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SOME IMPLICATIONS AND TESTS OF THE PERMANENT-INCOME HYPOTHESIS AND RATIONAL CONSUMER BEHAVIOUR.

by

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CONTENTS:

INTRODUCTION p. 1.

ESSAY I:

STOCHASTIC PROPERTIES OF AGGREGATE CONSUMPTION BEHAVIOUR.
A theoretical analysis and empirical investigation of data for some OECD-countries. p. 28.

ESSAY II:

CHANGES IN HOUSE PRICES AND THE PROPENSITY TO CONSUME.
A theoretical analysis and empirical study of a set of Norwegian microeconomic data. p.103.

ESSAY III:

EXCESS SMOOTHNESS AND MEASUREMENT ERRORS IN NORWEGIAN QUARTERLY INCOME AND CONSUMPTION DATA. p.151.

ESSAY IV:

SEASONAL IMPLICATIONS OF THE PERMANENT-INCOME HYPOTHESIS.
A comparison of Norwegian and Swedish data for consumption and retail sales.

Co-author: Knut Anton Mork. p.210.

INTRODUCTION.

1. Introduction.

What determines the consumption behaviour of individuals and households? How do they change this behaviour in relation to variations in economic policy and other central economic variables and conditions? Should consumption best be modelled as a fraction of current income or can we do better by applying more elaborate models? Questions like these are essential when we want to understand macroeconomic performance. Although much effort has been made to answer these questions, the lack of consensus in this field is still considerable. The questions are subject to debate and essential to different economic philosophies.

In this thesis, we investigate different aspects of the Permanent-Income Hypothesis. We elaborate on implications of this theory and derive testable hypotheses which we then test using mainly Norwegian data. First, however, we would like to highlight some important contributions that have led to the models that we analyse in this work.

2. Some history.

In the early days of economic thought, theory of aggregate consumption was more scattered and unstructured. The mercantilists had been concerned with the problem of under-consumption and insufficient demand. Adam Smith and the classical economists refuted this and claimed that the decision to save was linked to an investment decision so that aggregate demand would equal supply. Later, the theory of consumption was mainly concerned with the impact of the interest rate. Saving was assumed to be an increasing function of the real interest rate, r. This conclusion followed when we considered the household's intertemporal consumption decision as in the simple two-periodic model of the consumer introduced by Irving Fisher (1930). Although this model linked current consumption to both the interest rate r and life time wealth and resources, the main focus was on the interest rate.

2.1 Keynes' consumption function.

John Maynard Keynes (1936) devoted the third book (pp. 87 - 131) of his famous volume "The General Theory ..." to a discussion of the consumption function.

Compared to previous writers, Keynes was much less concerned with the intertemporal aspects of the consumer decision. Basically, he linked private consumption to the disposable income of households. Current consumption was assumed to be mainly a function of current income. Keynes did discuss the influence of the interest rate on current consumption but he concluded that:

...the total effect of changes in the rate of interest on the readiness to spend on present consumption is complex and uncertain, being dependent on conflicting tendencies,... There are not many people that will alter their way of living because the rate of interest has fallen from 5 to 4 per cent.(pp. 93-94)

Clearly, Keynes left little room for the interest rate in the consumption function. Further, he discussed what he labeled "*Changes in expectations of the relation between the present and the future level of income*". The conclusion was *...it is a matter about which there is, as a rule, too much uncertainty for it to exert much influence. (p.95)*

On the whole, intertemporal aspects of the consumption function were considered secondary and unimportant. His view of the consumption function can be neatly summarised in the following quotation:

... from our knowledge of human nature and from the detailed facts of experience , is that men are disposed , as a rule and on the average, to increase their consumption as their income increases, but not by as much as the income increase. (p.97.)

The consumption function described by Keynes in his book obeyed the following assumptions.

$$C_t = C(Y_t), \qquad 0 < {dC_t}/{dY_t} \equiv MPC < 1, \qquad \text{and} \qquad MPC < APC \equiv {C_t}/{Y_t}\,.$$

where C was aggregate consumption, Y was aggregate private disposable income and, t was the time subscript. Thus, consumption should be an increasing function of disposable income. Some of the increase was saved so that the marginal propensity to consume (MPC) was less than 1. Further, as people got richer they tended to save an increasing share of their disposable income so that MPC was less than the average propensity to consume (APC). Most of book III was devoted to the factors behind this function and how they influenced its position and slope.

2.1.1 The empirical evidence.

The "Keynesian revolution" emerged at the same time as quantitative methods in economics became available and more common. Inspired by the success of empirical studies in physics and natural science that revealed the laws of nature, economists wanted to reveal the laws of economics. The National Accounts were developed and formalised giving numbers to the major variables in Keynes' analysis. Time series data were constructed for aggregate measures such as GDP, private consumption, disposable income and savings, and some disaggregate data were available as well. The methods and data allowed testing of the new theories and hypotheses that were suggested by Keynes.

The first studies seemed to confirm the testable implications of Keynes' theory. The marginal propensity to consume out of income was between 0 and 1 and seemed to be less than the average consumption propensity. One of the major concerns in the early debate was whether or not aggregate consumption would eventually fall short of income as this increased and lead to under-consumption and unemployment.

In 1942, Simon Kuznets published data for US aggregates from 1869. From these data it became evident that there was a difference between the way data behaved in the short and in the long run. Although short run data were consistent with the theory, long run data were not[1]. The average propensity to consume did not decline as income increased, as predicted by the theory. On the contrary, it seemed to remain rather constant. This led to several refinements of the Keynesian model.

James Duesenberry (1949) had two suggestions as to how to resolve these problems. First, he claimed that people cared only for their relative consumption level. They could only increase their level of satisfaction if they increased their consumption level relatively to others of their group. A group could be students, an ethnic group or a neighbourhood. This would explain why the consumption curve shifted upwards over time as the general income and consumption level increased. Consequently, we would have a constant average propensity to consume as the general income level increased. Households with relatively low income would tend to have a larger propensity to consume as they tried to keep up with the consumption levels of others, while rich households would consume less on the margin. The marginal propensity to consume was consequently less than 1 when we looked at cross section data. Aspects of this hypotheses were confirmed by data.

His second suggestion was that it was more easy to increase consumption when income increased than to reduce the consumption level in bad times. The relative income hypothesis said that today's consumption was related to today's income relatively to previous peak income level. When income per capita was reduced it would seem that the marginal propensity to consume was small until we reach the previous peak income level. In particular, this was relevant for the period of the great depression. Many previous studies of the consumption function were heavily influenced by the data of this period and the hypothesis was successful in explaining the observed data. Duesenberry's two hypotheses explained much of the evidence found in previous studies.

2.2 Milton Friedman and the Permanent-Income Hypothesis.

Duesenberry reintroduced temporal aspects into the consumption function although in a somewhat ad hoc fashion. Milton Friedman made the consumption function a result of a fully dynamic optimisation behaviour of households in his Permanent-Income Hypothesis (PIH). In 1957, he published "A theory of the Consumption Function". In this book he challenged the Keynesian view of the consumption function. He returned to the classical economists prior to Keynes and emphasised the intertemporal aspects of

[1] Keynes was aware of the possibility that the marginal propensity to consume may differ in the short and

consumer behaviour which Keynes tended to disregard. According to the Permanent-Income Hypothesis, households scaled their consumption level to their life-time resources. Temporary changes in income should have no influence on the consumption level.

More formally, he divided both consumption and income into permanent and transitory components. In Friedmans own notation we had that

$c_p = k \cdot y_p$,

$y = y_p + y_t$ and

$c = c_p + c_t$

Subscripts p and t indicated permanent and transitory components of the variables. The propensity to consume (k) of permanent income depended on the interest rate, the ratio of non-human wealth to income[2] and consumers' preferences between wealth and consumption. He concluded that from data, k appeared to be rather constant over time.

Friedman was very unclear when he discussed the notion of permanent income (pages 23 - 25). He did not adapt the infinite horizon assumption that has been so tightly linked to the hypothesis in later studies and presentations. Further, he was vague on how this related to expected life time income and he appeared to have an open mind about the possibility that households may have quite short horizons. This vagueness seemed to be a deliberate intention. In later works, Friedman (1963) was more in line with what we consider to be the permanent-income concept of today. There he treated permanent income as a factor r of the human and non-human wealth in all future periods. Wealth was of course the present value of current assets and discounted values of future income; r was the rate of discount. Friedman related this partly to the subjective horizon

in the long run but concluded that the overall conclusion would apply also in the long run. (p.97).

[2] In our days, we would usually interpret non-human wealth into the measure of permanent income.

of the household and partly to the prevailing interest rate. The discount rate was assumed to be time dependent.

Crucial to Friedman's (1957) analysis were the correlation assumptions that he posed on his model. He assumed that the correlation coefficients between the various consumption and income measures were

$$\rho_{y_t,y_p} = \rho_{c_t,c_p} = \rho_{y_t,c_p} = 0$$

The two first assumptions seemed to be almost definitional and rather unimportant. The final restriction was what distinguishes Friedmans approach from earlier contributions. He assumed that transitory income and consumption were uncorrelated. The assumption was imposed on the system and not actually tested. Friedman was fully aware that he was not strictly testing his model against alternatives, and was mainly concerned with whether data were consistent with his hypothesis or not. Much of the discussion was meant to interpret previous results in the light of his hypothesis. The main testing problem was of course that the permanent and transitory components of consumption and income were unobservable and must be imputed from the data.

In his perhaps most well-known econometric model, Friedman assumed that the permanent part of income could be represented by the sum of a general trend and a geometrical lag structure in income deviations from the trend. This allowed him to estimate a measure of permanent income. Based on this measure he estimated that the propensity to consume of permanent income was 0.88. Transitory consumption constituted the residual of the regression. However, he performed no tests of his basic hypothesis that the transitory components should be uncorrelated. And as he wrote himself:

But, as is always the case in empirical work, there must be numerous other hypotheses with which this same evidence would be consistent; insofar as we choose ours, it is because we regard it as simpler and more fruitful than others....(p.157)

Obviously, others must have shared this opinion. Although, the empirical foundations for the Permanent-Income Hypothesis are vague and strict testing has been scarce, the hypothesis has shown itself to be more long-lived than most other hypotheses in economics.

The PIH explained much of the observed behaviour of the data and was, as Friedman said, consistent with the evidence. When transitory income was positive, the general income level tended to be high and the propensity to consume low since the transitory part of income had no impact on the consumption level. When it was low, the reverse was true, explaining the shape of the Keynesian short run consumption function. In the long run consumption would track income and the average propensity to consume was likely to stay rather constant. The hypothesis was able to explain smoothness of aggregate consumption.

2.3 The Life-Cycle model.

Brumberg and Modigliani (1954) developed a model that for aggregate data shared much of the same properties as the PIH. The difference between the two hypotheses was mainly in their focus. The Life-Cycle model focused on individual household behaviour as dependent on age while the Permanent-Income Hypothesis disregarded to some extent individual behaviour in favour of the average household. This was natural since Friedman's prime interest was the consumption function and aggregate behaviour. However, the implications for the aggregate economy showed more similarities than differences between the two hypotheses and they were often considered twin hypotheses. Often models on aggregate data were considered to represent the two hypotheses jointly.

The basic idea of Brumberg and Modigliani was that personal income was unevenly distributed over a person's life time. In the beginning income rose while we expected a decline in later years when people tended to have small or no income. On the other hand, people are likely to smooth their consumption level. Typically they would want to save for old age in order to keep up the consumption level when the income level droped.

In this model, the average propensity to consume would stay constant over time, since new generations would scale their consumption level to their life time income which in general would be higher than for previous generations. Aggregate consumption would follow aggregate income over time. A lower marginal propensity to consume would arise partly because people were at different stages of life (cross section data) since people would tend to save in high-income periods and partly because they would want to smooth transitory income (time series data).

The life-cycle model gave an easy answer to the question of saving. Aggregate saving in an economy was due to the fact that income tended to arise earlier in life than consumption. Also population growth and age composition tended to add to saving. Friedman was not very explicit about saving behaviour but it appeared that aggregate saving was explained by a wish to earn interest and perhaps by a precautionary savings motive. However, none of these were elaborated and included in the model.

2.3.1 From an empirical viewpoint.

In the two decades that followed, the twin hypotheses were central issues for economic debate. Refinements and modifications were introduced and the models were enriched. Consumption functions and other empirical studies were done based on the hypotheses.

However, it was difficult to discriminate between different models of consumer behaviour based on aggregate time series data. The problem was to go from observed data to construct measures of permanent income. Permanent income was closely related to the expectations that a household may have about future income. Depending on these assumptions several specifications of permanent income could be regarded as consistent with the theory. As a consequence, we have seen very little formal testing of this model versus other hypotheses of the consumption function.

The Life-Cycle model had more testable implications when we considered individual households and cross sectional data. The typical hump shape in savings with increased savings at early stages in life and dissaving in later parts, could be investigated. In general, studies indicated that dissaving in old age was small and inconsistent with the theory. Even if we add uncertainty about life length, the saving of older people indicated some altruistic behaviour and a bequest motive. The extent of this behaviour is still a matter of debate.

Some empirical studies considered more partial implications of the two theories. For instance, Cagan (1965) showed that people with pension plans tend to save more rather than less. Another example was Lee (1975). He treated dividends given to war veterans as windfall income[3]. He showed that the recipients had a clearly positive propensity to consume out of this income beyond the predictions from the two hypotheses.

Even though some parts of individual behaviour seemed to conflict with the theory, it remained an unsettled question whether aggregate data were in accordance with the

[3] Friedman (1957) stated these dividends as an example of windfall income.

Permanent-Income Hypothesis or not. The rather weak formulation of the notion of permanent income was an obstacle to implementing the model empirically. The main problem was how to represent the expectations that households held about future income. Often rather arbitrary assumptions were used. The introduction of rational expectations solved many of these problems.

3. Hall combines Rational Expectations and the Permanent-Income Hypothesis.

Hall's[4] (1978) inclusion of Rational Expectations on behalf of the households turned the study of the Permanent-Income Hypothesis (PIH) in a new direction. In order to consume according to permanent income, households must hold some expectations about their future income prospects. According to the Rational-Expectations approach, households should form their expectations so that there were no systematic biases in their predictions of future income. Households should make their best guesses based on all available information.

Some might argue that the assumptions of an infinite horizon and complete rational behaviour would violate the original ideas of Friedman. In his book Friedman had a much more modest set of assumptions and discussed both uncertainty and short-sightedness on behalf of consumers. On the other hand, the assumption of rational behaviour seems to comply with the inherit logic of the PIH and takes it to its logical consequence. Thus, it seems to be a natural way to extend the hypothesis. After all, in the PIH households already seem to behave rationally in order to smooth income. Why should they form expectations that were biased and false?

[4] In a parallel work, Sargent (1978) combined the same hypotheses. However, as Flavin (1981) pointed out, Sargent made some computational errors that made him deviate from Hall.

Hall was able to pin down several implications for consumer behaviour of the combined hypotheses. A theory that, from an empirical point of view, had been vague and difficult to track down, suddenly became well-structured and testable.

The major consequences of the Permanent-Income Hypothesis/ Rational Expectation approach (PIHRE) can be summarised in four properties. First, changes in permanent income beside those which are implicit from the household saving decisions must be unpredictable. This is the core of rational expectations. If permanent income is a best guess, it means that it is based on all available information about future income prospects. The expected permanent income in some future period can only change due to new information. As we assume that consumption is proportional to permanent income, deviations from the pre-planned consumption path should be unpredictable as well. The information we have today can not tell us how consumption will change in the future apart from the possible trend that households have planned for. Thus, adjusted for trend, consumption should be orthogonal to all lagged variables besides consumption itself.

Second, Hall combined the hypotheses with an assumption of quadratic utility so that marginal utility was linear in consumption and he set the rate of time preference equal to a certain interest rate. As a consequence, households should plan for a constant level of consumption equal to their permanent income. Then the evolution of household consumption should track permanent income and they should both become a random walk. Simple theory of non-stationary time series told us that the two properties mentioned should carry over to aggregate consumption.

Hall was primarily concerned with testing the unpredictability of consumption. His conclusion was that so was mainly the case. Changes in consumption seemed to be unpredictable from own past values and from household disposable income. When he tested for corporate stock values he found significant explanatory power all together, but concluded that the effect was small. In a later paper Flavin (1981) elaborated the model and showed that consumption was excessively sensitive to predictable changes in income and thus could be predicted from past information. Further, she showed the link between the time series properties of current income and household permanent income.

As a third implication, Deaton (1986) pointed out that permanent income did not need to be smooth. If income was non-stationary, income changes today might signal major changes in the future income path, and permanent income might perfectly well be more volatile than current income. Then the theory would have problems explaining why the marginal propensity to consume of current income should be less than the average propensity. Campbell and Deaton (1989) showed exactly that, and demonstrated that the smoothness of consumption should possibly be regarded as a puzzle to the PIH and not as an implication.

Finally, the fourth implication shown by Campbell (1987) is that it should be possible to predict future income by the size of consumption and saving. Households that expect future income to rise should have a lower level of saving than those who expect future income to decline. It should be possible to deduce household expectations about the future from their consumption/saving decision today.

Hall's model was actually a model of a single household but the studies mentioned above were all on aggregate data. In some cases the validity of the model on aggregate data was far from obvious. Hall and Mishkin (1982) investigated household data for USA. They made use of the simple covariance structure of income and consumption and showed that household consumption was sensitive to transitory income although to a much less extent than to changes in permanent income. Mork and Smith (1989) did a similar study on Norwegian data. Their data were of the same kind that we apply in our third essay, but from an earlier period. Their results were unfavourable to the PIHRE model.

In this thesis several of these implications are investigated in greater detail.

4. Outline of the thesis.

This thesis is completely within the tradition of the Permanent-Income and Life-Cycle Hypotheses. The three first papers elaborate different implications of the rational expectations assumption. In the first and the third papers we compare the Permanent-Income Hypothesis with the Life-Cycle Hypothesis and try to point to some differences in aggregate data behaviour. In the second paper the difference between the two is implicit and dependent on the assumed time horizon. The last paper discusses the smoothness implication common to both hypotheses. Here, we do not need the assumption of rational expectations. Households should smooth consumption relatively to expected fluctuations in income regardless of how they form their expectations.

All the papers contain empirical tests of these hypotheses. However, the papers differ in how they approach the hypotheses and which implications they test, as well as in the type of data that is being used.

In the first paper, «Stochastic Properties of Aggregate Consumption Expenditure» , we aim at studying the unit root properties of aggregate consumption. The maintained assumptions of the PIHRE model imply that aggregate consumption shall be close to a random walk. Issues such as existence of durable goods, aggregation bias or possibly slow reactions on behalf of households, may relax this. But these are mainly short run distortions and should not inflict with the long run unit root property of the data. In the paper we show that a Life-Cycle Hypothesis is likely to generate a consumption path that is very close to a unit root as well. From an econometrical point of view, the path should be indistinguishable from such a data-generating process. The two hypotheses do differ, however, in the general correlation pattern. The Life-Cycle model implies that changes in aggregate consumption should show positive correlation beyond the PIHRE model.

We test these assumptions on relatively long data series for seven OECD-countries. It is well-known from simulation studies that in order to obtain a more powerful test of an unit root hypothesis, the time span rather than the number of observations is what seems to be important.

