}^{\infty} [1/(1 + R)]^i E_t YL_{t+i} \equiv PI_t$.

Equation (6) contains the key implications of the linear-quadratic version of the forward looking hypothesis. It essentially says that optimal consumption is equal to the sum of the proceeds from non-human wealth and the expected present value of future labor income. Permanent income PI_t is defined to equal the right hand side of equation (6).

Finally, equation (6) and (2) imply that[3]

(7) $\Delta C_t = R(1 + R)^{-1} \sum_{i=0}^{\infty} [1/(1 + R)]^i (E_t - E_{t-1})YL_{t+i} \equiv \Delta PI_t$.

The change in consumption equals the change in permanent income, which equals the discounted value of the revisions in expectations about current and future labor income from last period to the present one. This result is intuitive because, according to equation (5), the consumer never makes any plans to undertake any future changes in consumption. Any change that may take place, must instead be rooted in new and unexpected information about how much the consumer can afford. Thus, it is unexpected news about labor income — revisions in expectations about future labor income — that lead to changes in consumption. Furthermore, when the consumption level is adjusted due to the advent of new information, the adjustment must be made in such a way that the new consumption level satisfies the new and revised budget constraint. Note also that the permanent income of course is unobservable as it contains expectations about future labor income.

This modern version of the forward looking hypothesis relies heavily on the assumptions of quadratic utility and perfect capital markets in the sense of absence of liquidity constraints. As we have seen, the assumption of quadratic utility has the great advantage that it allows us to derive a closed-form solution to the consumer's optimal consumption level. Furthermore, this solution contains the familiar property of certainty equivalence. However, the quadratic utility function also has an important awkward implication. Both common measures of risk aversion,

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[3] Equation (9) in Campbell and Mankiw's (1991) paper contains a typo.

that is the absolute and the relative rates of risk aversion, are increasing in consumption, indicating that the consumer becomes more risk averse the better off he or she is. This is counterintuitive from a perspective of realism. The other assumption that the consumer has access to perfect capital markets, with the possibility to borrow financial liquids, permits consumption to move freely with permanent income. However, Wilcox (1989) argues that lending criteria based on payment-to-income ratios constrain consumer borrowing, so that the consumer is often prevented from consuming as much as his or her permanent income justifies[4]. Thus, given the established empirical findings about the impact of liquidity constraints on consumption, it seems rather unrealistic to assume perfect capital markets when testing the permanent income hypothesis.

An Alternative Model

In the remainder of this section we briefly present the alternative model by Campbell and Mankiw (op.cit.), which attempts to counter the empirical critiques discussed above. Their model assumes that there is a probability λ that a given household does not follow the forward looking hypothesis, but instead lets its consumption change from period to period by the same amount as the change in its disposable income, ΔY_t[5]. Then, in the aggregate, we expect the change in consumption per capita to be a weighted average of ΔY_t and the expression on the right of equation (7), with weights λ and $(1-\lambda)$, respectively:

$$(8) \qquad \Delta C_t = \lambda \Delta Y_t + (1 - \lambda) \Delta PI_t .$$

An important advantage of this model is that it takes into account the possibility of credit rationed consumers resulting from financial regulation in the consumer markets. One can think of current income consumers as being credit rationed as they are prevented from consuming their permanent income[6]. As a consequence, a significant, positive λ found from estimation of equation (8) would then indicate a contradiction to the forward looking hypothesis due to imperfect capital markets. It is, however, important to emphazise that the interpretation of the coefficient λ in terms of the fraction of current income consumers may not be exact empirically. For instance, the coefficient λ may be understimated in those cases where the portion of credit rationed consumers also consists of *partly* rationed consumers. Such consumers are only able to follow the forward looking hypothesis until a prevailing credit ceiling is reached. The coefficient λ may be overestimated in those cases where current income consumers feel uncertainty about future credit rationing. In response to this uncertainty they may find it reasonable to limit their present loans somewhat (Jappelli and Pagano 1989). These potential effects on λ are not, however, measureable when estimating equation (8). As an empirical approximation we thus assume that any overestimation and underestimation of λ offset each other.

Campbell and Mankiw (op.cit.) reformulate the linear model in equation (8) by assuming that the representative consumer's preferences can be described by the following CRRA-utility function rather than quadratic utility:

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[4] Liquidity constraints, as a results of such lending policies, have also been analyzed rather extensively by other authors, for example by Zeldes (1989) and more recently by Deaton (1991). These authors conclude that the forward looking hypothesis can be rejected for a proportion of the population regarded to be rationed in the credit markets.

[5] Note that disposable income is defined as follows: $Y_t \equiv RA_t + YL_t$, that is the sum of the proceeds from non-human wealth and labor income.

[6] Of course, a fully worked out model with liquidity constraints predicts more complicated consumer behaviour (see for example Deaton 1991). As emphazised by Campbell and Mankiw (1991), their model can only serve as an approximation to a model in which liquidity constraints are explicitly modelled.

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(9) $u(C_t) = [1 / (1 - \gamma)]C_t^{1-\gamma}$, for $\gamma > 0, \gamma \neq 1$
          $= \ln C_t$, for $\gamma = 1$,

where γ is the coefficient of relative risk aversion, and $\sigma = 1/\gamma$ represents the elasticity of intertemporal substitution. The first order condition for optimal consumption choice now turns out to be

(10) $C_t^{-\gamma} = (1 + \theta)^{-1} E_t \left[(1 + R_{t+1})(C_{t+1})^{-\gamma} \right]$.

Note that equation (10) no longer assumes a constant interest rate. Campbell and Mankiw (op.cit.) make use of the assumption that consumption and the real interest rate are jointly lognormal and homoscedastic to show that this Euler equation simplifies to

(11) $E_{t-1} \Delta c_t = \phi + \sigma E_{t-1} r_t$,

or equivalently

(12) $\Delta c_t = \phi + \sigma r_t + \varepsilon_t$,

where the constant term ϕ reflects precautionary saving in response to uncertainty and lower case letters indicate the logs of variables. Note in particular that $r_t = \log(1 + R_t)$. ε_t is a normally distributed innovation term reflecting the new information that becomes available between times t-1 and t. Thus, $E_{t-1}\varepsilon_t = 0$. When equation (12) is weighted similarly as equation (7), we obtain the following log-linear version of the alternative model with stochastic real interest rate:

(13) $\Delta c_t = (1 - \lambda)\phi + \lambda \Delta y_t + (1 - \lambda)\sigma r_t + e_t$,

where $e_t = (1-\lambda)\varepsilon_t$. Finally, if expected real interest rate is assumed constant, then equation (13) becomes

(14) $\Delta c_t = \kappa + \lambda \Delta y_t + e_t$,

which is analogous to the linear model in equation (8) above. The constant term $\kappa = (1 - \lambda)(\phi + \sigma r)$. Our task in this paper can be stated simply as testing whether the value of λ for Norwegian consumers decreased after the financial deregulation in the mid-1980s. Because the error term $e_t = (1 - \lambda)\varepsilon_t$ reflects the new information about income and other relevant factors that arrives between t-1 and t, it is correlated with the current growth in disposable income, Δy_t. Thus, for consistent estimation, equation (13) and (14) must be estimated by instrumental variables. We discuss this and related problems in the following section.

3. Empirical Implementation

As indicated in the introduction, we follow the majority of the consumption literature by looking only at the consumption of nondurable goods and services. We thus implicitly assume that the utility function is separable between the consumption of durable goods on the one hand and of nondurable goods and services on the other. We can then estimate an equation like (13)

with data for the consumption of nondurable goods and services. For these categories, it is furthermore appropriate to equate consumption with spending, so that we can use expenditure data for these categories. We should note, however, that the omission of durables is not trivial in our case because Norwegian households may have used their newfound financial freedom to renew their stocks of durable goods. The data suggest this possibility, and we feel the investigation of durable consumption, albeit more complicated, should be the natural next step in investigating the effects of the liberalization.