The unit root hypothesis does not seem to perform well. In most data series the estimated unit root parameter is far from unity, but the Dickey-Fuller test observers are on the borderline of being significant due to high standard errors. In a Baysian test developed by DeJong and Whiteman the hypothesis is rejected in all cases. The paper compares the two tests. They both suffer from deficiencies. The main objection to the DF-test is that it is likely to have very low power, increasing the risk of accepting the unit root hypothesis although this is not true. The problem is due to the fact that the test utilises the sum of the estimated autoregressive parameters to represent the persistence in the series, rather than an exact measure of the lowest root of the process. If the "true" sum of the parameters is exactly 1, this implies that the lowest root of the estimated process is at the unit circle as well. But for the case where the lowest root of the estimated regression is different from 1 the sum is an imprecise measure of this root. As we show, the discrepancies between the two measures are possibly quite large. Thus, for a given significance level, we need wider confidence limits in order not to reject the null hypothesis of a unit root.

The Baysian test is problematic, since it seems to disregard that the asymptotic distribution of the estimated parameters changes discontinuously as the process goes from a stationary to a unit root process. As Phillips (1991) points out, this calls for a discontinuous confidence area.

In the second paper, «Changes in House Prices and the Propensity to Consume», we discuss the effects of a price boom and the following slump in the housing market and how these movements affected the consumption pattern of households. A theoretical model based on rational behaviour is outlined and implemented on a set of micro-data for Norwegian households. One of the main aspects of the model is that we should expect households to respond quite differently to price changes dependent on their status in the housing market. House-owners are likely to have a non-negative propensity to consume from house price changes, the size depending on the horizon that the household might have. The shorter the horizon the larger the propensity to consume since the wealth gain from the price increase is to a less extent offset by increases in future housing costs. The tenant should respond negatively to price changes since price increases are likely to change his future housing costs. Further, the more distant the horizon, the more negative the response is likely to be. We obtain this result for a certain specification of the utility structure of the households, but the implication seems obvious and should be far more general. The effect is mainly related to the budget constraint and should emerge from almost any reasonable utility structure.

In order to obtain this result, we apply an equilibrium model in the housing market that relates buying prices to renting prices in the housing market. The model of the housing market is based on rational behaviour from households when they make the decision of owning versus renting a house. As such, it seems a natural part of a model that tests rational behaviour.

For the purpose of this study, we have constructed data for Norwegian households for the period from 1986 to 1989. This was a period of large fluctuations in house prices, peaking in 1987 and the price changes led to considerable redistributions of wealth among households. Our results are not supportive to the basic assumption of rational behaviour. First, we are not able to distinguish the behaviour of the two groups of consumers. Tenants appear to have a large positive propensity to consume from house price changes; at least as high as for house-owners which is in strong contradiction to a model of rational[5] household behaviour. However, the results are on the borderline of significance, making it hard to obtain solid conclusions. A problem with the data set is that the time dimension is small, making it impossible to obtain a dynamic specification of the model. The instability of the results across different years suggests such an

[5] When using the term «rational», we think in standard economic terms throughout the thesis. Thus, irrational behaviour in this sense can perfectly well be rational in some wider sense as for instance when we include costs of computation and information collecting.

explanation. However, we also find it difficult to make up a credible dynamic story that would explain the results we observe. To the extent that we can draw conclusions from our results, they are highly unfavourable to an assumption that households relate to house prices as a determinant of their consumable wealth.

The third paper, «Excess Smoothness and Measurement Errors in Norwegian Quarterly Income and Consumption Data», came about as a wish to investigate the excess smoothness property of Norwegian data. Campbell and Deaton show that if aggregate income is integrated of order 1, permanent income does not need to be smooth compared to current income. Actually, it appears that it is much more volatile. Thus, the PIHRE-model can not serve as an explanation for the smoothness of consumption compared to current income. The smoothness of consumption was probably one of the reasons that Friedman introduced the Permanent-Income Hypothesis. Given the maintained assumption of disposable labour income as I(1), consumption in quarterly US data is certainly smoother than the PIHRE model predicts. We suspected that Norwegian data would show a different pattern of a more volatile consumption path. And this turned out to be quite true. However, instead of finding evidence to contradict Campbell and Deaton, we believe that we have found short run information in Norwegian data to be of poor quality. Norwegian quarterly consumption data appear to suffer from measurement errors.

From a theoretical point of view, we show that excess smoothness follows from a life-cycle model as well. The model used is the same as in Galí (1990). We show the extent to which this smoothness depends on essential parameters in our model. We suggest that this is likely to be a partial explanation for the results that Campbell and Deaton obtain. We argue that the parameter values applied by Galí are probably too small and tend to marginalize the effect of excess smoothness in this model. By empirical testing, we show that in Norwegian quarterly data the reverse conclusion emerges. Consumption appears to be more volatile than permanent income. However, the behaviour of the data strongly suggests that this is an artefact due to measurement errors. The quarterly data show an odd behaviour, for which we find no other plausible explanation but such errors. Primarily, changes in the consumption level are strongly negatively correlated at lag 1. If we turn to annual data, the Norwegian results resemble the ones for US annual data and confirm the excess smoothness property of consumption. The changes in consumption are positively correlated over time and more congruent with results from other countries. The same problems of measurement errors seem to apply for income data as well, but perhaps to a somewhat smaller extent. This gives a negative bias in the AR(1)-parameter which tends to understate the permanence of an innovation. Thus, its permanent-income equivalent will be reduced. It follows that we have more confidence in the annual data and base our conclusions on that.

We do not actually try to pin down the reasons why the data contain errors. A possible explanation could be the seasonal adjustment procedure. The data have been seasonally adjusted by the ARIMA-X11 method. However, our prime suspects are errors in the retail sales index and errors due to the conversion process that makes quarterly data consistent with the annual data. Imposing long run consistency on the data may have detrimental effects on the short run information that they contain.

The last paper «Seasonal Implications of the Permanent Income Hypothesis» is a joint work with professor Knut Anton Mork. In this paper we investigate the smoothing property from a somewhat different angle. If people behave according to the permanent-income/life-cycle theory, expected seasonal variations in income ought to have no bearing on the way households distribute their consumption expenditure over the seasons. This paper was actually written before the others. It differs from the three first papers in that it does not rely on the assumption of rational behaviour. The original and weaker assumptions of Friedman are sufficient in this paper. Note further, that this smoothing property by no means is affected by non-stationarity in income data, in the way it happens for Campbell and Deaton, since we consider smoothing over the seasons and not generally over time.

Further, the problem of measurement errors that we found in the third paper is likely not to cause a serious problem in this paper, since the overall seasonal pattern is a long run property of the data and only vaguely related to short run errors. As long as the errors do not show a seasonal pattern, for which we see no reason, errors should cancel out over time.

We cannot exclude the possibility of a seasonal pattern in preferences so we need to compare two groups of people that have a different seasonal pattern of income, but which are similar in most other characteristics. Using Sweden as a control group, we find that the seasonal pattern in Norwegian quarterly consumption is influenced by the specific tax withholding system that we have in Norway. Here, tax withholding is suspended during June or July and done at half rate in December. Compared to Sweden, Norwegian quarterly consumption data imply that Norwegians tend to allocate their consumption to the second, third and fourth quarters in proportion to the differences in tax withholding. The propensity to consume is approximately unity.

When we look behind the consumption data, we do find problems with the construction of the seasonal pattern in the data from the raw data. The construction of quarterly consumption data is based on the retail sales index (RSI) and the seasonal pattern in consumption data is totally dependent on the similar pattern in the raw data. We try to give some insight in the construction process of the RSI and how this is transformed into aggregate data. We highlight the differences between the Norwegian and Swedish construction methods and deduct the implications for the time series data.

For Norway, we find that the overall seasonal pattern is well preserved in the conversion of the retail sales index into aggregate consumption data. For Sweden, however, we find that the correspondence is less than satisfactory. We show that this is probably related to how Swedish data are constructed, but we are not able to track down the exact sources of the discrepancies. When our hypothesis is tested on the raw data, we find that the fit is less convincing for what we call the naive hypothesis although the evidence still points in the same direction as the previous conclusions.

5. Reflections on the PIHRE model.

The main conclusions from this thesis are that the evidence from the data is unsupportive to the Permanent-Income Hypothesis and in particular to the PIHRE-model. This is not a new discovery and the model has failed in several previous studies. Even though we apply different sets of data and to a large extent investigate different implications of the theory than previous studies, we derive the somewhat same conclusions. *Confronted with data, the PIHRE model does not perform in a satisfactory way!*

Among the problems are the excess smoothness/sensitivity issues raised by Flavin (1981) and Campbell and Deaton (1989). Aggregate consumption appears to respond slowly to aggregate income and track the short run evolution in income too closely. Several studies have found that a share of the aggregate consumption is better described by a Keynesian consumption function as in Campbell and Mankiw (1991). Further, consumption is sensitive to lagged information. As far as we have investigated this, Norwegian data are consistent with these conclusions.

However, some of our results indicate that there is a deviation from the theory also in the long run, as it appears that aggregate consumption is trend stationary and does not contain a unit root. If this is the case, then this is in some sense a stronger violation of the model implications than the ones described by Flavin and Campbell and Deaton.

Several explanations have been tried in order to save and modify the theory and settle the differences with the data. Problems such as aggregation bias, precautionary savings, various data generating structures and liquidity constraints among others, have all been suggested as solutions to the deficiencies of the model. Most of the reasons given in order to explain previous results tend to imply temporary deviations from the long run path and cannot explain why the data fail on the unit root test. Galí (1990) and ourselves have investigated the implications of a life-cycle model that in the limit implies that consumption tracks the path of aggregate income. However, as we show, the deviation from a unit root process is likely to be so small that we should probably not be able to discover it by the use of econometrical tests.

The obvious question is whether we can defend to go on using the PIHRE-model in empirical and theoretical work. If so, in what way should we modify the theory to comply with the observed behaviour of the data. Here, we will comment on an issue that has received little attention. The PIHRE-model is a partial model in the sense that it assumes that aggregate income and the interest rate are exogenous to the household and assumed known. A standard assumption is that the real rate of interest is assumed constant. For the individual household the exogeneity assumption appears to be reasonable. However, in the aggregate this assumption is less plausible. It disregards the fact that the real rate of interest might come as a result of the consumption and investment decisions made by the agents. If demand today is high compared to aggregate production this will tend to increase the price of consumption and investment today in relation to the future. The interest rate should then rise.

Let us consider a closed economy. We have the simple identity that total amount of goods and services produced can be used either for investments or for consumption purposes. But as this trivial relationship applies, a given theory for the consumption path and a fixed interest rate give severe restrictions to the income-investment relationship. In the most common and simple version of the PIHRE-model, we have that $C = Y^P$, where Y^P is the permanent component of income which is assumed to be the annuity of future income streams. If we disregard the public sector, this implies that investments are dictated by the transitory part of income, $I = C - Y = Y^T$[6]. First, this leaves no independent option for rational investment behaviour. The investment decision is not linked directly to the interest rate. A general increase in the real rate of interest will lower present consumption and thus tend to increase investments. This seems to

[6] In the crudest sense, this implies that investments on average are zero and often will be negative. This reflects the problems with explaining saving in the simple PIHRE. However, if we introduce a more sophisticated model that accounts for saving either through life-cycle effects or through precautionary motives, the above reasonning and critique will still apply.

contradict reasonable investment functions and behaviour. Further, to the extent that investments are linked to transitory income, it means that investments by definition cannot have a permanent impact on permanent income. This in turn means that the expected yield must be exactly $r \cdot I$. If the economy-wide signals are that future income will be higher than previously assumed, this will both trigger an increase in demand for consumer goods and be likely to increase investment demand. However, these are mutually incompatible and the two kinds of demand are likely to crowd each other out, the real interest rate being the likely instrument.

This is certainly an important aspect when we consider large regions/countries, and to the extent that innovations and shocks are world-wide. But even if income shocks are of a local nature, it is reasonable to assume that this effect will be of some importance. Only if a country is small, and foreign production are perfect substitutes for domestically products and services, will the restrictions on the consumption/investment allocation be totally alleviated, and the be independent of the amount of home-made goods and services.

Further, we consider data over the last century. In this period world goods and capital markets were much less integrated than they are today. We have observed trade barriers and direct capital control, and also we must assume that transaction costs were much higher during this period. This is particularly important when we consider the capital markets. Thus, during most of the period that we observe, the assumption of a closed economy is more accurate than during the last decades.

In some studies consumption is regressed on the interest rate. However, from the discussion above, it is far from obvious that interest rates can be treated as exogenous in such a regression. The causality is far from obvious. Christiano (1987) considers a general equilibrium model of a closed economy and makes simulations of this economy. He finds that in such a setting consumption and investments are smoothed and that this might explain the behaviour of the consumption data. On the other hand, he finds it problematic that investments are smooth since this is in contradiction with empirical evidence. Investments are known to be volatile.

We believe that this line of research is not fully exploited. The interaction of rational consumption/investment behaviour and imperfect trade should possibly give plausible descriptions of the data that we observe. Further, in such an economy it is likely that consumption will stick to the income path also in the long run. Thus, if income is stationary, the restrictions should probably make consumption stationary as well. Otherwise, consumption and investments must be co-integrated. This is certainly a project that I wish to pursue further in the future.

6. Concluding remarks.

The PIHRE-model introduced by Hall gave a much more precise and formal representation than previous specifications of the Permanent-Income Hypothesis. This was certainly needed in order to increase the applicability of the theory. The model is appealing since it is derived from "rational" behaviour by individual agents in the economy. It has strong implications on both micro and macro level and is useful in building economic models. The model should be considered a benchmark and a reference point for other models.

However, the empirical evidence so far is not overwhelming in its support. The evidence that we show in this work is consistent with that view. In the first paper, we find that consumption expenditure tends to be trend stationary in the long run, which contradicts the assumption of a long horizon and rational behaviour. In the next, we investigate the impact that wealth gains and losses in the housing market have on household consumption expenditure. In general, we are not able to observe rational consumption behaviour to the changes in the budget constraint. Then we turn to the excess smoothness problem of consumption. Based on annual data, we conclude that this problem seems to apply to Norwegian data as well. Finally, we show that the seasonal pattern of consumption expenditure seems to be related to the tax-withholding scheme that we have in Norway contradicting the predictions of the Permanent-Income Hypothesis.

It is important to find why real world data behave differently from the model predictions. There seems to be an abundance of possible explanations that can cover for the observed discrepancies. They all appear to be more or less consistent with the data and all together are amply sufficient to account for real world behaviour. Apparently, there is no agreement as to how this model should be modified in order to fit with reality. Based on aggregate data it is probably impossible to distinguish between the different explanations given. Thus, if we want to decide on the actual causes for the empirical results we need data that carry more information about actual household behaviour. Disaggregate data of good quality are needed.

It seems premature to make the final judgements about this model at this stage. With proper modifications, it still remains to be seen if we are able to make the theory operational so that it can make a better job at explaining existing observations than its rivals. It seems that a considerable effort is still needed on this model.

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ESSAY I.

STOCHASTIC PROPERTIES OF AGGREGATE CONSUMPTION EXPENDITURE.

A theoretical analysis and empirical investigation of data for some OECD-countries.

1 INTRODUCTION

The purpose of this paper[1] is to answer the question of whether aggregate consumption expenditure contains a unit root or not. The answer to this question may have important implications for economic understanding. Hall(1978) shows in a very influential paper that the consumption path of a household acting according to the permanent income hypothesis (PIH) and forming rational expectations (RE) about future income streams, should contain or at least be very close to a unit root process. A lack of such roots in an empirical analysis would question the validity of this combined hypothesis (PIHRE).

We compare this model with a model proposed by Galí (1990) in which he considers aggregate consumption based on life cycle behaviour of households. He assumes that households have finite lives and that they leave no bequest to their descendants. Aggregate consumption ought to track aggregate income more closely in this framework than in the PIHRE model. However, we show that even if income is trend stationary, we would expect consumption to be rather close to a unit root process. The two models seem indistinguishable with the use of unit root tests. However, they differ to some extent in the additional correlation of the data. The Life-Cycle model should contain a positive correlation structure beyond the model proposed by Hall.

[1]I would like to thank professor Knut A. Mork for valuable comments on earlier drafts of this paper.

The question of stationarity is of vital interest in economic modelling. If the series contain a unit root, statistical shocks or innovations may have permanent and everlasting effects. In traditional macroeconomics, variables are expected to follow a deterministic trend. This was presented as a linear trend (or linear in logarithms) with possibly one or two breaks in the growth rate. Any deviation from this trend was considered temporary. Thus, a unit root was excluded in advance.

We approach the questions with data for seven OECD countries. Since unit roots have fundamental implications for economic modelling and understanding, we would like to be able to make conclusive statements of whether private consumption does or does not contain a unit root more or less independently of time and space. However, our results seem to diverge considerably between countries. When we look at the data after World War II, the data seem to fit the unit root hypothesis very well for some countries. For others, the results seem to be very close to rejecting the hypothesis having low values for the estimated unit root parameter with large standard errors. Thus, the power of the test towards interesting alternatives is low. When we include the period of the Great Depression, we reject the unit root hypothesis firmly for both North American countries. All things considered, the data do not make a convincing case for unit roots in the aggregate consumption expenditure. The additional autocorrelation seems to fit better the predictions of the Life Cycle model.

2 THEORETICAL BASIS FOR A UNIT ROOT

2.1 The Hall model and extensions

In an innovating paper, Hall (1978) shows that if a household abides the Permanent Income Hypothesis (PIH) and forms rational expectations about its future income streams, this would imply that the household consumption path is close to a random walk. His model shows that all the information about future income and wealth available to the household at any time should be reflected in the consumption level.

The Hall framework applies the traditional assumptions of Permanent Income models of a single infinitely living household which represents the aggregate of consumers. Thus, the implications of the model apply to the macro economy so that testing of the model can be performed on aggregate data. The majority of the tests and discussions of the Hall model have been performed at the aggregate economy level. This paper fits into this tradition as we analyse data for aggregate consumption.

The model is based on the utility (expected) maximising household subjected to the usual budget constraint. The household is representative for the whole economy and no life cycle effects are involved. Consequently, the time horizon is set to infinity.

The preference ordering is expected to be additively separable in consumption over time. Preferences can then be represented by a utility function of the form

$$(2.1)\quad U(C)=\sum_{t=T_0}^{T=\infty}\left(\frac{1}{1+\delta}\right)^{t-T_0} u(c_t)$$

in which c_t is consumption at time t of the representative household, δ is the rate of time preferences and the remaining symbols have an obvious interpretation. C is a vector of all future consumption levels, c_t. Let $E(y_t \mid \Phi_{T_0})$ be the expected income at time t based on the information set Φ available to the household at time T_0. Further, let r be the economy-wide real rate of interest that the household expects in the future. It is assumed to be exogenous and we assume the financial markets to be perfect so that the interest rates on deposits and loans facing the household are equal. A_0 is the initial wealth at time T_0. We may write the budget constraint perceived by households as[2]

$$(2.2)\quad \sum_{t=T_0}^{T=\infty}\left(\frac{1}{1+r_t}\right)^{t-T_0} c_t^{(T_0)}=(1+r)\cdot A_{T_0}+\sum_{t=T_0}^{T=\infty}\left(\frac{1}{1+r_t}\right)^{t-T_0} E\left(y_t \middle| \Phi_{T_0}\right)$$

[2] Here, we adopt the timing convention that consumption and income take place at the end of the year while wealth is measured at the beginning. This is comparable to Galí.