We use quarterly data. Seasonally adjusted data are available; however, the methods used for this adjustment (X11-ARIMA), while highly sophisticated and widely recognized, are known to be less reliable for short series than for long ones. We thus decided to follow Campbell and Mankiw (op.cit.) in their work with European data and work with fourth differences instead of first differences. The theory behind equations (13) and (14) has the same implications for fourth differences, except that the error term now obeys $E_{t-4}e_t = 0$, so that it may have serial correlation of up to the fourth order. We compensate for this potential problem by correcting the variance-covariance matrix of our Two-Stage Least Squares (2SLS) estimates according to the Generalized Method of Moments (GMM), as recommended by Newey and West (1987).

For the same reason, our instruments need to be lagged by four quarters. Moreover, the fact that the quarterly observations of the consumption levels really are time aggregates of the instantaneous consumption rates for which the theory actually is made, may introduce an additional element of first order serial correlation, as demonstrated by Working (1960). In order to avoid this problem, we lag our instruments by five quarters rather than four.

Preliminary investigations indicated that we — like many other researchers — unfortunately were unable to obtain informative estimates of the elasticity of intertemporal substitution, σ. The estimates tended to be negative and had large standard errors. We were both surprised and disappointed by this finding because one of the effects of financial deregulation should be expected to be a clearer consumption response to changes in the real interest rate. However, we had to conclude that our data did not have the power to detect such effects. We thus decided to treat the expected real interest rate as constant, so that we actually estimated

$$(15) \qquad \Delta_4 c_t = \kappa + \lambda \Delta_4 y_t + e_t,$$

where Δ_4 indicates the fourth difference. Both consumption and income are in per capita terms and expressed in 1991 prices.

Campbell and Mankiw (op. cit.) also experimented with a couple of other specifications, one where the current income growth is replaced by a weighted average of current and lagged income growths, and another where public consumption is analyzed as a substitute for private consumption. We tried these specifications, but as they did not seem to provide informative results we decided not to present them.

An important decision was how to split our sample to obtain subperiods before and after deregulation. Grønvik (1992) describes the deregulation process as consisting of the following steps:

- *Removal of the ceiling on deposit rates in January 1978. A milder declaration of the lending rate was introduced as a substitute in September 1980 and eventually abolished in September 1985.*
- *Liberalization of the bond market in several steps between 1982 and 1985, allowing competition among banks and other lending institutions in the household and business markets.*

- *Abolition of reserve requirements in 1987.*
- *Gradual deregulation of the foreign-exchange market, completed in July 1990.*

Reserve requirements and the foreign-exchange regulations do not normally imply liquidity constraints for households. The interest ceilings could if they resulted in capital shortages. But the most binding regulations seem to have been in the bond market. As the removal of these regulations was gradual, however, the historical record does not make a clearcut suggestion of a breaking point in our sample. To be on the safe side, we could choose to put the break towards the end of 1985. However, the competition seems to have started well before the last regulation was removed. Thus, for example, Brodin and Nymoen (op. cit.) and Brubakk (op.cit.) break their sample at the beginning of 1984 and 1985 respectively.

In this situation, we find it impossible to avoid a certain amount of arbitrariness. We have chosen to be a little more conservative than Brodin and Nymoen, but do not feel it necessary to defer the break until the end of 1985 because we then probably would ignore a good many data points from a regime that was pretty much deregulated already. As a compromise, we split our sample in the middle of 1984. However, careful sensitivity tests confirm that the exact date of the breaking point is not important.

In contrast, the choice of instrument list *is* important because the power of our tests depends on the correlation between the ability of the instruments to predict disposable income growth. We conducted a comprehensive search among a wide list of candidate variables. Based on this search, we arrived at two alternative lists, where all the variables were used in fourth-difference form and lagged five quarters:

IV_1:   Log real disposable income per capita ($\Delta_4 y_{t-5}$)

      Log real government expenditure per capita ($\Delta_4 g_{t-5}$)

      The unemployment rate ($\Delta_4 u_{t-5}$)

IV_2:   Log real disposable income per capita ($\Delta_4 y_{t-5}$)

      Log real government expenditure per capita ($\Delta_4 g_{t-5}$)

      The unemployment rate ($\Delta_4 u_{t-5}$)

      The nominal interest rate ($\Delta_4 r_{t-5}$)

      The log of the Oslo Stock Exchange Total Index ($\Delta_4 s_{t-5}$)

The results of the regression of $\Delta_4 y_t$ on these instruments are presented in the Appendix. These results indicate that the longer instrument list IV_2 does a better job at predicting income changes, both in terms of $\overline{R}^2$ and a Wald χ^2 test based on GMM-corrected standard errors. Nevertheless, the shorter list IV_1 gives lower standard errors in the instrumental-variable estimation (with GMM correction). Thus, we report the results of both lists but tend to prefer those based on the shorter list. The fact that our instruments can predict income growth with a relatively high $\overline{R}^2$, means that we are able to distinguish empirically between current and permanent income[7]. Hence, the coefficient λ is identifiable in Norwegian data.

The data for consumption, disposable income, and government expenditure were taken from the quarterly national accounts of Statistics Norway. These series were deflated by the consumer price index for nondurable goods and services taken from the same data source. Per-capita figures were obtained using interpolations of the annual population data of Statistics Norway.

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[7] Except for Japan, Campbell and Mankiw (op.cit.) were also able to predict income growth significantly, but with lower $\overline{R}^2$ figures. For the Unites States, the United Kingdom, Canada and France the adjusted R^2 statistic were found to be 9%, 11%, 12%, and 13% respectively.

The unemployment data come from Statistics Norway's regular labor market survey. The interest-rate data and the stock-exchange data were obtained from the Bank of Norway. Our interest rate is the average lending rate of Norwegian commercial and savings banks. The quarterly national accounts for Norway go back to the beginning of 1966. However, because of the use of fourth differences and the lagging of the instruments, 1968:2 is the effective starting date of our sample. The ending date is 1994:4.

4. Results[8]

Table 1 shows the results of the estimation of equation (15) with the two alternative lists of instruments, on the entire sample 1968:2–1994:4 as well as the subsamples 1968:2–1984:2 representing the strictly regulated regime and 1984:3–1994:4 representing the deregulated regime.

For the period as a whole, the results resemble those of Campbell and Mankiw (op. cit.). The only difference is that, for the same specification, Campbell and Mankiw obtain somewhat lower λ-values except on non-seasonally-adjusted data for the United Kingdom. This difference may have been caused by the financial regulations having been somewhat stricter for Norway, although it is a little hard to see why our results should be very different from those for Sweden, where they obtain a λ–value of 0.36, compared to our 0.6.

We subject our results to a battery of tests, the p-values of which are displayed in the last four columns. The model test is a test of the exogeneity of the overidentifying instruments, obtained as a Wald test in a regression of the IV residuals on the superfluous instruments. We find no indication of the instruments from either list being inappropriate with respect to exogeneity. The LM test checks for serial correlation of up to the fourth order in the residuals. As expected, this test indicates considerable serial correlation, which is one of the reasons why we use GMM-modified estimates of the variance-covariance matrix. Another reason is potential heteroskedasticity, whose presence seems to be confirmed by the low p-values of the White test. As we shall see, at least part of the heteroskedasticity is due to the higher variability of consumption changes after deregulation.

Because of the heteroskedasticity, the empirical standard error of the residual is not really an estimate of any well-defined parameter. We include it nevertheless as a rough measure of the average variability of the residuals in the sample period. The last column, labelled «Normality», reports the p-values of tests of the skewness and kurtosis of the residuals, with the values for kurtosis in parentheses. We find no significant departure from normality.