We see that the present value of planned future consumption must equal the present value of expected future human and non-human wealth[3]. Superscript (T_0) indicates that this is an expected/planned value at time T_0. The Euler equations maximising the expected utility are

$$(2.3) \qquad E_{T_0}(mu_s) = E_{T_0}(mu_t) \cdot \left[\frac{(1+\delta)}{(1+r_t)}\right]^{(s-t)} \equiv \lambda^{(s-t)} \cdot E_{T_0}(mu_t)$$

where mu is the marginal utility of consumption for the household (s and t are time subscripts).

At time T_0 a plan for future consumption c_t is made so as to satisfy the Euler conditions based on the information about future income available to the household. The consumption c_{T_0} is realised along with the income y_{T_0}. In the next period, the initial wealth will change according to $A_{T_0+1} = A_{T_0} \cdot (1+r_t) + y_{T_0} - c_{T_0}$. This change in initial wealth is expected by the household and, thus, will not have any consequences for future consumption plans. If the household receives no new information in period $T_0 + 1$, consumption will go on as planned. However, if new information arrives, the household updates the information set and revises its future income prospects. The perceived change in the budget constraint as seen from point $T_0 + 1$ on will be

$$(2.4) \qquad \sum_{t=T_0+1}^{T=\infty}\left(\frac{1}{1+r}\right)^{t-T_0} E\left(y_t \middle| \Phi_{T_0+1}\right) - \sum_{t=T_0+1}^{T=\infty}\left(\frac{1}{1+r}\right)^{t-T_0} E\left(y_t \middle| \Phi_{T_0}\right) = \varepsilon_{T_0+1}$$

$$= \sum_{t=T_0+1}^{T=\infty}\left(\frac{1}{1+r}\right)^{t-T_0} E\left(c_t \middle| \Phi_{T_0+1}\right) - \sum_{t=T_0+1}^{T=\infty}\left(\frac{1}{1+r}\right)^{t-T_0} E\left(c_t \middle| \Phi_{T_0}\right)$$

Since households are assumed to form rational expectations, E(ε) must be 0. Households make a new plan contingent on the revision of future income possibilities. Furthermore, the

[3]In principle we could imagine an irrational household making impossible plans violating this constraint. The actual realization will have to satisfy this constraint and we shall disregard this possibility.

forming of rational expectations implies that the revision of consumption is based strictly on new information so that all information available in the previous period is unrelated to the present change. One immediate consequence is that changes in the planned consumption path should be orthogonal to all lagged variables. (*The orthogonality property*).

2.2 Quadratic utility

In his paper Hall discusses different types of utility functions. He concludes that the stochastic nature of consumption is not sensitive to the specification of the utility function, at least if the time span is not too long. We assume that utility is quadratic within periods. This is the standard assumption in much of this literature, and it is what Hall did in his empirical application. We have

(2.5) $$u(c_t) = -\frac{1}{2}\left(\bar{c} - c_t\right)^2$$

where $\bar{c}$ is a constant. This function has the convenient property that

(2.6) $$mu_t = \bar{c} - c_t$$

i.e. marginal utility is linear in consumption. Further, the function exhibits certainty equivalence. Uncertainty about future income should not affect the consumer's consumption/ saving decision today and only expected present value of future income is relevant for this decision. If we assume that the interest rate is perceived as constant and equal to households' rate of time preference $r = \delta$, households will plan for a constant level of consumption. Assuming normality of the error term, this implies that consumption should be a random walk[4].

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[4]If we allow that $\lambda \neq 1$, we have that consumption evolves according to

$$(2.7) \qquad \Delta c_t = e_t$$

e_t is the unplanned component due to new information about future income possibilities assumed to be white noise. Δ is the first difference operator. We refer to this implication of the model as *the unit root property,* since a random walk has a unit root and we shall abandon the strict random walk later.

In order to satisfy the budget constraint, a household that wishes a constant level of consumption, will consume the annuity of its total wealth (= the right hand side of the budget constraint given in 2.2). When the household has an infinite time horizon, this equals the real rate of interest. Consumption and income are linked by

$$(2.8) \quad c_t = r \cdot \left\{ A_{T_0} + \sum_{t=T_0}^{T=\infty} \left(\frac{1}{1+r} \right)^{t-T_0+1} E\left(y_t \middle| \Phi_{T_0} \right) \right\}$$

i.e. consumption is exactly proportional to the right hand side of the budged constraint (*the proportionality property*). Previously, the innovation in expression 2.8 was defined as ε_t. The implications are that

$$(2.9) \quad \Delta c_t = e_t = r \cdot \varepsilon_t$$

The change in consumption is proportional to the change in the expected value of the right hand side of the budged constraint.

$$\Delta c_t = (\lambda - 1) \cdot c_{t-1} + \bar{c} \frac{(r - \delta)}{(1+r)} + e_t \equiv (\lambda - 1) \cdot c_{t-1} + \beta_0 + e_t$$

The two additional terms are the planned changes in consumption.

2.3 Constant relative risk aversion utility

A possible alternative specification of the utility function is the constant relative risk aversion (CRRA) form used by Nelson (1987). He argues that this implies more reasonable behaviour.

$$(2.10) \qquad U(c_t) = \frac{c_t^{\gamma}}{\gamma}, \qquad \gamma < 1$$

and similar we have that

$$(2.11) \qquad mu_t = c_t^{\gamma - 1}$$

By using the Euler-condition expressed in marginal utility in equation 2.3 and continuing to assume that interest rates are constant, we see that for the CRRA - utility specification, we have that consumption will evolve according to

$$(2.12) \qquad \Delta \ln(c_t) = \frac{\ln(\lambda)}{(\gamma - 1)} + e_t$$

If we want the error term to be normally distributed, we have to assume that ε_t, the innovation in information about future income, has a multiplicative log-normal distribution. Now, the log of consumption evolves like a random walk with drift. If we assume $r = \delta$, the drift disappears and we will have that

$$(2.13) \qquad \Delta \ln(c_t) = e_t$$

Again we have a pure random walk. However in this case, it is the logarithm of consumption that has got this property.

If λ differs from 1 the deviation from a random walk could be significant in the long run in the Hall model. This could indicate that the assumption of a unit root in consumption expenditure on a large time span is not an exact representation. It is reasonable to believe that λ is not far from 1 so that the dominant (smallest in absolute value) root is very close to 1.

2.4 Liquidity constraints

An extension to the plain PIHRE model is the inclusion of liquidity constraints. Hayashi (1985) assumes that a portion of the population earning a share of ω of the aggregate income is liquidity constrained. The constrained part of the population consumes according to

$$(2.14) \quad c_t = y_t$$

Assuming that the remainder of the population consumes according to the PIHRE model, aggregate consumption behaviour could be more sensitive to income than in the pure Hall model. Now, consumption will be a weighted average of previous consumption behaviour and income according to

$$(2.15) \quad \Delta c_t = (1-\omega)\cdot r \cdot \varepsilon_t + \omega \cdot \Delta y_t$$

Whether consumption is more or less sensitive to changes in current income in the presence of liquidity constraints depends on the relative size of $r\varepsilon_t$ and Δy_t. If income is a pure random walk, they are equal and the two models cannot be distinguished. If an income change today implies expectations of further increases in the future, the consumption will actually be less sensitive to changes in income than is implied by the PIHRE. However, as long as some of the changes in income are regarded as temporary, we would expect an increased sensitivity of current consumption to current income. We can see that as long as we have a linear relationship between Δy_t and ε_t, the proportionality condition still holds true, but the factor of proportionality depends on the stochastic properties of income and the share of consumers being liquidity constrained. Furthermore, since changes in current income might be autocorrelated and correlated with other lagged variables, the orthogonality condition now fails. The portion of the population that is not liquidity constrained, will continue to consume according to a random walk. Since the sum of a unit root process and another non-explosive series contains a unit root, we have that the unit root property of the model survives. The correlation of the actual income path may add to the short run movements of aggregate consumption.

2.5 Durable goods

Hall disregards the existence of durable goods. Mankiw (1981) extended the work of Hall to show that the existence of durable goods forces the durable part of consumption expenditure to contain a negative moving average term in addition to the unit root. The reason is simple. If k is durable and the single consumption good in the economy and we use the quadratic utility function above, we have that the desired capital stock of durable goods in the household is

(2.16) $$k_t = \lambda \cdot k_{t-1} + \bar{k}\frac{(r-\delta)}{(1+r)} + e_t = \lambda \cdot k_{t-1} + \beta_0 + e_t$$

where e is the change in desired stock of durable goods due to news about future income possibilities. k has much of the same properties as we discussed for consumption expenditure earlier. The error term e will still be proportional to ε. If we let the durable good have a depreciation rate of d, we see that investment in durable goods must be

(2.17) $$c_t^i = k_t^i - (1-d)\cdot k_{t-1}^t$$

and we will have that the change in investments will follow

(2.18) $$\Delta c_t = e_t - (1-d)\cdot e_{t-1}$$

and if $\lambda \neq 1$

$$= (\lambda - 1)c_{t-1}^i + d\beta_0 + e_t - (1-d)\cdot e_{t-1}$$

Therefore, if the assumptions hold true, we would expect the expenditure series for individual households to be of the type ARMA(1,1). It still contains a unit root if λ is 1. We may transform this model to the same form as the augmented Dickey Fuller regression that we shall apply later.

$$(2.19)\quad c_t = \lambda \cdot c_{t-1} + d\beta_0 + e_t - \eta \cdot e_{t-1}$$

$$= d\beta_0 \cdot \left[1+\eta+\eta^2+...+\eta^i+...\right] + (\lambda-\eta)\cdot c_{t-1} + \eta\cdot(\lambda-\eta)\cdot c_{t-2}+...+\eta^i\cdot(\lambda-\eta)\cdot c_{t-i}+...+e_t$$

This can be transformed into

$$= \frac{d\beta_0}{(1-\eta)} + \frac{(\lambda-\eta)}{(1-\eta)}\cdot c_{t-1} - \sum_{j=1}^{\infty}\eta^j \cdot \frac{(\lambda-\eta)}{(1-\eta)}\cdot \Delta c_{t-j} + e_t$$

and if $\lambda = 1$

$$= c_{t-1} - \sum_{j=1}^{\infty}\eta^j \cdot \Delta c_{t-j} + e_t$$

This is the same form as we use in our estimations. We should expect the negative moving average term to appear as negative estimated parameters for the lags of first differences added to the equation. The more durable the goods, the higher the value be will for the estimated lag coefficients.

2.6 Other extensions

Bernanke (1985) develops a model including both a durable and a non-durable good. He considers a non-separable utility function and adjustment costs[5]. He shows that this is a much more complicated structure and that the results from the Hall model do not have any straight forward extension in this case. In order to be able to implement the model, Bernanke abstracts from changes in relative prices between the two goods. He demonstrates that with this simplification his consumption expenditure functions for the two goods have the same unit root property. Startz (1989) explores a similar model where he considers adjustment costs in durables purchases and non-separability between durables and non-durables in consumption. As long as he does not consider changes in relative prices between the two goods, $\lambda = 1$ keeps the unit root property in this case. If he allows changes in relative prices, the real rate of interest can be measured in both prices and it is not clear what is meant by $r = \delta$.

Orzag and Staroselsky (1993) look into the assumption of a representative agent. Their point is that even if we aggregate over a set of different consumers, we may get a unit root model even though individual behaviour does not follow the Hall model. A unit root in aggregate consumption should therefore be considered a more general property than just being a result of this model. As a consequence, the finding of a unit root is not a proof of Hall's model.

In the empirical section later on, we work with consumption in logarithmic form. The main reason for this is not that we have a strong preference for the CRRA utility function over the quadratic one. However, the data show a steady growth over time. We believe that this trend is best approximated by an exponential growth rate. Further, it seems more reasonable to assume that the variability in the data is constant in percentage terms rather than in absolute terms. When data are transformed by logarithms we probably avoid a possible problem of heteroscedasticity.

Besides the fact that λ might not be equal to 1, for several reasons we do not expect the IMA(1,1) representation to be an exact description of consumption behaviour in the aggregate. First, we are looking at the sum of several households and persons. The population growth and family structure change over time and may cause additional serial correlation in the data. We try to some extent to correct for this by working with consumption expenditure

[5]Adjustment costs are needed in the formulation to smooth out investments in durables. Mankiw (1981) among others shows that durables are too smooth according to the theoretical predictions.

per capita. Second, there may be adjustment costs contributing to the autocorrelation structure. Third, the replacement of old generations by new ones having a higher lifetime expected wealth, gives a steady rise in per capita consumption over time. This effect may be the main driving force of the movement of aggregate consumption over time.

Fourth, the IMA(1,1) model is derived for a single household. When we apply this on aggregate data, we make strong assumptions about co-ordination of news and timing among different households. This is not quite realistic. Fifth, the assumption about additive separability in preferences over time may be too restrictive. As a consequence, the consumption path could be close to, but not actually contain, a unit root. Finally, time aggregation may induce a moving average term to the series, and Working (1960) showed that, in the limit, when households reoptimise their consumption level continuously, this parameter should equal 0.25[6]. All this said, the unit root should probably still remain in the aggregate series. However, we do approach the unit root and correlation testing with an open mind uncertain of what to expect.

2.7 The Galí model

Galí (1990) uses a life cycle formulation of the household optimisation problem. In this setting, he demonstrates that even though individual behaviour appears to be the same as in the Hall model, the behaviour of aggregate consumption will track the trend of aggregate income more closely. If income is stationary, we should expect that consumption is stationary as well[7]. We show that, strictly speaking, this assumption is correct, but the aggregate consumption path is still very close to a unit root process. For practical purposes, we would expect the two models to be indistinguishable in this regard. We show, however, that the life cycle model will have additional positive autocorrelation in the augmented Dickey-Fuller regression in contrast to the model of Hall.

[6]Christiano et al.(1991) investigate this assumption further and find clear evidence in support of this. In several of our models, we find similar results of a second order autoregressive parameter close to .25, but some countries like Sweden and Japan show a much lower value and UK a much higher value for this parameter.

[7]In the discussion Galí limits himself mostly to the case where income is non-stationary. There is fairly strong evidence indicating that income is a stationary process. In de Haan and Zelhorst (1993), they show that on a set of 12 OECD countries (which comprised all the countries we consider here except Japan), using yearly data for a long time period, 1870-1989, the unit root was rejected at a high level in all cases.

The model of Galí can be regarded as a generalisation of the Hall model since it contains this as a limiting case. Basically, the main difference between the models is the way they treat the bequest motive. In the Hall model, households are assumed to be perfectly altruistic in the sense that any shocks and innovations in aggregate human and non-human wealth are handed over to future generations. Only the real rate of return is consumed today. In the Galí model, the bequest motive is left out by assumption. Although a change in labour income might indicate a change in human wealth for future generations, the present households care only about themselves. They consume all of their change in wealth within their lifetime. Every new generation starts out with no non-human wealth. The possibility of postponing consumption to future generations in the Hall model permits a separation of the development of aggregate consumption from that of aggregate income. In this model the two must be linked in the long run.

2.7.1 The model and household behaviour

Galí assumes that the household has a probability p of dying in every period independent of age[8]. Let p be the number of members in the new-born generation as well, so that the population sum to 1. In order to remove uncertainty about future time of death and to leave no bequest to future generations, he assumes that there exist annuity markets. A firm inherits the financial wealth or debt of a consumer when he dies. The remaining members of this generation will share this wealth so that the firm has no costs and makes no profit. Since all households within a generation are equal and have the same amount of non-human wealth, the effective interest rate facing a household will be $z = (1+r)/(1-p) - 1$[9].

The household maximises expected utility as in the Hall model but in this model future utility must be weighted by the probability of staying alive since people receive no utility when they are dead. We apply a similar quadratic utility function as before,

$$(2.20) \qquad U(C) = -\frac{1}{2}\sum_{t=T_0}^{\infty}\left(\frac{1-p}{1+\delta}\right)^{t-T_0}\left(\bar{c} - c_t\right)^2,$$

[8]This assumption do not seem critical to the results and could have been relaxed to make p dependant on age. This would make the model more complicated and less tractable and would probably not add much insight.

[9]The surviving (1-p) households receive the return on capital r from the disappering (p) housholds so that $(1-p)(1+z) = 1+r$. Thus, we have the effective rate of interest above.

and the budget constraint can be expressed as

$$(2.21) \qquad A_{s,t+1+j} = (1+z)\cdot A_{s,t+j} + y_{s,t+j} - c_{s,t+j}$$

and

$$(2.22) \qquad \lim_{j\to\infty}(1+z)^{-j} A_{s,t+j} = 0$$

Thus, $A_{s,t}$ is the non-human wealth in period t of the household born in period s.

Optimising behaviour gives us the same Euler condition as before since the 1-p in the utility function and the same 1 - p as a part of z in the budged constraint cancel out. This seems intuitively reasonable since the increased return on savings is the same as the increased possibility of not benefiting from it. If we assume that the real interest rate r equals the rate of time preferences δ, we have that households plan for a constant level of consumption. However, the household will consume a fraction z of its total wealth instead of r. This is sustainable due to transfers in the annuity market. The consumption level of a household born in period s will be

$$(2.23) \qquad c_{s,t} = z\cdot\left(A_{s,t} + H_{s,t}\right)$$

where $H_{s,t} \equiv \sum_{j=1}^{\infty}(1+z)^{-j} E\left(y_{s,t+j-1}\middle|\Phi_t\right)$ is the human wealth. Household income is governed by two factors. First, we have a stochastic process for wages to which we shall return later. Labour is, of course, assumed to be homogeneous so that the same wage rate applies for all labour. Second, household labour supply is assumed to be exogenous and a geometrically decaying function of time with α as decaying parameter

$$(2.24) \quad L_{s,t} = \frac{\Gamma}{p} \cdot (1-\alpha)^{t-s}$$

Γ is a normalising factor equal to (1-(1-α)(1-p)) to ensure that total labour supply equals 1. Γ is between p and 1 depending on the value of α. α gives rise to life cycle saving. We see that if α and p are set to zero, we return to the model proposed by Hall.

2.7.2 Aggregate behaviour of the model

Since individual households consume in proportion to their aggregate wealth, this will apply to the aggregate as well

$$(2.25) \quad c_t = z \cdot (A_t + H_t),$$

Although individual consumption behaves like a random walk, aggregate consumption will not inherit this property. The random walk governs the evolution of individual consumption over its life cycle. New generations will start with no financial wealth received from their predecessors and their consumption will be scaled according to their expected future income possibilities in the economy.

$$(2.26) \quad c_{t,t} = z \cdot [H_{t,t}]$$

They will not continue the consumption path of the last period's deceased. In order to trace the effect that an innovation in the income process has on aggregate consumption, we need to describe the total human wealth of the present generations. Household human wealth of different generations (s and s') will differ with a proportionality factor of $(1-\alpha)^{s-s'}$. The only difference in their income is the level of labour supply. We can add the human wealth of all presently living households to get

$$(2.27)\quad H_t \equiv \sum_{i=t}^{-\infty} H_{i,t} \cdot N_{i,t} = H_{t,t} \sum_{i=0}^{-\infty} (1-\alpha)^i \cdot p(1-p)^i = H_{t,t} \cdot p/\Gamma$$

$N_{i,t} = p(1-p)^{t-j}$ is the population of generation i that is alive at time t. Aggregate human wealth is proportional to the wealth of members of the new-born generation. If we replace the expression for $H_{t,t}$, we have that total human wealth as a function of future expected income will be

$$(2.28)\quad H_t = \sum_{j=1}^{\infty} (1+z)^{-j} (1-\alpha)^{j-1} E\left(y_{t+j-1} \middle| \Phi_t\right)$$

The households that live today and survive to the next period are expected to keep the same level of consumption that they have in the present period. Since the probability of dying in each period (p) equals the number of new consumers, we have that the expected consumption level in the next period will be

$$(2.29)\quad E_t c_{t+1} = (1-p)c_t + pE_t c_{t+1,t+1}$$

Although individual households do not plan a growth in their consumption path, this does not imply that aggregate consumption is stationary. A general rise in the income level over time, will mean that new generations start at a higher level of consumption, which causes aggregate consumption to grow.

2.7.3 The income process and its effects on aggregate consumption

Galí restricts his attention to the case where the income innovation process has a unit root. He applies a general non-specific income process with a trend. We intend to deviate from this and shall consider the relation between the income process and the evolution of aggregate consumption. Any trend in income is disregarded for simplicity. Further, there seems to be no need to restrict the attention to unit roots in income and we wish to be more general on this issue. We believe that there is a considerable evidence showing that income might be stationary (see de Haan and Zelhorst (1993)). Let us consider the simple AR(1)-process in income[10].

$$(2.30) \quad y_t = \mu + \rho y_{t-1} + e_t$$

We intend to trace out the isolated effect that the innovation e_t has on all future consumption. The innovation in income will not make changes to current non-human wealth and will only affect wealth in future periods. The relationship between the stochastic innovation in current aggregate income and that of current aggregate consumption can be expressed by looking at the impact of the current income change on current human wealth. That is, the wealth of currently living generations. The innovation will change future income level according to

$$(2.31) \quad \frac{\partial y_{t+i}}{\partial e_t} = \rho^i$$

The aggregate consumption and human wealth for current generations will consequently change.

$$(2.32) \quad \frac{\partial c_t}{\partial e_t} = z \cdot \frac{\partial H_t}{\partial e_t} = \frac{z}{(1+z)} \sum_{i=0}^{\infty} \left(\frac{(1-\alpha)\rho}{1+z} \right)^i = \frac{z}{1+z-\rho+p\cdot\alpha} \equiv \Psi_0$$

The currently living generations plan to maintain a constant level of consumption at the household level and will not change their future consumption unless they are exposed to a

[10] We tried to generalize this to an AR(2)-model, which is commonly used when income is modelled. However, it seems difficult to solve explicitly for the implications for aggregate income in any convinient form. This problem could be solved by simulation. It is my guess that this would not change the fundamental conclusions in any important manner.

new shock. The human capital of the new generation born in the next period will be effected, however. This can be realised if we consider

$$(2.33)\qquad \frac{\partial H_{t+1,t+1}}{\partial e_t}=\frac{\Gamma}{p}\cdot\frac{1}{(1+z)}\cdot\frac{\partial\left\{\sum_{i=0}^{\infty}\left(\frac{(1-\alpha)}{(1+z)}\right)^{i} y_{t+1}\right\}}{\partial e_t}$$
$$=\frac{\Gamma\rho}{p}\frac{\partial H_t}{\partial e_t}=\frac{\Gamma\rho}{p}\cdot\frac{\Psi_0}{z}$$

An innovation in the income process will change the human wealth of the next periods' new-born generation in proportion to the change it causes to the current consumers. The proportionality factor consists of ρ since the effect is reduced on future income if income is stationary and Γ/p which relates human capital of one new-born household to the total human capital of all generations. The income innovation shock will consequently have the following impact on aggregate consumption in future periods

$$(2.34)\qquad \frac{\partial c_{t+1}}{\partial e_t}=(1-p)\cdot\frac{\partial c_t}{\partial e_t}+p\cdot z\cdot\frac{\partial H_{t+1,t+1}}{\partial e_t}$$

$$=(1-p)\cdot\Psi_0+\Gamma\cdot\rho\cdot\Psi_0\equiv\Psi_1$$

$$\frac{\partial c_{t+2}}{\partial e_t}=(1-p)\cdot\frac{\partial c_{t+1}}{\partial e_t}+p\cdot z\cdot\frac{\partial H_{t+2,t+2}}{\partial e_t}$$

$$=(1-p)\cdot\Psi_1+\Gamma\cdot\rho^2\cdot\Psi_0\equiv\Psi_2$$

$$\frac{\partial c_{t+3}}{\partial e_t}=(1-p)\cdot\Psi_2+\Gamma\cdot\rho^3\cdot\Psi_0\equiv\Psi_{3l}$$

$.......etc.$

Since Γ is less than 1, we have that

$$\text{(2.35)} \quad \Psi_1 \leq (2-p)\cdot\Psi_0, \Psi_2 \leq (1-p)(2-p)\cdot\Psi_0 + \rho^2\cdot\Psi_0, \\ \Psi_3 \leq (1-p)^2(2-p)\cdot\Psi_0 + \rho^3\cdot\Psi_0, etc.$$

As long as the income process is stationary so that ρ < 1, the strict inequality applies. The effects of a shock will eventually die out and the maximum limit will start decaying from period t+2. Old generations will die out and so will the unit root in their consumption path. New generations will start out in direct proportion to the income stream. But as long as the income innovation process is stationary, a shock will eventually die out in the consumption of new generations as well. As old generations are replaced by new, there will be a tendency for the aggregate consumption to track the income evolution and if income is stationary then consumption will be stationary as well.

Furthermore, if we continue to leave out any constant and trend terms, the moving average process of consumption can be expressed as

$$\text{(2.36)} \quad c_t = \Psi_0\cdot e_t + \Psi_1\cdot e_{t-1} + \Psi_2\cdot e_{t-2} + \Psi_3\cdot e_{t-3} + etc.$$

$$= \Psi_0\cdot e_t + (1-p)\cdot\Psi_0\cdot e_{t-1} + \Gamma\cdot\rho\cdot\Psi_{0,}\cdot e_{t-1} + (1-p)^2\cdot\Psi_0\cdot e_{t-2} + (1-p)\cdot\Gamma\cdot\rho\cdot\Psi_0\cdot e_{t-2} + \Gamma\cdot\rho^2\cdot\Psi_0\cdot e_{t-2}$$

$$+ (1-p)^3\cdot\Psi_0\cdot e_{t-3} + (1-p)^2\cdot\Gamma\cdot\rho\cdot\Psi_0\cdot e_{t-3} + (1-p)\cdot\Gamma\cdot\rho^2\cdot\Psi_0\cdot e_{t-3} + \Gamma\cdot\rho^3\cdot\Psi_0\cdot e_{t-3} + etc.$$

If we subtract $(1-p)\cdot c_{t-1}$ from the equation above we have that

$$\text{(2.37)} \quad c_t - (1-p)\cdot c_{t-1} = \Psi_0\cdot e_t + \Gamma\cdot\rho\cdot\Psi_0\cdot e_{t-1} + \Gamma\cdot\rho^2\cdot\Psi_0\cdot e_{t-2} + \Gamma\cdot\rho^3\cdot\Psi_0\cdot e_{t-3} + etc.$$

Further manipulation gives us

$$(2.38)\quad c_t-(1-p)\cdot c_{t-1}=\Psi_0\cdot\left[e_t+\Gamma\sum_{j=1}^{\infty}\rho^j\cdot e_{t-j}\right]$$

$$\frac{c_t-(1-p)\cdot c_{t-1}-(1-\Gamma)\cdot\Psi_0\cdot e_t}{\Gamma}=\Psi_0\cdot\sum_{j=0}^{\infty}\rho^j\cdot e_{t-j}$$

The right-hand side of this equation can be recognised as the moving average representation of a first order autoregressive process. We may then transform the above equation to

$$(2.39)\qquad \frac{c_t-(1-p)\cdot c_{t-1}-(1-\Gamma)\Psi_0\cdot e_t}{\Gamma}=\rho\cdot\frac{c_{t-1}-(1-p)\cdot c_{t-2}-(1-\Gamma)\Psi_0\cdot e_{t-1}}{\Gamma}+\Psi_0\cdot e_t$$

When we solve this, we end up with an ARMA(2,1) representation of aggregate consumption

$$(2.40)\qquad c_t=(1-p+\rho)\cdot c_{t-1}-\rho\cdot(1-p)\cdot c_{t-2}+\Psi_0\cdot e_t-\rho\cdot(1-\Gamma)\cdot\Psi_0\cdot e_{t-1}$$

If we define $\theta_1\equiv(1-p+\rho)$, $\theta_2\equiv-\rho(1-p)$, $\varphi_1\equiv-\rho(1-\Gamma)$ and let $\omega\equiv\theta_2-\varphi_1\theta_1-\varphi_1^2$, we can put the consumption equation on the same form as the augmented Dickey-Fuller regressions

$$(2.41)\quad c_t=(\theta_1+\varphi_1)\cdot c_{t-1}+\omega\cdot c_{t-2}-\varphi_1\cdot\omega\cdot c_{t-3}+\ldots+(-\varphi_1)^j\cdot\omega\cdot c_{t-j}+\ldots+\Psi_0e_t$$

$$=\left[(\theta_1+\varphi_1)+\frac{\omega}{(1+\varphi_1)}\right]\cdot c_{t-1}-\frac{\omega}{(1+\varphi_1)}\cdot\Delta c_{t-1}+\varphi_1\cdot\frac{\omega}{(1+\varphi_1)}\cdot\Delta c_{t-2}$$

$$-\varphi_1{}^2\cdot\frac{\omega}{(1+\varphi_1)}\cdot\Delta c_{t-3}-\ldots-(-\varphi_1)^j\cdot\frac{\omega}{(1+\varphi_1)}\cdot\Delta c_{t-j-1}-\ldots+\Psi_0e_t$$

If we take a closer look at the autoregressive parameter at lag 1, we have that $1>(\theta_1+\varphi_1)+\omega/(1+\varphi_1)=1-p(1-\rho)/(1-\rho+\rho\Gamma)>1-p$. This process is a non-explosive process since the unit root test observer is less than or equal to one. However, it is larger than 1-p. If we imagine that an average life expectancy is 50 years for a household, this gives us a value of p of .02. If income is a unit root process, ρ = 1 and aggregate consumption will contain a unit root. It will not be a random walk, however. At the other extreme, we can imagine that income innovations today are purely transitory and have no implications for future income. Then the parameter will be 1-p. The households that survive will smooth their consumption of the transitory income. It will have no effect on future generations. Although the process appears to be stationary, we see that it is very close to a unit root independent of the income process and empirically it will be indistinguishable from such a unit root alternative.

The additional autocorrelation distinguishes the Life-Cycle model from the model of Hall. In the pure Hall model which is comparable to this one, there should be no additional autocorrelation. Here, the parameters are clearly positive. We have that ω is described by $-\Gamma\rho<\omega=-\Gamma\rho((1-p)-\rho(1-\Gamma))<-\Gamma\rho(1-p-p\rho)<0$. For the whole parameter it must be that $-\Gamma\rho<\omega/(1+\varphi_1)=-\Gamma\rho+\Gamma\rho p/(1-\rho+\rho\Gamma)<0$.(The last inequality is strictly speaking only correct as long as p is less than .5.) Consequently, the parameters for the lagged terms in first differences are positive and less than 1. Reasonable values of p indicate that ω is closer to $-\Gamma\rho$ than 0. The actual value of this parameter depends crucially on the importance of the size of the life cycle saving motive. If α is high these parameters will be high. A value of 0.1 implies that $-\omega/(1+\varphi_1)$ is about 0.12 (p=0.02, ρ=1). As we shall see from the empirical results later on, these predictions of positive values of the lagged first differences, are quite in agreement with the actual results obtained from data, and differ from what we would expect from the Hall model with durable goods. However, left alone these are not sufficient to explain all of the observed behaviour of the data.

The model above could be adopted to accommodate durable goods like in the Hall model. We would have to interpret c as the desired capital stock and substitute by using the investment function as we did in the Hall model. However, the expressions became very complex and we were not able to reduce it to any transparent formulation.

The main source behind the differences between the two models is not that the bequest is set to zero in the Galí model, but that the bequest is unrelated to the income innovation process. To be able to disconnect the stochastic process of aggregate consumption and aggregate income in the long run, the expected bequest must contains a unit root if income is stationary. This is the only way that saving can become non-stationary (except for a possible trend).

3 STATISTICAL CONSEQUENCES OF UNIT ROOTS

As we saw above, the existence or non-existence of unit roots in aggregate consumption may have important theoretical implications. Apart from that, there are well-founded empirical reasons for the interest in unit roots. If a stochastic process contains a unit root, a quite different asymptotic distribution theory applies than in the stationary processes normally considered. If the error term contains a unit root, ordinary least squares estimation is biased and the distribution of the estimates may be far from normal. Hence, the t-ratios can be highly misleading. Furthermore, problems may arise with spurious regression effects. Possibly, some of the regression variables may be co-integrated so that the error term may be stationary even if some of the included variables are not. These are issues necessary to consider before modelling consumption.

Since a unit root in a time series means that a shock has permanent effects, unit roots are important if we want to make forecasts. If a unit root is present, the forecast error in some distant future will have a much larger variance, and the possibility of forecasting will be much more limited than if the series is stationary.

All this makes it important to try to decide whether private consumption actually contains a unit root or not. It is also interesting to consider if the same result applies when we look at different countries.

4 SOME EARLIER WORKS ON UNIT ROOTS IN PRIVATE CONSUMPTION

A considerable amount of work has been done to test the implications of the combined hypothesis of Permanent Income and Rational Expectations using aggregate data. Hall looked at the orthogonality implication of the PIHRE model and found that most other series failed in predicting future consumption beyond the predictions from lagged consumption itself. His tests were not able to reject the hypothesis using lagged income data, but did find a significant, although small, impact from the stock market. Later, several authors have claimed that his tests probably lack the necessary power and have looked at models containing more structure between consumption and income (like Flavin (1981), Campbell (1987), Deaton (1987) and Campbell and Deaton (1989) among others). In general, these studies have shed some doubt on the PIHRE model.

Common to the tests of the PIHRE model is that they make use of the orthogonality or the proportionality condition between income and consumption. Even though unit roots in time series have been very much on the agenda during the last decade, the PIHRE hypothesis has to a less extent been evaluated on the basis of its unit root implications. Most of the work on unit roots has concerned GDP, exchange rates, interest rates and monetary data. The famous study of Nelson and Plosser (1982) which led to the debate of unit roots in macroeconomic time series, did not include any measures of private consumption among their 14 different series!

Still, some work on unit roots has been done on data for private consumption expenditure as well, mostly as part of other work like testing for co-integration of consumption and income. As a result of that, it is difficult to find the estimates done for different countries and my overview is probably not complete.

Common to tests on private consumption is the use of quarterly data after World War II. One exception is Peel (1992). He tests nominal data for UK from 1830 to 1990 on a yearly basis. He does not offer the tests much comment and does not show the results but it seems as if he has accepted the hypothesis for all different segments of the data[11]. In his testing procedure it might very well be the price level that has a unit root.

Several of the quarterly data series have been seasonally adjusted (SA) in advance which induces a bias towards non-rejection of a unit root (See Ghysels and Perron (1993) for a discussion of this problem).

Campbell (1987) tests the unit root hypothesis for consumption as a part of a co-integration test of consumption and income. He uses data from the USA (Q(SA)[12] 1953:2 to 1984:4). He cannot reject the hypothesis on these data at a 90%-level using Phillips-Perron tests, although the test (τ) values are as high as 2.51 for total and 2.78 for non-durables.

[11]He uses a Phillips-Perron testing framework and states that in no case is the null hypothesis not rejected. However, as he continues with a cointegration analysis he cannot have rejected the unit root.
[12]Q = quarterly data, (SA) = seasonally adjusted.

MacDonald & Speight (1989) treat the co-integration problem in UK data (Q(SA) 1966:1-1986:3). Their test values are low and the unit root hypothesis is not rejected by their augmented Dickey-Fuller test neither for non-durables nor for total consumption. MacDonald and Kearney (1990) take a closer look at Australia (Q 1959:3 -1989:2). There is no sign of rejection of the unit root hypothesis in these data.

Otto and Wirjanto (1990) regard the seasonal unit roots using the framework of Hylleberg et al.(1988) in different time series for Canada (Q 1955:1-1988:4). At the zero frequency, they must reject the hypothesis of a unit root for durables and services. Total private consumption expenditure is barely accepted at a 95%-level having a test value of 3.42 (critical value = 3.50) casting doubt about the unit root in these data.

5 THE DATA

We wanted to do a cross-country study comparing the behaviour of aggregate consumption expenditure across different economies. After all, the theory is quite general and should apply to different subsets of consumers like separate countries. In doing so, it was important to get comparable data for the countries included so as to simplify the interpretation of the results. We decided to make use of the aggregate private consumption expenditure in real[13] terms for each country, and we collected the official data that could be obtained from the national statistical bureaux.

We wanted to use the G7-countries since these economies cover a considerable part of the world economy. In addition, we include Norway and Sweden since small countries could show a different pattern and because rather good statistical data are available. The Scandinavian countries have had some periods with tight credit rationing, but most of these countries have had a well developed financial sector in the period we are considering, reducing the risk of any distortions in the results from credit rationing[14]. In the end, we obtained data for total real private consumption expenditure for seven OECD-countries;

[13]The use of nominal data may bias the results towards non-rejection of the unit root hypothesis, since the price level is likely to contain a unit root.

[14]Even though some households are liquidity constrained, this should not invalidate the unit root implication of the PIHRE- model.

Canada (1926-1990), Germany (West)(1950-1990), Japan (1955-1990), UK (1948-1990), US (1929-1990), Sweden (1950-1991) and Norway (1865-1939 and 1946-1991)[15].

In the German data, there is a break in 1960 when West-Berlin and Saarland are included in the data. This raises the level of the series by about 6.8%. Since we have received the observations for 1960 with and without these regions, we have simply rescaled the observations for the period before 1960 to be comparable to the rest[16].

Jaeger (1990) concludes that in pre-war GNP data (1869-1929) for USA the use of interpolating techniques in construction of the data has biased the results towards rejecting the unit root hypothesis. In Swedish and UK data, however, he finds that interpolation is used to a much smaller extent and he cannot reject the hypothesis for these data. In the construction of the Norwegian consumption data, some interpolation has been used in the early periods.

In addition to the consumption series, we have collected population statistics for the countries in the relevant years. We analyse all our series in per capita terms.

As advocated earlier, we analyse the logarithm of the consumption data. In the figures at the end, we have drawn the levels of the series and the first differences of the logarithms ΔC_t (= $\log(C_t/C_{t-1})$. USA and Canada show a lot more fluctuations in the data before 1946. This may indicate that a structural change has taken place after the war. Also, we see the great depression in the thirties showing up as large negative growth values.