We performed conventional Chow tests of the stability of the estimates, with a breaking point between the second and third quarter of 1984. These tests did not lead to rejection and thus suggest no effect of the deregulation. However, this test is not quite appropriate here because the error terms are dirty. Thus, we decided to ignore this result and went on to estimate equation (15) separately on the two subsamples 1968:2–1984:2 and 1984:3–1994:4. The results are displayed in the last four rows of Table 1.

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[8] All estimations and statistical tests have been conducted by use of the econometric package RATS 4 (Doan 1992).

Table 1. Estimation equation: $\Delta_4 c_t = \kappa + \lambda \Delta_4 y_t + e_t$

Estimation period	IV-list	κ	λ	SEE	Model test	LM test	White test	Normality test
1968:2-1994:4	IV_1	0.007 (0.008)	0.603 (0.237)	0.027	0.457	0.000	0.035	0.414 (0.706)
	IV_2	0.007 (0.008)	0.559 (0.254)	0.027	0.633	0.000	0.033	0.520 (0.617)
1968:2-1984:2	IV_1	0.003 (0.007)	0.747 (0.256)	0.024	0.358	0.001	0.068	0.551 (0.822)
	IV_2	0.012 (0.007)	0.369 (0.266)	0.022	0.159	0.020	0.379	0.544 (0.270)
1984:3-1994:4	IV_1	0.015 (0.013)	0.021 (0.317)	0.035	0.177	0.000	0.058	0.543 (0.631)
	IV_2	0.015 (0.016)	0.032 (0.520)	0.035	0.379	0.000	0.061	0.560 (0.643)

Note: The lists of instrumental variables, IV_1 and IV_2 are as defined in the text. The third and fourth columns show IV estimates of κ and λ, with standard errors in parentheses. The fifth column reports the standard error of estimate. The sixth column gives the p-values for a Wald test of the model, where the IV residuals are regressed against the superfluous instruments, and the null hypothesis is that the coefficients of this regression be zero. The columns headed LM and White report significance levels associated with the Lagrange Multiplier test for no serial correlation up to 4th order and the White test for no heteroskedasticity. The last column gives significance levels for the hypotheses of no skewness and no excess kurtosis in the IV residuals, with the latter in parentheses. All standard errors and test statistics are Newey & West (1987) consistent with respect to heteroskedasticity and serial correlation.

These results are strikingly different from the ones for the full sample. For the first period, the λ-estimates are even larger than for the full period when we use the short instrument list. In fact, this estimate is not far from being significantly different from unity, whereas the estimate for the full sample, with the same instrument list, is significantly different from unity on the 5% level with a one-sided test. The estimate of 0.75, suggests that liquidity constraints affected the consumer behavior for a majority of the Norwegian consumers during the regulated regime. With the second instrument list, the estimate for the first subsample is somewhat lower than for the full sample. It still is positive, although not significantly so. We note, however, that the standard errors are slightly higher for the estimates with this instrument list.

The most striking results are obtained for the second subsample, however. For this period, the estimated λ-values are practically zero for both instrument lists. Although the standard errors are too large to permit bold conclusions, we find these estimates remarkable. They suggest that the deregulation of the Norwegian financial markets in the 1980s may have removed completely the cause of any deviation from rational, forward-looking behavior by Norwegian consumers.

This result becomes even more striking when we compare it to the corresponding estimates for the United States consumers, who presumably have lived within quite a liberal financial-market system for a long time. Despite this regime, Campbell and Mankiw (op.cit.) find a tightly estimated λ-value for the United States of 0.35. We suspect that the difference may be due to the discrepancies in the distribution of income for the two countries. Because the United States has a larger proportion of households with incomes far below the national mean, we expect that

the share of households that are constrained by market-imposed borrowing constraints also is larger. Because the incomes of Norwegian consumers are much more alike, the large majority of people now are able to obtain loans from the financial institutions when they need them.