In the Norwegian data, we apparently have the same structure during the whole series except during the First World War and in the following years, when the aggregate consumption showed large fluctuations. Even though Norway was neutral during World War I, there was a strong rationing of goods in 1917 and 1918. Combined with a large increase in the money stock during the war, this created a consumption boom in 1919 and 1920 with high inflation. It was followed by some unstable years. When we include this period in the sample, it is likely to have great influence on the estimation results. We shall consider different sample selections. The other series are somewhat more difficult to interpret. Germany and Japan show a tendency of a falling growth rate during the sample period. As we shall see later when

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[15]We also tried to obtain data from Italy and France. It seemed difficult to obtain long enough data series for France while the Italian bureau was like "The Castle" of Kafka.

[16]We tried to use a dummy variable for this year as well. The consequence was marginal and did not have any effect on the unit root testing.

we use the Bayesian estimation procedure, for Germany the unit root hypothesis can only be supported when we allow a negative trend.

As is obvious, the real aggregate consumption expenditure shows steady growth during the period we are looking at. The main alternative to a unit root will be a stationary series with a linear trend.

6 TESTING PROCEDURE

According to Campbell and Perron (citing earlier works of Perron), the time span rather than the number of observations per se is what decides the power of the unit root tests. Since the power issue is important in these tests, this has motivated our use of annual data. Annual data are available for a much longer period than quarterly data. We end up using only 36 - 65 observations (more for Norway). Although in quarterly data we often have more than 100 observations, there is reason to believe tests on annual data to be at least as powerful as the tests based on more frequent data. In addition, we avoid the problems connected to seasonal patterns in consumption and adjustments for these. Of course, analysing a longer time period increases the possibility of structural changes possibly biasing the test results. Such changes not comprised in the alternative hypothesis, will often tend to bias the results towards non-rejection as was the case in the paper of Perron (1989).

When we test for a unit root, we cannot apply the usual standard tests from traditional econometrics. If we have an autoregressive model in a variable X like

$$X_t = \mu + \rho \cdot X_{t-1} + e_t \tag{6.1}$$

the estimated parameters will be biased if we apply OLS and the usual t-statistics will not be correct. If $\rho < 1$ in absolute value, we have a stationary series and X will tend towards μ in the long run. If $\rho = 1$, however, this is no longer true. There is no mean reverting tendency for this process so that in the long run the process may very well drift away from zero.

Asymptotically, the two different models, $\rho < 1$ and $\rho = 1$, are of a quite different nature[17]. OLS-estimation will bias the estimated parameter ρ^* towards zero if the true parameter is 1.

In testing for a unit root, we make use of the standard unit root tests of Dickey and Fuller (DF-(τ,ρ)) and the comparable tests of Phillips and Perron (PP-(τ,ρ)). These tests are based on the standard ordinary least squares regression, but the confidence limits are adopted to accommodate the bias in the estimated values if the unit root is true.

In addition to these classical tests, we consider the Bayesian test proposed by DeJong and Whiteman (1991a). While the classical tests treat the unit root as the null hypothesis and the tests are based on the assumption that the unit root is actually true, the Bayesian test is closely related to the likelihood function. The likelihood function is based on normal distribution theory and is strictly correct only if the model is stationary.

Our basic assumption in the analysis is that consumption expenditure can be expressed within the equation

$$(6.2) \qquad c_t = \alpha + \beta_t t + \sum_{j=1}^{k} \varphi_j c_{t-j} + e_t$$

where c_t is our measure of aggregate consumption expenditure and e_t is white noise. This model contains the possibility of a unit root process with $\beta_t = 0$ and $\sum_j \varphi_j = 1$ (smallest root = 1). α is then a drift parameter making the series c grow steadily[18]. Furthermore, the model includes the possibility of a trend stationary process where $\beta_t > 0$ and the smallest root is above 1. The roots are the roots of the polynomial p(z) that we define below.

To evaluate the persistence in the error structure, we are seeking the roots of the polynomial

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[17]As Phillips (1991) among others points out, this may lead to split confidence intervals concerning the parameter ρ. This should not be considered as a weakness of the testing procedure since the model behaves so differently in the two cases.

[18]In case of a unit root we do not restrict β to be zero in the estimation procedure. However, the combination of a unit root and a time trend in the long run does not seem plausible, nor is it implied by our theoretical models. A quadratic trend in the stationary case is not convincing for the same reason. β is a necessary parameter in the estimation since else the plausible alternative hypothesis would be suppressed.

$$(6.3) \qquad p(z) = 1 - \sum_{j=1}^{k} \varphi_j z^j = \prod_{j=1}^{k} (1 - \lambda_j z)$$

Here, $1/\lambda_j$ are the roots. If all the roots exceed 1 in absolute value, c is a trend stationary series. If one of the roots is less than 1 in absolute value the series is explosive, while if the smallest root is exactly on the unit circle, consumption expenditure has a unit root. We denote the inverse of the smallest root by Λ (the highest λ_i). If we insert z = 1 in p, we see that if c has a unit root, then $\Phi \equiv \sum_{j=1}^{k} \varphi_j = 1$. In the classical tests, we exploit this directly by testing whether this sum equals 1, while in the Bayesian test we look at $\Lambda = 1$.

In the theory section, we saw that the Galí model and the Hall model with durable goods both contained a MA(1)-term. We showed in these sections how these models can be written as an infinite autoregressive model. However, the basic model for estimation is truncated at k. In Said and Dickey (1984), they show that mixed ARMA-models can be consistently estimated by using the augmented Dickey-Fuller estimation procedure as long as k, the number of included lags, increases sufficiently slowly compared to the number of observations[19].

6.1 Set-up for the classical tests

These tests should be rather familiar to most readers. We apply the augmented Dickey-Fuller (DF) tests and the tests developed by Phillips and Perron (PP). When we include both post- and pre-war data, we consider the possibility of a break in trend and in the constant term a model which Perron (1989) developed in his paper on the great crash in 1929 and on the oil price shock in 1973.

When we look at the post-World-War II data, we assume that the null hypothesis is a unit root process with a drift. The alternative hypothesis is a stationary process around a linear trend. We make use of a one-step estimation procedure. It means that for all countries we estimate Dickey-Fuller regressions of the type:

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[19]They assume that $n^{-1/3}k \to 0$ and that there exist c > 0, r > 0 such that $ck > n^{1/r}$. In this case the Dickey-Fuller τ-test is valid but the confidence intervals for the ρ-test depends on some of the unknown parameters.

$$(M1) \qquad c_t = \alpha + \beta_1 \cdot t + \pi \cdot c_{t-1} + \sum_{j=1}^{k} \gamma_j \cdot \Delta c_{t-j} + e_t$$

c_t now represents the natural logarithm of total aggregate consumption expenditure per capita at time t and Δ gives the first difference of the series. The first difference terms compensate for additional serial correlation in the error structure. This formula is equivalent to the AR(p)-model (6.2) used above, since $\pi=\Phi$, j=p-1 and $\gamma_j = -\Phi + \sum_{i=1}^{j} \varphi_i$. Our null hypothesis is that the series contain a unit root meaning $\pi = 1$. The alternative is that there is no unit root present in the series, which is equivalent to saying that $\pi < 1$. As long as we cannot reject the possibility of $\pi = 1$, we shall not reject the hypothesis of a unit root.

We use the τ-test and the normalised bias (ρ-) test developed by Dickey and Fuller in several papers and extended by Said and Dickey (1985), to include additional autocorrelation. The τ-test is simply based on the same test statistic as the traditional t-test, but because of the different asymptotic nature of the unit root models and traditional models different distribution must be applied in order to evaluate the test. The ρ-test is calculated using

$$(6.4) \qquad test\ stat. = \frac{n(\pi - 1)}{\left(1 - \sum_j \gamma_j\right)}$$

In addition, we apply the Phillips-Perron τ- and ρ-tests. These are based on the same principle as the Dickey-Fuller tests, but use a different method to correct for the additional autocorrelation above. Instead of including the lagged first differences of the series in the estimated regression, they use a Newey-West transformation of the estimated parameters using the estimated error terms. The tests are developed in Phillips (1987) and Phillips and Perron (1988). The same asymptotic distribution theory applies to both the DF- and the PP-tests and the relevant probability tables are in Fuller (1976).

For Norway, USA and Canada we estimate using both the whole sample and a shortened one of 1946-1990. Because of the many observations for Norway we try several different periods here.

Due to the Second World War, we consider the possibility that a change has occurred in the series not part of the underlying data generating process. For Norway, this seems like a reasonable assumption due to considerable war damages. Perron (1989) discusses three different estimation models[20]:

(M2) $$c_t = \pi \cdot c_{t-1} + \alpha_1 + \alpha_2 \cdot d_2 + \alpha_3 \cdot d_3 + \beta_1 \cdot t + \sum_{j=1}^{k} \gamma_j \cdot \Delta c_{t-j} + e_t$$

(M3) $$c_t = \alpha_1 + \beta_1 \cdot t + \beta_2 \cdot t_2 + c_t^{'} \quad c_t^{'} = \pi \cdot c_t^{'} + \sum_{j=1}^{k} \gamma_j \cdot \Delta c_{t-j}^{'} + e_t$$

(M4) $$c_t = \pi \cdot c_{t-1} + \alpha_1 + \alpha_2 \cdot d_2 + \alpha_3 \cdot d_3 + \beta_1 \cdot t + \beta_2 \cdot t_2 + \sum_{j=1}^{k} \gamma_j \cdot \Delta c_{t-j} + e_t$$

where T_B = time of structural change.
$d_2 = 0$ if $t < T_B$ and 1 otherwise.
$d_3 = 1$ if $t = T_B + 1$ and 0 otherwise,
$t_2 = 0$ if $t < T_B$ and t otherwise.

Model 2 (M2) allows for a one time jump in the series not part of the general data generating process. In case of a unit root, this amounts to α_3 different from zero and in a stationary case to α_2 different from zero. In the third model (M3), we have included the possibility for a change in the general trend in the case of a stationary model. There is, however, not included a term $\alpha_2 \cdot d_2$ which would change the drift in the case of a unit root process. This would

[20] Perron and Vogelsang: Erratum (1993) makes a correction to models M2 and M4. Since the sudden change in intercept at time T_B will occur in the terms $\Delta\ C_{t-j}$ at time T_{B+j} additional dummy variables are needed to neutrilize these terms. This makes no difference in our case since these terms are lost anyway, because of the missing values in our Norwegian data series. When we use model M3, we have to compare it to the significance tables of the erratum.

imply that we would have a break in the intercept in the stationary alternative as in model 2. The model must be estimated in a two-step procedure (see Perron). The final model combines the properties of the two others, allowing a break in the intercept and time trend in case of a stationary process and a one-time-jump and a change in the drift of a unit root process. In these models, Perron makes use of Dickey-Fuller type tests.

A natural break point will be after the Second World War. For the North American countries we estimate the whole sample using the model M1 and M3 since these seem to be the most relevant models here. We have no particular reason to believe in a sudden change in the intercept for these countries. For Norway we estimate models M2 and M4, allowing for a break in the constant term as a natural consequence of the war damages. For Norway this break point is necessary because of the missing years. The missing observations for Norway imply that we lose 2*k+2 (k is the number of included lags of the first differences) observations when we correct for the additional autocorrelation[21]. In the case of a break in the constant term and time trend, the relevant asymptotic probability tables can be found in Perron(1989) (model M3; Perron (1993)).

We shall limit k to be 0,1,2 or 3. With only 36 - 65 observations, this seems to be the highest number of lags supportable. However, the results indicate that this is amply sufficient to remove the autocorrelation in the error term. Looking at the post-war data, only Canada has a γ_3 which is significant and only Germany needs k = 2. In some of the Norwegian models including data for the First World War, k = 3 is necessary. This is also necessary for one of the Canadian models. For the remaining series, we have only one significant lag.

It is recommended (see Campbell and Perron (1991)) to start from the highest value of k and diminish k until the last γ is significant according to its t-value. In the tables at the end of the paper, the test values are reported for k = 0,1,2,3 together with the estimated model for the last significant k.

[21] We lose twice as many observations because we need initial values in both the pre- and the post-war sample. k observations are lost due to lags and one is lost because of first differences of the lagged variables.

6.2 The Bayesian testing framework

In a paper in 1988, Sims criticised the use of unit root tests and advocated the use of a Bayesian model for analysing the stochastic properties of time series. The Bayesian test considered here is the one used by DeJong and Whiteman in several papers. DeJong and Whiteman (1991a) give a good description of the procedure. Here, I shall give a short introduction. My procedure is very similar to theirs, except that I consider the possibility of a break in trend/constant term as comparable with the models of Perron above.

In Bayesian analysis, the observed data are treated as fixed and the likelihood of different parameter values is considered subject to some prior beliefs. The likelihood is calculated according to Bayes theorem

$$(6.5) \quad P(\theta|data) = \frac{P(data|\theta)\cdot P(\theta)}{P(data)}$$

where θ is the set of parameters in which we are interested. Since the data are given and known, P(data) is treated as a constant. P(θ) reflects the prior beliefs of the researcher. The third term on the right hand side corresponds to the traditional inference problem where parameters are treated as fixed and the data as one possible realisation of the underlying structure. This term gives the likelihood of the given set of observed data for different possible combinations of the parameters.

In their analysis DeJong and Whiteman consider the stochastic model

$$(B1) \qquad c_t = \alpha + \beta\cdot t + \varphi_1\cdot c_{t-1} + \varphi_2\cdot c_{t-2} + \varphi_3\cdot c_{t-3} + e_t$$

generating the data. We follow DeJong and Whiteman (DW) using three lags as long as our results from the use of the classical tests show that this seems quite sufficient for removing the autocorrelation from the error term. In the cases where γ_3 is significant in the classical models, we add a term $\varphi_4 \cdot C_{t-4}$ to the equation[22].

In Bayesian analysis, it is common practice to specify a flat prior over the parameter space giving equal weight to all possibilities. This is assumed to represent prior ignorance about the parameters of interest. Normally, this is defended by the observation that the prior is often dominated by the term P(data|θ). Thus, the prior becomes less important. DW also utilise a flat prior in their analysis. This approach is strongly criticised by P.C.B. Phillips (1991) defending the classical models. He specifies what he calls an "objective" prior. However, as many of the following discussants point out, this prior puts an extremely high weight on values of Λ (the inverse of the smallest root) above 1 and it seems very implausible. We follow DW and use a flat prior in this paper.

Assuming a flat prior over the parameters $(\alpha,\beta,\varphi_1,\varphi_2,\varphi_3,(\varphi_4))$, DW show in an appendix that the posterior probability density (P(θ|data)) will be gamma-normal. This means that it can be written on the separable form

$$(6.6) \quad P(\theta|data) \propto P(\sigma|data) \cdot P(\theta'|\sigma,data)$$

in which θ is the vector of parameters to be evaluated, $(\alpha,\beta,\varphi_1,\varphi_2,\varphi_3,(\varphi_4),\sigma)$
σ is the standard error and
θ' is the vector of parameters with σ excluded, $(\alpha,\beta,\varphi_1,\varphi_2,\varphi_3,(\varphi_4))$.

The probability distribution of σ, the "true" standard error, is independent of the estimate of the other parameters. The distribution of the "true" standard error is shown to be

$$(6.7) \quad P(\sigma|data) \propto \sigma^{-\nu-1} e^{\left(-\nu s^2/2\sigma^2\right)}$$

[22]Since the classical models use lags of first differenced data, 3 lags in this model corresponds to 4 lags in the Bayesian set-up.

where ν is degrees of freedom and S is the estimated standard deviation of the estimated autoregressive equation. This distribution is proportional to a gamma distribution. DW show that the transformation of σ, $u = \nu S^2/\sigma^2$ is distributed as χ^2_ν. This can be exploited to make drawings of σ. Finally, the distribution of $P(\theta'|\sigma,\text{data})$ is shown to be normal. Together, this gives us the posterior distribution of θ.

This posterior is not easy to explore analytically but it may be studied using a Monte Carlo simulation framework. As we saw from the gamma-normal distribution above, the distribution of σ did not depend on the other parameters. This can be exploited so that a value for σ can be simulated before the other parameters.

The procedure of the simulation is as follows. First the assumed model for consumption is estimated using ordinary least squares, which gives us an initial value for θ, θ^*. Then a draw is made for the term u above using a χ^2-distribution. The u-value translates to a value for σ when we use the estimated variance from the regression, S^2. Third, the draw of the standard error σ together with the actual $(X'X)^{-1}$ from the data (X is the matrix of explanatory variables) give one simultaneous drawing of the errors in the estimated parameters θ'^* in our equation. We add these errors to the estimated values of the parameters θ'^* to get one simulated outcome for θ'. Eventually, Λ may be calculated from the simulated φ's (elements of θ').

In the actual simulation we use 100,000 independent replications of independent drawings. Since the normal distribution is symmetric around zero, we use both the positive and the negative value of the simulated error, σ. This implies that from one draw of an error, we get two simulated values for the parameter set itself, $\theta'^* + \sigma^*(X'X)^{-1}$ and $\theta'^* - \sigma^*(X'X)^{-1}$. This gives a total of 200,000 simulated values. This should be amply sufficient for our purpose. The resulting distribution over β the time trend parameter and Λ is graphed in the end of the paper.

Along with DeJong and Whiteman, we use a truncated prior for β and Λ. The truncation used by DeJong and Whiteman seems to be suitable for our purposes. We apply a flat prior over the φ's but truncate $\Lambda \in [.55, 1.055)$ and a flat prior $\beta \in [0, 0.016)$. With a Λ above 1.055, the consumption series will have an unreasonable explosive generating process. A β less than zero does not appear to be an attractive part of a unit root process either.

We adopt the procedure used by DeJong and Whiteman by looking at 0.975 as a critical value for Λ in order to conclude that we have a non-stationary series. It means that we interpret the calculated probability for Λ above 0.975 as the probability that the series is non-stationary. DeJong and Whiteman claim that the intention is to be generous to the unit root hypothesis. We prefer to regard it as a compensation for the bias in the estimated coefficients. As we discussed earlier, if the model has a unit root, there is an inherent bias in the estimated trend coefficient when we ordinary least squares are applied. As far as I can see, this testing framework disregards this problem. Whether the compensation should be considered to be enough or not, is hard to tell.

Along with the simultaneous graph of Λ and β we have a graph of the posterior density for Λ and β individually and the density of Λ if β is restricted to equal 0 in the estimation. We also present some of our results in tables 5 and 6. In these tables, we have reported the mean of Λ and β and their standard deviations as they appear in the simulated data after the truncation.

We report the share of the simulated values that we lose due to the truncation. The prior on β and Λ is not implemented in the simulation, but is used to filter the simulated result afterwards. All observations that violate the prior restrictions are removed. This procedure is equivalent to a procedure that had implemented this in the simulation framework, except for the change in number of observations. Since we have so many drawings of the parameters, this should be of no importance.

The number of observations lost due to the truncation is reported and so is the reason for the loss. Of great interest is the probability that Λ is above 0.975, and we give this value with the truncation and without any truncation. Finally, we report the measure Φ/Λ. Φ is the sum of the autoregressive coefficients used in the Dickey-Fuller test, while Λ is calculated directly from the estimated regression. This gives a good indication of the impreciseness inherit in the use of Φ as an approximation for Λ.