We should add, however, that Norwegian lending institutions may have been particularly lax about lending during our sample period. In hindsight, it has become apparent that banks' risk management in the period immediately following deregulation was lacking, and that this laxness contributed to the severe banking crisis of the early 1990s. The negative aggregate saving rates for households in the mid-1980s[9] similarly have been considered a social problem that — it is hoped — will not be repeated. Thus, we will not be surprised if someone finds a somewhat larger λ-estimate on Norwegian data for the next decade. If anything, our λ-estimates for the second sample period emphasize the extreme change in behavior that took place among banks as well as consumers after deregulation. We expect the pains of the subsequent crisis to have modified this change.

In order to study the statistical significance of the changes in the estimates, we repeat the sample split with dummy variables. Thus, we replace equation (15) by

$$(16) \qquad \Delta_4 c_t = \kappa_0 + \Delta\kappa \cdot d_t + \lambda_0 \Delta_4 y_t + \Delta\lambda \cdot \Delta_4 y_t \cdot d_t + e_t,$$

where d_t is a variable whose values are zero for all the observations in the first subsample and one in the second. The estimation of this equation is numerically equivalent to the subsample-by-subsample estimation in Table 1 provided that we expand the instrument list to also include d_t times each of the original instruments. Implicitly, we then include d_t itself as an instrument, which makes the GMM-corrected variance-covariance matrix produce the same standard-error estimates as if we weighted the observations of the two subperiods by their respective residual standard errors[10]. For the same reason, the estimated standard errors for κ_0 and λ_0 in Table 2 become identical to those of the first-period estimates of κ and λ in Table 1, respectively. This feature is of some importance because of the higher residual variance in the second subsample, as can be seen from Table 1 (c.f. the SEE figures) and the F-statistics below for equality of the error variances.

$$\text{IV}_1\text{: } F = \left(\frac{\hat{\sigma}_2}{\hat{\sigma}_1}\right)^2 = \left(\frac{0.035}{0.024}\right)^2 = 2.13$$

$$\text{IV}_2\text{: } F = \left(\frac{\hat{\sigma}_2}{\hat{\sigma}_1}\right)^2 = \left(\frac{0.035}{0.022}\right)^2 = 2.53$$

$\hat{\sigma}_1$ and $\hat{\sigma}_2$ denote the standard error of estimate in the first and second subsample respectively. The 5% critical value with degrees of freedom 40 and 63 is 1.59, and we thus reject the hypothesis of equality of the error variances in the two subsamples for both lists of instrumental variables.

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[9] For the years 1984–1987 the household saving rates were –2.7%, –6.1%, –6.2%, and –2.4%, respectively.

[10] The key to seeing this point lies in the fact that the GMM correction estimates the fourth-order matrix $E(Z'ee'Z)^{-1}$, where Z is the matrix of instruments. This operation allows the variance to depend on the right-hand variables, which in our case means that the variance differs depending whether the dummy variable is 0 or 1.

Table 2. Estimation equation: $\Delta_4 c_t = \kappa_0 + \Delta\kappa \cdot d_t + \lambda_0 \Delta_4 y_t + \Delta\lambda \cdot \Delta_4 y_t \cdot d_t + e_t$
Estimation period: 1968:2–1994:4

IV-list	κ_0	$\Delta\kappa$	λ_0	$\Delta\lambda$	SEE	Model test	LM test	White test	Normality test
IV$_1$	0.003 (0.007)	0.012 (0.015)	0.747 (0.256)	−0.726 (0.407)	0.029	1.000	0.000	0.006	0.683 (0.149)
IV$_2$	0.012 (0.007)	0.003 (0.018)	0.369 (0.266)	−0.337 (0.584)	0.028	0.313	0.000	0.006	0.341 (0.034)
IV$_1$	0.010 (0.008)		0.494 (0.293)	−0.232 (0.338)	0.027	0.679	0.000	0.066	0.903 (0.210)
IV$_2$	0.014 (0.009)		0.311 (0.320)	−0.218 (0.259)	0.028	0.292	0.000	0.027	0.460 (0.056)

Note: The columns headed κ_0 and λ_0 show IV estimates of the pre-1984:3 coefficients, while $\Delta\kappa$ and $\Delta\lambda$ show IV estimates of the change in the coefficients from 1984:3 onwards. The figures in parentheses are GMM-corrected standard errors. The columns otherwise are defined as in Table 1.