In some of our models, our truncation removes a considerable part of the simulated results. This is mainly due to an observation of β less than 0. This implies that the assumed normality of the parameters is a dubious assumption even if one believes in the truncation. The results in these cases should be treated with great caution.

6.3 A comparison of the two tests

As we saw above, the two tests are quite different in nature. The classical tests treat the unit root as the null hypothesis and apply the distribution theory for unit root processes. The Bayesian procedure makes use of the likelihood function which is strictly speaking only valid in the case of stationarity. Even though this test "accepts" values above .975 as an indication of a unit root, this may be too strict on the unit root hypothesis.

On the other hand, the two types of tests make use of a different measure of the persistence in the estimated equation. While the Bayesian test applies the inverse of the exact root of the estimated autoregressive equation, the classical tests make use of the sum of the autoregressive coefficients. If the true process has a unit root, the two measures are similar and should equal 1. Also, if we have the AR(1)-model the two procedures should be equal. However, if the true process is stationary and we have several lags, the two may diverge. The deviation is considerable in some of our models.

As long as our prime interest is the size properties of our test, this might not be a serious question. However, when we deal with power issues, this is more likely to be of importance. The interesting question is in what terms should the alternative hypothesis be specified. If we think that the sum of autoregressive coefficients is a good way of specifying the alternative, the classical procedure is correct. However, if we think the highest root of the process to be the relevant unit of measure of our alternative, the imprecise root measure used by the classical tests will erode some of their power. The highest root seems to be the appropriate measure to use. It seems to better represent the persistence of the error structure in the series. DeJong and Whiteman (1991c) emphasise this.

As a consequence, the classical tests may appear too liberal and the Bayesian test somewhat strict on the unit root hypothesis.

7 ESTIMATION RESULTS

We divide the comments on the estimation results into several sections. First, we consider the period after World War II and make a comparison between all the countries. Second, we include the data for the depression in the early thirties for Norway and the North-American countries. Finally, we turn our attention towards the Norwegian series more fully. The results are given in the tables at the end of the paper where tables 1 to 4 contain the results from the classical tests, while 5 and 6 give the main information for the Bayesian test. At the end, we have included the graphs of the Bayesian test.

7.1 Section I: Unit roots in the post-war data

From the results of the classical tests in table 1, we can not decisively reject the null hypothesis for any of the series. Sweden, Germany and Japan seem to fit the unit root hypothesis very well with sharp estimates of π close to 1 and with all the test values far from being significant at a 90% level. For Germany, it does not matter if we just rescale the data from before 1960 or use a dummy variable. The results are influenced by this to a very small extent, and we only treat the rescaling alternative.

The four other series have higher test values and Norway, UK and Canada fail the DF-ρ-test at a high level. This reflects the fact that all these series have an estimated π in the range between 0.66 and 0.77 with rather high standard errors which surely question the power of these tests against interesting stationary alternatives. The unit root hypothesis is rather dubious for these series.

When we look at the additional autocorrelation in the series (the parameter γ), we find that this is overall positive. For Germany and UK, γ_1 is about 0.5 and is highly significant. The North-American countries are close to .4, while Norway gives .3. The remaining two show insignificant signs of additional autocorrelation[23]. Considering the expectation of a negative moving average term from the combined hypothesis of rational expectation, permanent income and durable goods, the results are not too favourable to this hypothesis. In table 4, we give the same autocorrelation parameters when we have restricted our models to contain a unit root. That is π is set to 1. We do this in order to show that the positive additional correlation is not an artifact of our estimated model failing to show unit roots. The correlation is still positive, but in some cases less than before. The positive autocorrelation gives support to the Life Cycle model.

[23]In both of these series γ_1 was positive and above .2, but not significant.

In table 5, the Bayesian tests show a somewhat different pattern. The differences are partly caused by the truncation working differently for the series, in particular the demand that $\beta_t>0$. However, as can be seen from the tables, Φ is an imprecise measure of Λ.

Actually, the only series for which the unit root is not rejected is Canada, for which the test predicts that there is a 6% chance that Λ is above .975. This is rather surprising if we compare with the results from the classical tests. The main reason is that the Λ from the estimated OLS-regression is highly underestimated, using Φ for Canada. For USA, giving very similar results to Canada when we applied the classical tests, the reverse is true. Λ is highly overestimated by Φ, and the unit root seems very implausible for this series. The main cause for truncation here is the $\Lambda>.55$ restriction.

UK and Norway are both rejected at the normal 5%-level, although not as convincingly as for the USA. If we look at the untruncated results, we no longer reject the unit root hypothesis for the Norwegian series, while we just reject it in the UK data.

The three series that showed results close to unit root when we looked at the classical tests, all suffer from a strong truncation of negative trend coefficients. The truncation is between 15% and 35% of the simulated sample and this may distort the results considerably. The assumption of normality of the error terms in the original regression may be a dubious assumption. Caution should be exercised when we interpret these results.

If we take the estimated models literally, we reject the unit root for Germany and Japan strongly. Looking at the no truncation model the hypothesis is however, accepted at a 10%-level (20% for Japan). Estimating the two series restricting the trend coefficient to zero, gives a probability of $\Lambda>.975$ of .064 for Japan and .032 for Germany. It seems that Germany needs a negative trend coefficient to accept the unit root hypothesis, while for Japan this is not the case.

In the case of Sweden, we have very much the same pattern as for Japan and Germany. The unit root is not as strongly rejected and gets a probability of .14 in the case of no truncation. With the trend coefficient restricted at zero, the probability rises to .344, making the rejection of a unit root implausible.

7.2 Section II: Unit roots when the pre-war data are included in the sample

First of all, we want to find out whether the behaviour of the consumption path during the great depression makes any difference to our previous conclusions and, secondly, we try to find out whether any changes are general or if there are differences between the separate series. For this period, we have estimated two different classical models for every series, one with no break in trend and a second one with a break in trend in 1946 (models M1 and M3 for the North American countries). For the Norwegian consumption series we keep a break in the constant term in both models (models M2 and M4).

First, looking at the models with no trend break, we see from table 2 that the unit root hypothesis now breaks down for USA and to some extent for Canada. All the test values for USA are significantly different from zero at a 95% level (most on a 99% level), leading to a clear rejection of the unit root hypothesis. A drop in the estimated π is mainly the cause of the change in conclusion, but the estimated standard deviation of π falls a little as well. For Canada the estimated π increases, but standard deviation falls so much that the Dickey-Fuller tests give the same conclusions at an almost equally high level as for the USA. The Phillips-Perron tests are not significant at a 90%-level. These differences among the tests stem from the different treatment of the additional autocorrelation of course and γ_1 is as high as .525. The unit root hypothesis is therefore dubious for Canada. Allowing for a break in the trend does apparently not alter these conclusions and the estimated break in trend seems to be insignificant for both countries. The test values are much the same, but the estimated π falls and has a larger standard error. Once again, the results seem to fit the Life-Cycle model of consumption much better than the representative agent model of Hall.

For Norway, the results are rather similar to the post-war sample alone. Although the estimated change in the constant term is negative and significant, there are only small changes in the remaining estimated parameters and the test values show a very similar pattern although somewhat lower. The DF-ρ test is lower and for k=1 does not reject the unit root at a 95%-level. It is very close to the 90%-level in both models. The change in trend is not significant. Thus, according to the classical tests, there seems to be no reason to alter the conclusions for Norway and the unit root hypothesis has not been weakened by including this period.

In the Bayesian test, the unit root is rejected at a very high significance level. For Canada, Λ shows a drop from the results from the post-war sample causing the rejection. For the US data the change from the post-war sample mainly results in a smaller standard deviation. The share of the simulated values lost due to truncation is so small that this should not cause any problems for this conclusion. One interesting observation is the large disagreement between Φ and Λ in the US data. While in the post-war sample the Φ highly overstates the Λ, the reverse is true when the pre-war sample is included. However, Λ is remarkably stable across the two sub-samples.

The Norwegian series shows a similar pattern when the period of the great depression is included as in the post-war sample alone. The unit root seems less plausible and is also rejected in the case of an untruncated model. The rejection is not as clear as for the North American countries.

7.3 Section III: Discussing the Norwegian series more fully

As can be seen from the plot of the first difference, the Norwegian data show large fluctuations at the end of World War I and in the following years. Using ordinary least squares estimation procedure, this period is likely to be highly influential on the estimation results when the sample includes this period. We therefore test for unit roots using different periods of time and 1914 as a cut point in addition to the Second World War[24]. As we shall see, the assumption that the period after the First World War is highly influential is confirmed by our estimated models.

First, I like to comment on the pre 1939 data separately and then discuss the results of the models using data from both periods.

Using only one trend, the period 1865 to 1939 questions the unit root hypothesis. We need three lags in the augmented regression in order to remove the autocorrelation and the tests based on normalised bias both reject the hypothesis. Also the PP-τ test makes a clear rejection, while the DF-τ at two and three lags not quite reach the significant values. When we limit the period to 1865-1914, we see that some of these problems disappear. Now, we need only one lag to remove the autocorrelation and the calculated test values are not significant at a 90%-level. The only exceptions are high values of DF-ρ at two and three lags.

[24]The Perron framework does not allow more than one break in the sample used for estimation. In our model, this makes a sample period of 1865-1914, 1926-1939 and 1946-1991 in one estimation impossible.

The results are rather similar to the post-war sample although the estimates are somewhat sharper. The estimated trend coefficient is lower than in the post-war sample indicating the need for a break in the trend looking at the two sub-samples simultaneously. Still with a $\pi =$ .785, the power of the test is not impressive.

In the models in which we consider both the pre- and post-war data, much of the same pattern repeats itself. When the erratic period at the end of the First World War is included in the sample, we need three lags to remove the autocorrelation and the estimated π is lower. However, the standard error increases accordingly making a clear rejection difficult. The DF-ρ seems to reject when a break in trend is allowed. When the sample period is 1915 to 1991, $\pi = .56$ gives very little confidence to the unit root hypothesis. When leaving out the break in trend which is not quite significant, the unit root is not rejected.

Combining the two periods before World War I and after World War II seems to give some support to the hypothesis of a unit root. The test values do not reject the hypothesis at a 95%-level for the τ-test, but for DF-ρ at 1-3 lags the unit root is rejected. The estimated π's are higher than when the period around the First World War is included, but still the power is not good. If we leave out the break in the trend term, the unit root seems highly plausible. Both the trend terms appear significant, however. The main conclusion for the Norwegian data is that although highly dubious, on the basis of the classical tests, we can not completely exclude a unit root. The period 1915-1925 shows such low values for π that no interesting stationary alternatives can be excluded.

The γ-parameters still show a clear sign of positive additional autocorrelation giving support to the Life Cycle model. When the period around the first World War is included this evidence is less clear however. Is this case these parameters are highly dominated by this erratic period.

Applying the Bayesian test procedure on the same sub-samples, we reject the unit root strongly in all the cases and it does not seem to add much to the classical tests. The Φ/Λ is quite stable across different sub-samples.

8 MAIN CONCLUSIONS

The purpose of this paper is to investigate the general stochastic properties of aggregate consumption expenditure. The main question is if this variable shows any signs of a unit root or not. As was shown in Hall (1978), the combined hypothesis of Permanent Income, Representative Agent and Rational Expectations should give an aggregate consumption series containing a unit root, or at least be very close to one. We show that replacing this model with a Life Cycle formulation gives us a stationary model. This model is so close to a unit root process that it seems indistinguisable from the non-stationary process, at least in our length of data.

However, the results do not show any clear pattern across countries. Although some countries (notably Japan and Sweden) show clear signs of a possible unit root in the series, the period of the Great Depression shows a clear rejection for the North American countries. The Bayesian test makes the unit root quite unlikely in most of our series.

Perhaps one could argue that the period of the great depression was a period of exceptional events making severe restrictions on household consumption behaviour, and that one should not reject the unit roots in more "normal" times. However, when we take a closer look at the data after World War II, we do not get a convincing case for the unit root hypothesis. Using the classical tests, four out of seven countries have imprecise estimates giving the tests low power against stationary alternatives. Still, they give low confidence to the null hypothesis of a unit root and is close to rejection in some. When we apply a Bayesian testing framework, we reject the unit root in all but the Canadian series. We can explain this by a bias due to the truncation for Japan and Sweden, but although the German consumption series appear to be close to a unit root in the classical framework, the Bayesian testing procedure shows that we can only support this if we accept a negative coefficient for the time trend. Combined with the unit root properties, this is not a very satisfactory assumption. We are therefore by no means willing to accept unit roots in aggregate consumption expenditure as a "stylised" fact.

The two theoretical models discussed in this paper had very different implications for the additional autocorrelation parameters in the model. While the Hall model indicated no or negative additional autocorrelation, the Life Cycle model implied positive parameter values for these additional lags. The parameters of the lagged first differences of the consumption series show signs and sizes that are contradictive to the Hall model. In all of the estimated models, the additional autocorrelation is positive and mostly quite significant. The only models where this is not the case, are the Norwegian models containing the erratic period of the First World War. In the post war sample, all models have a γ_1 of size .30 or larger. Germany does have a γ_2 of -.33 but the first is as high as .51. Turning to the data from before 1940, the results are equally convincing.

According to the model that includes durable goods, we should expect to find evidence of negative sign on these coefficients. Time aggregation bias could balance this effect to some extent, but cannot explain that we estimate such high positive values.

On the other hand, the results seem to be quite in agreement with what we would expect from a Life Cycle model, perhaps in combination with aggregation bias. The prediction of this model is clearly in favour of positive values for these parameters. Thus, the life cycle model seems to be more in line with the stochastic properties of actual observed consumption in many OECD-countries. Both models, however, fail to explain why we do not find a unit root (or close to) behaviour in the consumption data. This remains a topic for further research.

APPENDIX: DATASOURCES

Germany: *Consumption data:* Volkswirtschaftliches Gesamtrechnungen, 1990 and 1950 bis 1990.
Population data: International Financial Statistics.

Norway: *Consumption data:* Nasjonalregnskap 1865-1961, Printouts from The National Account.
Population data: Historisk Statistikk and Statistisk Årbok.

Sweden: *Consumption data:* Nationalräkenskapen 1970-1989
Population data: International Financial Statistics.

UK: *Consumption data:* Printout of the Central Statistical Office central database.
Population data: International Financial Statistics.

Japan: *Consumption data:* Printout from the Economic Planning Agency.
Population data: International Financial Statistics.

Canada: *Consumption data:* National Income and Expenditure Accounts, Annual Estimates 1979-1990 and 1926-1986.
Population data: Demographic Yearbook, United Nations. (1989, 1979 and Special Issue 1979). Canadian Economic Observer April 1992. Vital Statistics 1951.

USA: *Consumption data:* The National Income and Products Account.
Population data: Historical Statitics of the United States. Colonial times to 1970. Statistical Abstract of the United States, 1991.

TABLE 1: RESULTS FOR THE POSTWAR DATA OF THE CLASSICAL TESTING PROCEDURE.

MODEL 1 (M1): $C_t = \pi C_{t-1} + \alpha + \beta t + \Sigma^{k}_{j=1} \gamma_{t-j} \Delta C_{t-j}$

DF = Dickey-Fuller, τ_i = Tautest with i lags for autocorrelasion adjustments.
PP = Phillips-Perron, ρ_i= The Rhotest using i lags for adjustment

Crit. val. for 25, 50	5% level	τ: -3.60 , -3.50 -3.45	ρ: -17.9 , -19.8 -20.7
and 100 # of obs.	10%level	τ: -3.24 , -3.18 -3.15	ρ: -15.6 , -16.8 -17.5

		ESTIMATION RESULTS.							TEST VALUES.							
	Sample	π	α	β	γ_1	γ_2	γ_3		τ_0	τ_1	τ_2	τ_3	ρ_0	ρ_1	ρ_2	ρ_3
Germany	1950-90	.9398	.5684	.00097	.5095	-.3283	-	DF	-1.82	-2,15	-1.85	-1.50	-2.21	-4.09	-2.79	-2.42
		(.0326)	(.2781)	(.00119)	(.1565)	(.1552)		PP		-1.81	-1.80	-1.83		-2.47	-2.47	-2.39
Norway	1946-91	.7697	1.870	.0060	.2980	-	-	DF	-1.77	-2.95	-2.91	-2.18	-9.4	-30.7	-62.9	-46.2
		(.1208)	(.9607)	(.0033)	(.1660)			PP		-2,28	-2.18	-2.15		-12.6	-12.5	-12.0
Sweden	1950-90	.9517	.5196	.00063	-	-	-	DF	-.779	-1.36	-1.39	-1.22	-1.93	-4.14	-5.25	-4.29
		(.0620)	(.6351)	(.00124)				PP		-.977	-1.07	-1.10		-1.16	-1.19	-1.26
UK	1948-90	.6650	2.503	.0073	.5200	-	-	DF	-1.63	-3.08	-2.28	-1.91	-6.99	-28.6	-26.8	-34.5
		(.1087)	(.8084)	(.0023)	(.1651)			PP		-2.06	-2.10	-2.03		-9.88	-10.1	-9.56
Japan	1955-90	.9696	.4028	-.00027	-	-	-	DF	-1.02	-1.20	-1.35	-2.26	-1.06	-1.62	-1.83	4.22
		(.0298)	(.3118)	(.0016)				PP		-1.06	-1.06	-1.07		-1.16	-1.19	-1.26
Canada	1946-90	.7521	1.909	.0068	.3520	-.0032	.3901	DF	-2.48	-2.07	-1.90	-2.54	-9.30	-10.3	-10.6	-32.3
		(.0851)	(.6485)	(.0024)	(1486)	(.1288)	(.1262)	PP		-1.85	-1.93	-2.13		-7.07	-7.54	-8.82
USA	1946-90	.7556	.2703	.0052	.3117	-	-	DF	-2.40	-2.57	-2.31	-2.33	-9.14	-14.9	-14.7	-19.1
		(.0861)	(.0909)	(.0018)	(.1454)			PP		-2.50	-2.35	-2.53		-10.6	-10.3	-10.9

TABLE 2. RESULTS OF THE CLASSICAL TESTING PROCEDURE WHEN THE PERIOD OF THE DEPRESSION IS INCLUDED IN THE SAMPLE.

MODEL 2 (M2) $C_t = \pi C_{t-1} + \alpha_1 + \alpha_2 d_2 + \alpha_3 d_3 + \beta t + \Sigma_{j=1}^{k} \gamma_{t-j} \Delta C_{t-j}$

MODEL 3 (M3) $C_t = \alpha_1 + \beta t + \beta_2 t_2 + C'_t,$ $C'_t = \pi C'_{t-1} + \Sigma_{j=1}^{k} \gamma_{t-j} \Delta C'_{t-j}$

MODEL 4 (M4) $C_t = \pi C_{t-1} + \alpha_1 + \alpha_2 d_2 + \alpha_3 d_3 + \beta t + \beta_2 t_2 + \Sigma_{j=1}^{k} \gamma_{t-j} \Delta C_{t-j}$

ESTIMATION RESULTS. / TEST VALUES. (Definitions and critical values as in table 1.)

	Sample	π	α_1	α_2	β_1	β_2	γ_1	γ_2	γ_3		τ_0	τ_1	τ_2	τ_3	ρ_0	ρ_1	ρ_2	ρ_3
Canada (M1	1926-90	.8056	1.496		.00549		.525			DF	-2.17	-4,27	-4.49	-5.77	-7.72	-25.78	-26.89	-43.71
		(.0455)	(.3485)		(.00126)		(.0995)			PP		-2.50	-2.59	-2.60		-10.41	-11.38	-11.46
(M3)		.1221	1.003		.00342	-.599E-5	.3565	.1249	.2766	DF	-5.19	-5.68	-5.25	-5.00	-34.7	- 55.4	-75.3	-217.7
		(.1755)	(.4347)		(.0015)	(.811-5)	(.1461)	(126)	(.1131)									
USA (M1)	1929-90	.6043	.4272		.00846		.3947			DF	-6.30	-5.59	-6.08	-3.59	-23.9	-39.2	-67.4	-36.5
		(.0708)	(.0749)		(.0015)		(.0894)			PP		-6.10	-6.04	-6.05		-26.5	-27.2	-27.0
(M3)		.3807	.4179		.00867	-.00021				DF	-5.09	-5.14	-5.15	-4.22	-37.2	-56.3	-74.5	-62.6
		(.1217)	(.0723)		(.00144)	(.00083)												
Norw. (M2)	1926-39	.7608	1.994	-.0499	.0062		.3019			DF	-1.66	-2.43	-2.38	-1.88	-9.03	-19.5	-28.7	19.8
	+1946-91	(.0984)	(.8069)	(.0256)	(.0026)		(.1403)											
(M4)		.7257	2.383	-.1628	.0057	.00150	.3287			DF	-1.58	-2.51	-2.59	-1.99	-9.35	-23.3	-49.7	-43.7
		(.1092)	(.9591)	(.1513)	(.0027)	(.0020)	(.1453)											

TABLE 3. RESULTS OF THE CLASSICAL TESTING PROCEDURE FOR DIFFERENT SUBSAMPLES OF THE NORWEGIAN SERIES.

ESTIMATION RESULTS. TEST VALUES. (Definitions as in table 1.)

	Sample	π	α_1	α_2	β_1	β_2	γ_1	γ_2	γ_3		τ_0	τ_1	τ_2	τ_3	ρ_0	ρ_1	ρ_2	ρ_3
(M2)	1865-39	.9426	.5340	.0142	.00093		-.0083	-.4042	.2394	DF	-2.60	-2,19	-2.00	-1.90	-9.52	-7.60	-4.42	-5.52
	+1946-91	(.0302)	(.2736)	(.0143)	(.00030)		(.0956)	(.0850)	(.0948)									
(M4)	1865-39	.6883	2.864	-.3421	.00401	.00415	.1962	-.2560	.3395	DF	-4.31	-4.18	-1.79	-3.04	-39.6	-47.7	-14.3	-48.9
	+1946-91	(.1025)	(.9376)	(.1382)	(.0013)	(.0016)	(.1220)	(.1006)	(.1001)									
(M2)	1865-14	.9694	.2873	.0112	.00047		.1924			DF	-1.81	-1.45	-1.62	-1.38	-3.57	-3.48	-4.11	-3.05
	+1946-91	(.0210)	(.1898)	(.0148)	(.00041)		(.1018)											
(M4)	1865-14	.7772	2.058	-.2488	.00245	-.00336	.3181			DF	-2.13	-3.21	-3.60	-3.24	-13.5	-30.1	-55.7	-60.4
	+1946-91	(.0693)	(.6377)	(.0909)	(.00079)	(.00116)	(.1069)											
(M2)	1915-39	.7876	1.796	-.0381	.00532		.0028	-.3950	.2405	DF	-3.63	-3.37	-2.41	-2.54	-21.8	-20.5	-8.54	12 .1
	+1946-91	(.0837)	(.6974)	(.0282)	(.00217)		(.1266)	(.1034)	(.1174)									
(M4)	1915-39	.5598	3.916	-.3730	.00764	.00399	.2296	-.2574	.3309	DF	-4.16	-4.15	-1.11	-2.26	-34.2	-47.1	-8.07	-40.4
	+1946-91	(.1950)	(1.781)	(.2608)	(.0028)	(.00309)	(.2007)	(.1480)	(.1361)									
(M1)	1865-39	.7143	2.627		.0037		.1048	-.3487	.3343	DF	-3.99	-3.68	-1.49	-2.33	-29.7	-33.3	-7.95	-22.3
		(.1225)	(1.121)		(.0015)		(.1511)	(.1205)	(.1237)	PP		-4.13	-3.82	-3.99		-30.6	-25.6	-28.3
(M1)	1865-14	.7850	1.986		.0024		.3527			DF	-1.94	-2.68	-3.25	-3.06	-7.63	-15.9	-31.0	-39.3
		(.0802)	(.7377)		(.0009)		(.1419)			PP		-2.26	-2.42	-2.48		-9.75	-11.3	-11.9

TABLE 4. ADDITIONAL AUTOCORRELATION STRUCTURE WHEN A UNIT ROOT IS IMPOSED ON THE DATA.

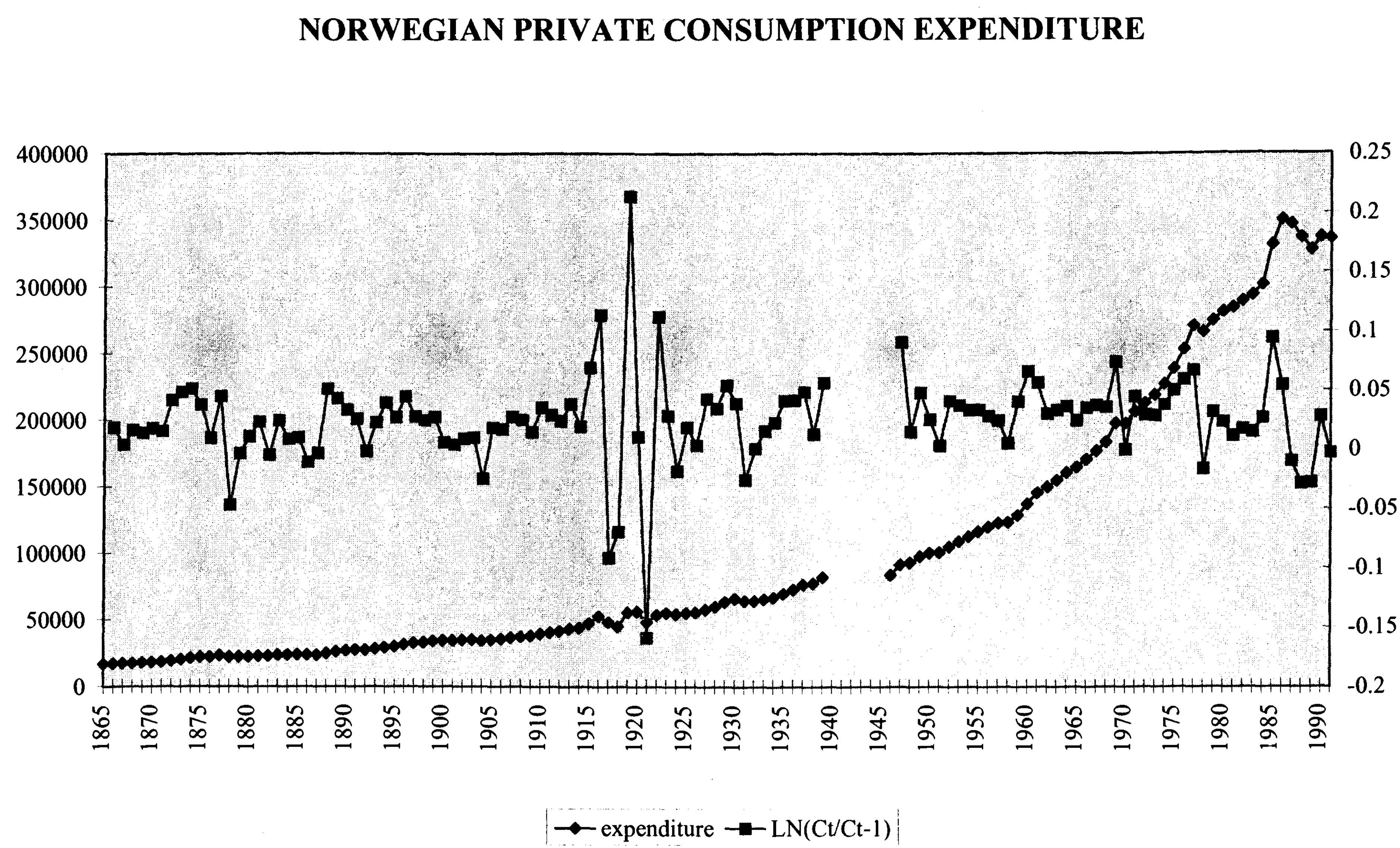
	γ_1	γ_2	γ_3
Germany	.544 (.161)	-.344 (.160)	
Norway	.144 (.150)		
Sweden	.227 (.151)		
UK	.270 (.159)		
Japan	.207 (.178)		
Canada	.373 (.144)	-.054 (.145)	.289 (.144)
USA	.241 (.155)		

TABLE 5. RESULTS FROM THE ESTIMATION OF A BAYESIAN UNIT ROOT MODEL.

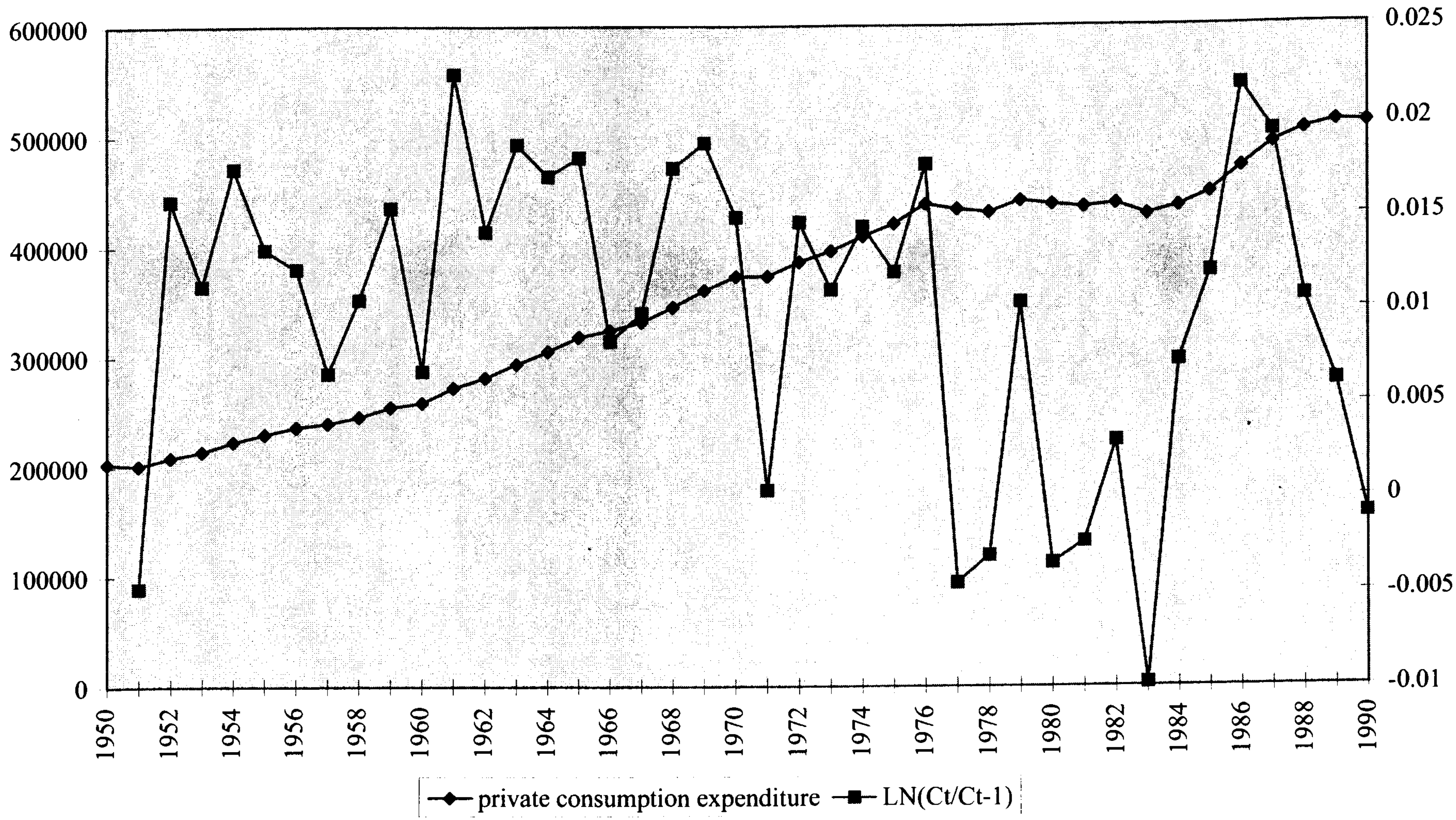
PREWAR RESULTS	Λ-mean	Λ-st. dev.	β-mean	β-st.dev.	share out of sample	# of Λ > 1.055	# of Λ < .55	# of β > .016	# of β < .0	P(Λ> .975 | prior)	P(Λ> .975| no prior)	Φ/Λ
GERMANY	.907	.0376	.0014	.0009	.2108	461	7	0	42145	.0016	.1174	1.01
NORWAY	.797	.1000	.0078	.0037	.0884	3576	4002	5929	7557	.0244	.0671	.93
SWEDEN	.843	.0805	.0016	.0010	.1514	5246	108	0	30095	.0205	.1422	1.10
UK	.783	.1082	.0075	.0030	.0522	1784	6770	899	2289	.0337	.0478	.95
JAPAN	.906	.0511	.0017	.0012	.3577	1397	3	0	71504	.0067	.2381	1.01
CANADA*	.882	.0642	.0068	.0024	.0047	478	23	34	578	.0611	.0652	.85
USA	.736	.0990	.0052	.0021	.0808	425	14649	1	1508	.0068	.0138	1.31
PREWAR RESULTS												
CANADA*	.772	.0670	.0092	.0016	.0095	1	1888	6	1	.0000	.0000	.88
USA	.740	.0636	.0111	.0018	.0192	0	2795	1042	0	.0000	.0000	.65
NORWAY 1926-91 (B2)	.781	.0929	.0084	.0032	.0435	441	4573	2642	1443	.0099	.0184	.89

Table 6. Estimates of autocorrelation parameters when a unit root is imposed on the system. The models are based on M1 and M2.

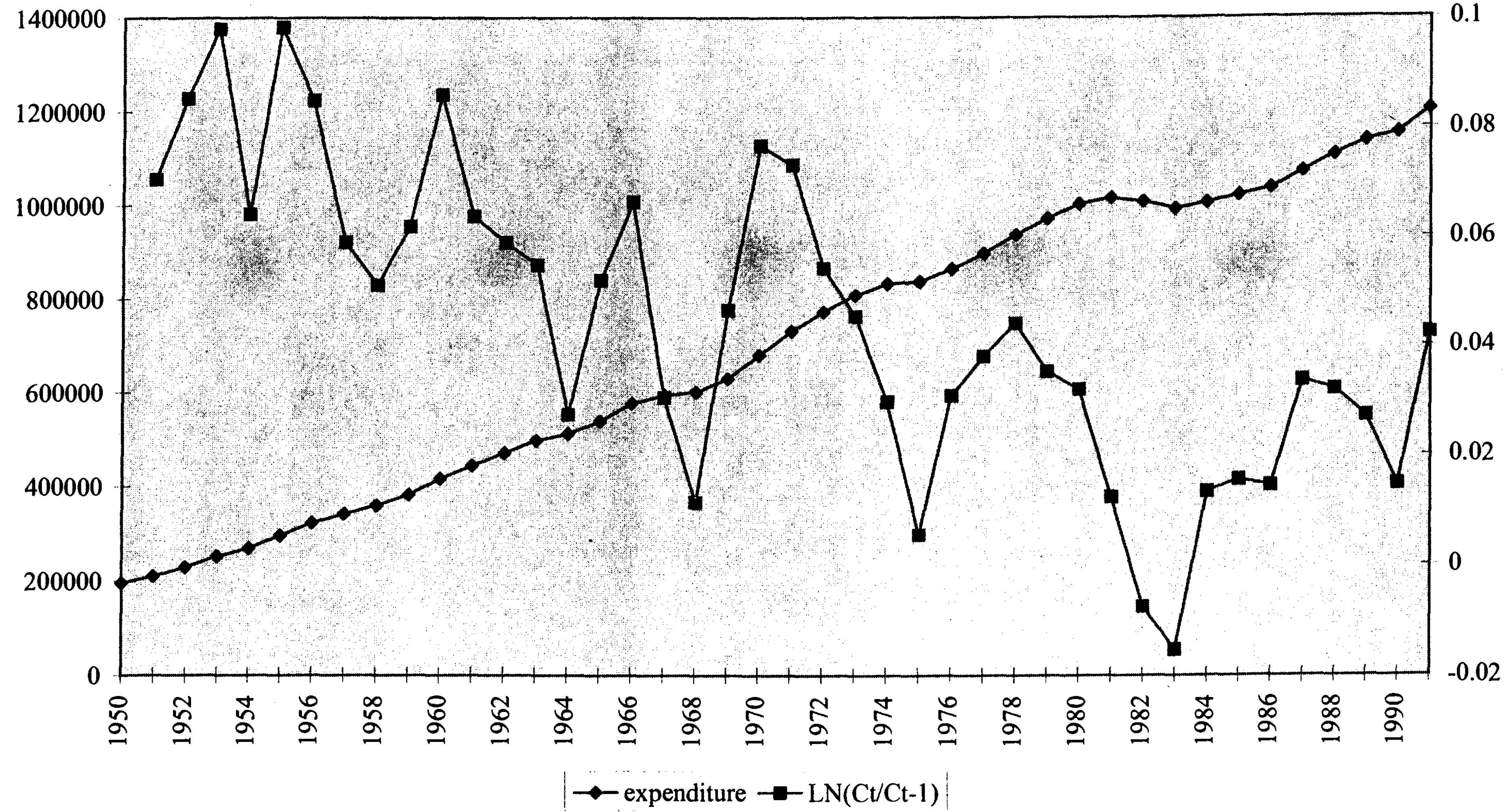
	γ_1	γ_2	γ_3
Germany 1950-90	.506 (.178)	-.306 (.191)	-.031 (.177)
Sweden 1950-90	.278 (.178)	.047 (.182)	-.128 (.162)
UK 1948-1990	.348 (.164)	-.304 (.169)	-.105 (.174)
Japan 1955-1990	.208 (.183)	-.108 (.186)	.261 (.189)
Norway 1946-1990	.219 (.164)	-.083 (.164)	-.173 (.162)
Norway 1865-1914 1946-1990	.233 (.110)	-.001 (.112)	-.153 (.109)
Norway 1926-1939 1946-1990	.223 (.148)	-.068 (.149)	-.196 (.146)
Canada 1946-1990	.232 (.157)	-.100 (.137)	.288 (.133)
Canada 1926-1990	.338 (.206)	.172 (.226)	.662 (.197)
USA 1946-1990	.223 (.159)	-.119 (.132)	-.128 (.130)
USA 1929-1990	.520 (.129)	.227 (.137)	.164 (.119)