The results of the dummy-variable estimation are presented as the first two rows of Table 2. For the short instrument list, the shift in λ has a t-value of −1.782. This value can be considered statistically significant because, for a one-sided test, its p-value is 0.037. For the longer instrument list, the shift in λ naturally is lower because the first-period estimate is lower as well. More importantly, the standard error of the shift is about 50% higher than for the shorter instrument list, so that the t-value is quite low. In addition to the larger standard error, we note the significant deviation from normal kurtosis for this instrument list. For these reasons, we are inclined to place a somewhat higher confidence in the IV$_1$ results, despite the fact that the longer instrument list predicts income growth slightly better.

The short instrument list also estimates a sizeable, positive shift in the intercept κ. Although this shift is insignificant statistically, we find it plausible because $\kappa = (1 - \lambda)(\phi + \sigma r)$ and λ is estimated to shift downward. Furthermore, the increase in the residual variance that seems to have followed deregulation (c.f. the SEE estimates in Table 1) would provide an increased incentive for precautionary saving, so that an increase in ϕ also would seem plausible[11]. This observation provides an additional reason for believing in the IV$_1$ results.

However, given the statistical insignificance of the shift in the intercept, it would seem reasonable to see if the shift in λ could be estimated more sharply if the intercept shift is forced to be zero. The results of this exercise are shown in the last two rows of Table 2. The instrument lists for these results are identical to the ones in the first two rows; in other words, the dummy d_t is used as an instrument even though, by itself, it no longer is a right-hand variable in the equation. The reason is to maintain the mimicking of weighted estimation in the GMM-corrected variance-covariance matrix.

These results do not seem as attractive. The standard errors for the shifts in λ do indeed decline, but so do the point estimates, so that the shift no longer is significant for the short instrument list. As expected, the change in the point estimate is less dramatic for the longer instrument list; and the standard error is cut in half for this case. Even so, however, the t-value remains below

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[11] One possible explanation for the increased uncertainty in terms of higher residual variance after the deregulation may be the dramatic increase in the unemployment since 1987. In the period between 1987 and 1993 the unemployment rate changed from 2% to nearly 6.5% in Norway.

unity. We conclude that we obtain clearer results by allowing the intercept to shift and that doing so is in closer accordance with theory.

5. Conclusions

The purpose of financial-market deregulation normally is to allow agents to act more in accordance with the theory of rational behavior. In this manner, the deregulation effort in Norway aimed at improved efficiency. Although the public debate mostly focused on the issue of allocating capital to those purposes where it is the most productive, household behavior in regard to consumption and saving was very much part of the picture. After ten years of deregulated markets, we ought to be able to see from aggregate time-series data whether households indeed changed their behavior.

Our results suggest that they did. Whereas Norwegian consumer behavior showed evidence of substantial departures from rational, forward-looking behavior before deregulation, this evidence has essentially disappeared after deregulation. This evidence is so much more remarkable in that hardly any other aggregate time-series data set, from any country, conforms this closely with the forward-looking hypothesis.

Even so, our results do not prove that Norwegian households now act more rationally. Rational, forward-looking theory requires not only that consumption changes be unpredictable and uncorrelated with expected income changes, but also that the consumption changes should be determined by news about current and future income, subject to the intertemporal budget constraint. All we have found, is that Norwegian consumption changes have become unpredictable and uncorrelated with expected income changes.

In fact, an interesting alternative interpretation of our findings is that, after deregulation, Norwegian consumers started to behave more erratically, simply because they — or banks lending to them — had not yet learned how to adapt to a deregulated environment. The spending spree that followed deregulation — with negative, aggregate saving rates — can easily be interpreted this way. And even if this spending spree was not erratic, it may have reflected a one-time adjustment rather than a permanent, new pattern. Either way, it is entirely possible that the first decade after deregulation, which is the period we have studied, may turn out to be atypical.

Even so, we find it interesting to see that deregulation seems to have had such a strong effect on consumer behavior during this decade. Furthermore, we find it important to note that we only have looked at the spending on nondurable goods and services. An investigation of durables consumption would be a natural next step.

Appendix

The following table summarizes the results of the regressions showing the fit between the income changes and the respective instrument lists.

Prediction of $\Delta_4 y_t$

Estimation period	1968:2–1994:4		1968:2–1984:2		1984:3–1994:4	
	IV_1	IV_2	IV_1	IV_2	IV_1	IV_2
Constant	0.006	0.009	0.010	0.025	0.004	0.007
	(0.005)	(0.004)	(0.008)	(0.009)	(0.006)	(0.005)
$\Delta_4 y_{t-5}$	0.284	0.245	0.315	0.381	0.270	0.223
	(0.110)	(0.099)	(0.149)	(0.142)	(0.168)	(0.154)
$\Delta_4 g_{t-5}$	0.186	0.206	0.119	−0.007	0.204	0.257
	(0.093)	(0.103)	(0.172)	(0.193)	(0.097)	(0.096)
$\Delta_4 u_{t-5}$	0.029	0.026	0.022	0.033	0.037	0.032
	(0.010)	(0.009)	(0.013)	(0.013)	(0.012)	(0.012)
$\Delta_4 s_{t-5}$		−0.012		−0.002		−0.026
		(0.012)		(0.022)		(0.012)
$\Delta_4 r_{t-5}$		−0.679		−2.321		−0.592
		(0.393)		(0.829)		(0.372)
$\bar{R}^2$	0.164	0.190	0.114	0.200	0.150	0.239
SEE	0.023	0.023	0.025	0.024	0.021	0.020
Wald χ^2 test	0.001	0.001	0.015	0.001	0.016	0.000
LM test	0.000	0.000	0.000	0.000	0.051	0.105
White test	0.869	0.781	0.635	0.555	0.426	0.701
Normality test	0.940	0.562	0.981	0.706	0.868	0.756
	(0.862)	(0.739)	(0.791)	(0.665)	(0.921)	(0.586)

Each column shows the results of a regression of the four-quarter change in log disposable income on a set of instruments. The first row shows the estimation period, the second the name of the instrument list, the next six rows the coefficient estimates with GMM-corrected standard errors in parentheses, and the last five rows show summary statistics. The figures in the rows of $\bar{R}^2$ and SEE are the adjusted R^2 statistics and the standard error of estimates respectively. The p-values of the Wald test of the hypothesis that all the slopes are zero are computed from the robust GMM estimate of the variance-covariance matrix. The row of the LM test shows the p-values of the Lagrange-multiplier test of no serial correlation up to the fourth order, whereas the White test gives similar numbers for the White test of heteroskedasticity. The last row gives the p-values for test that the skewness and (in parentheses) kurtosis of the estimated residuals are normal.

The variables are defined as follows:

$\Delta_4 y_{t-5}$:   The fourth difference of log real disposable income per capita, lagged five quarters.

$\Delta_4 g_{t-5}$:   The fourth difference of log real government spending, lagged five quarters.

$\Delta_4 u_{t-5}$:   The fourth difference of the unemployment rate, lagged five quarters.

$\Delta_4 s_{t-5}$:   The fourth difference of the log stock index, lagged five quarters.

$\Delta_4 r_{t-5}$:   The fourth difference of the nominal interest rate, lagged five quarters.

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