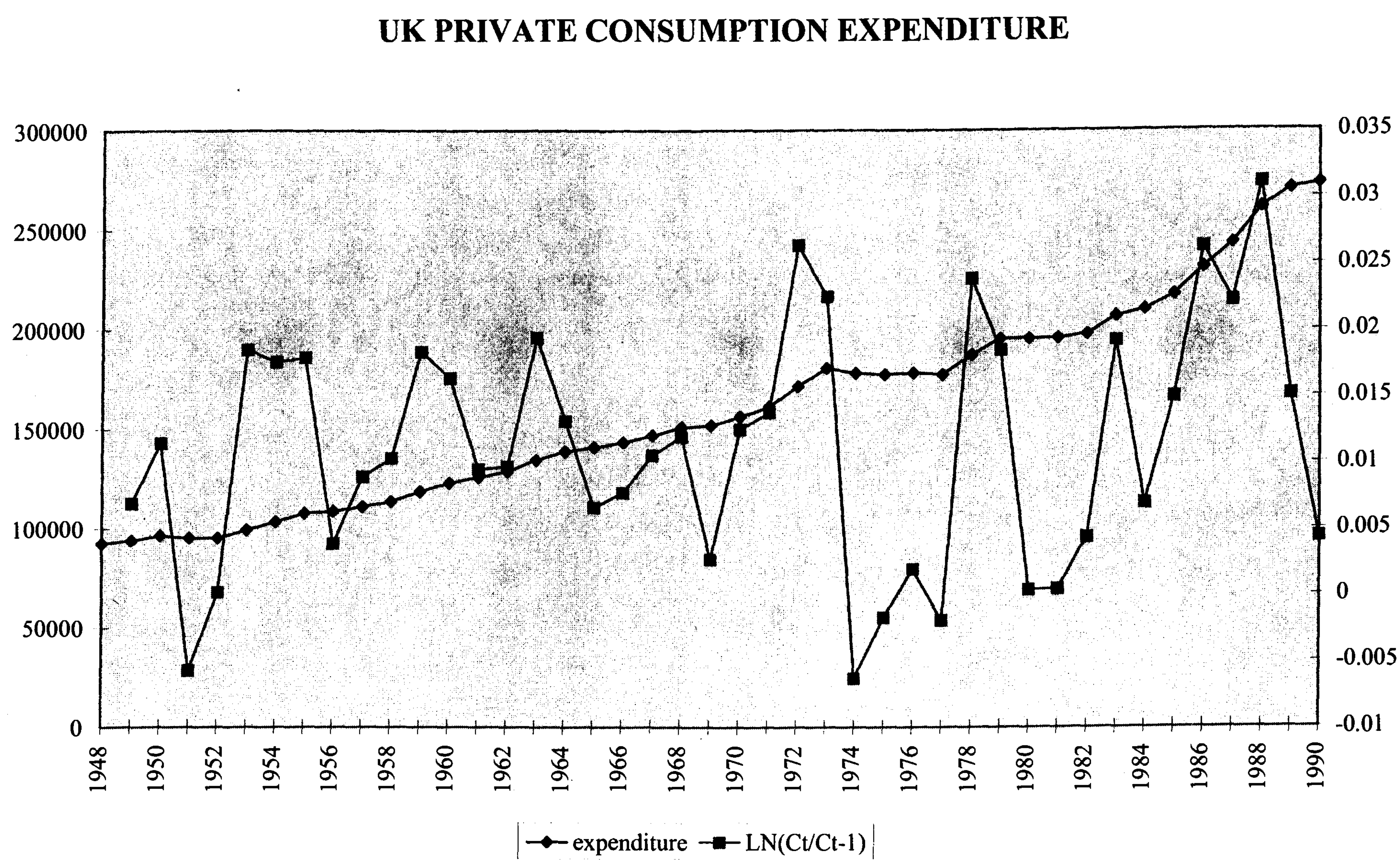

NORWEGIAN PRIVATE CONSUMPTION EXPENDITURE
400000
350000
300000
250000
200000
150000
100000
50000
0
0.25
0.2
0.15
0.1
0.05
0
-0.05
-0.1
-0.15
-0.2
1865
1870
1875
1880
1885
1890
1895
1900
1905
1910
1915
1920
1925
1930
1935
1940
1945
1950
1955
1960
1965
1970
1975
1980
1985
1990
expenditure
LN(Ct/Ct-1)

SWEDISH PRIVATE CONSUMPTION EXPENDITURE.

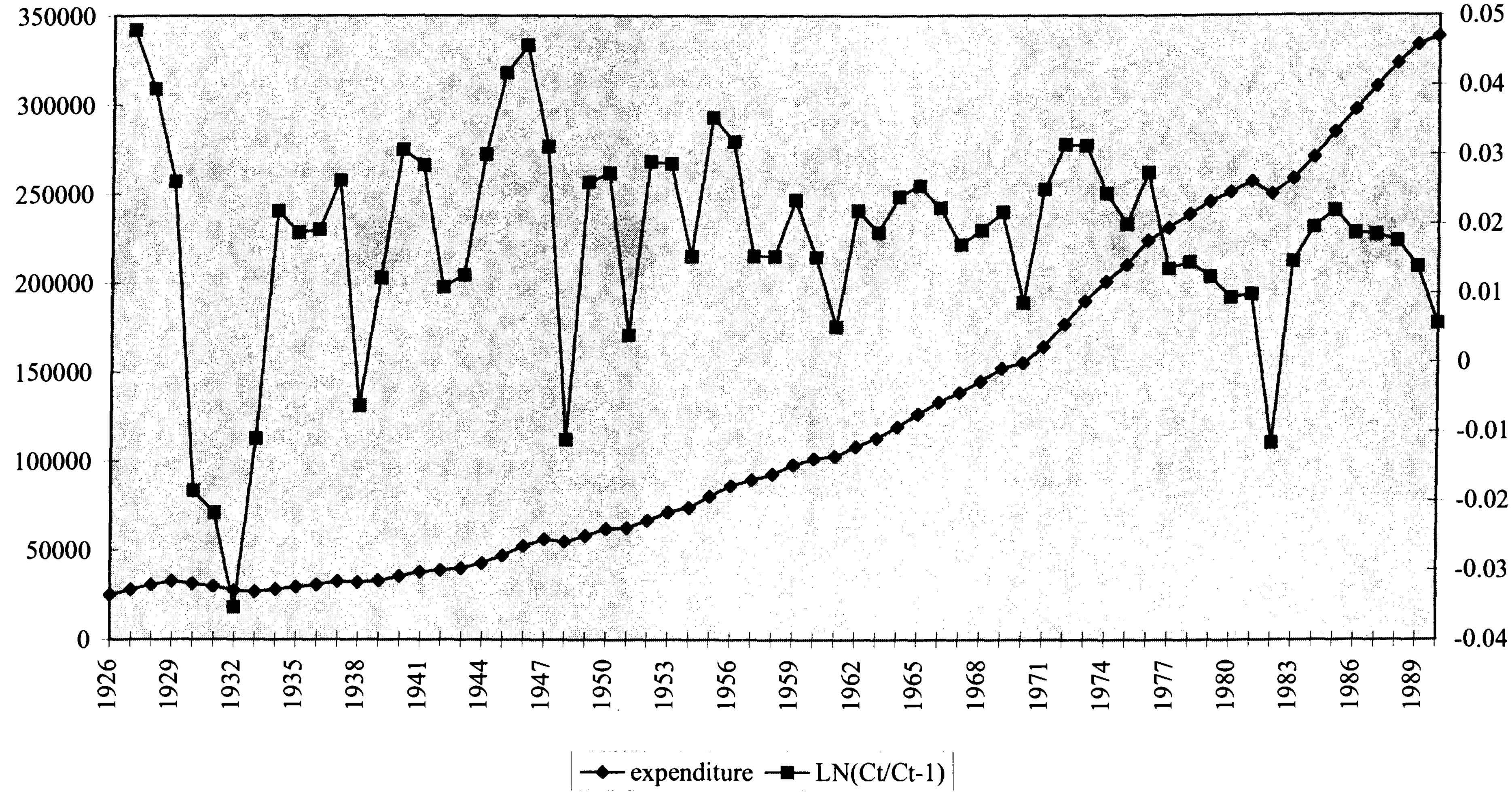


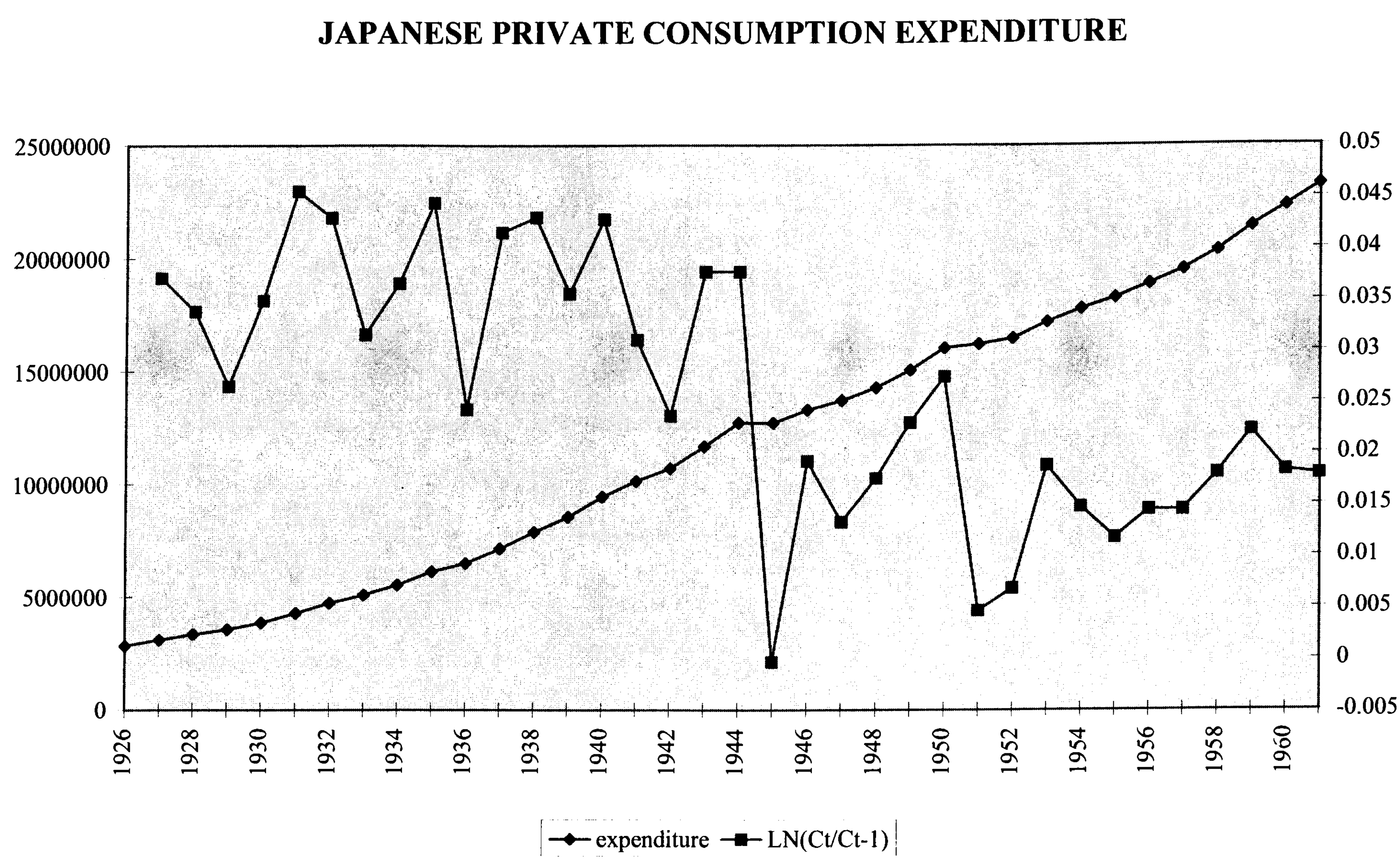
GERMAN PRIVATE CONSUMPTION EXPENDITURE



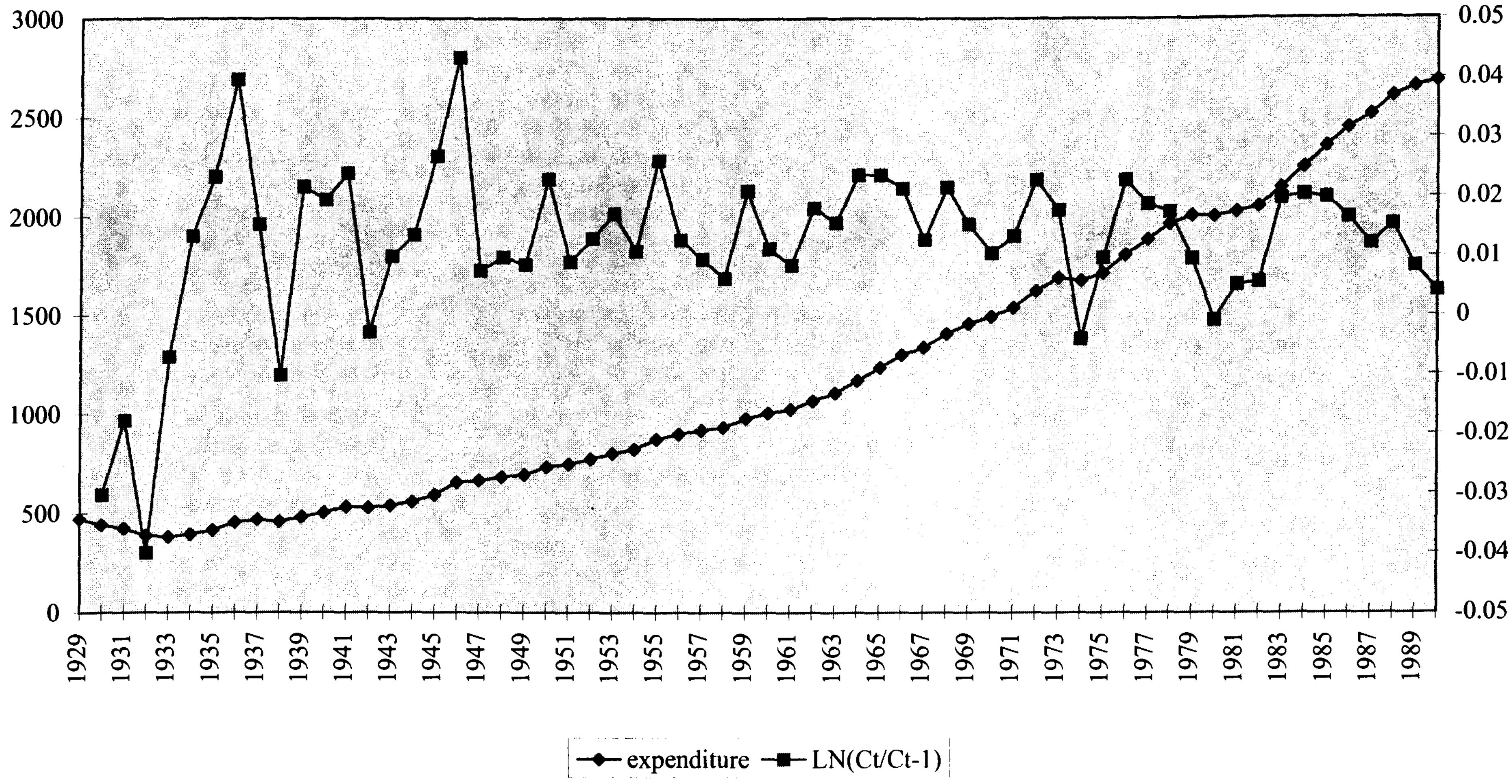

UK PRIVATE CONSUMPTION EXPENDITURE
300000
250000
200000
150000
100000
50000
0
0.035
0.03
0.025
0.02
0.015
0.01
0.005
0
-0.005
-0.01
1948
1950
1952
1954
1956
1958
1960
1962
1964
1966
1968
1970
1972
1974
1976
1978
1980
1982
1984
1986
1988
1990
expenditure
LN(Ct/Ct-1)

CANADIAN PRIVATE CONSUMPTION EXPENDITURE

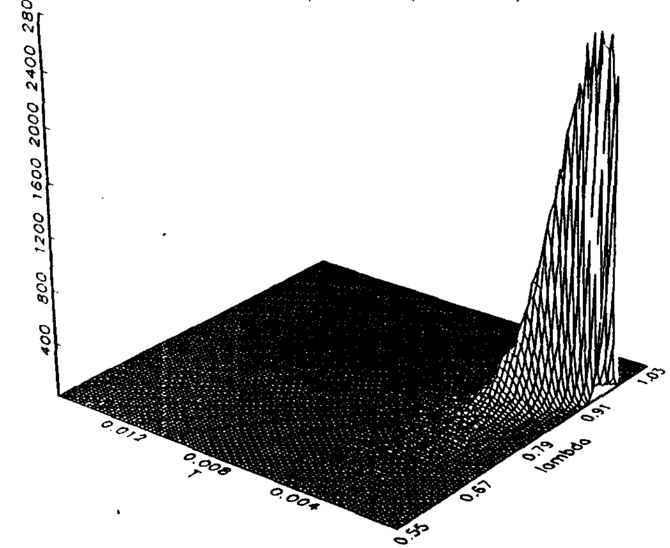



JAPANESE PRIVATE CONSUMPTION EXPENDITURE
25000000
20000000
15000000
10000000
5000000
0
0.05
0.045
0.04
0.035
0.03
0.025
0.02
0.015
0.01
0.005
0
-0.005
1926
1928
1930
1932
1934
1936
1938
1940
1942
1944
1946
1948
1950
1952
1954
1956
1958
1960
expenditure
LN(Ct/Ct-1)

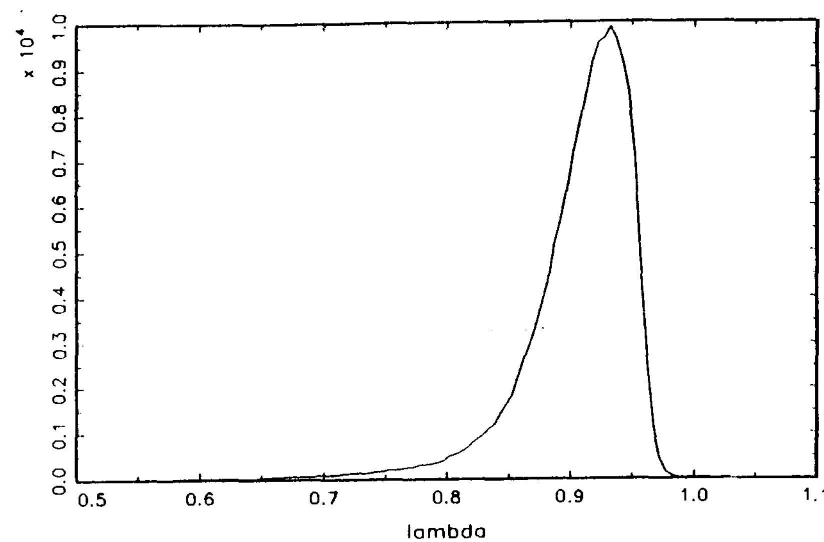
US PRIVATE CONSUMPTION EXPENDITURE



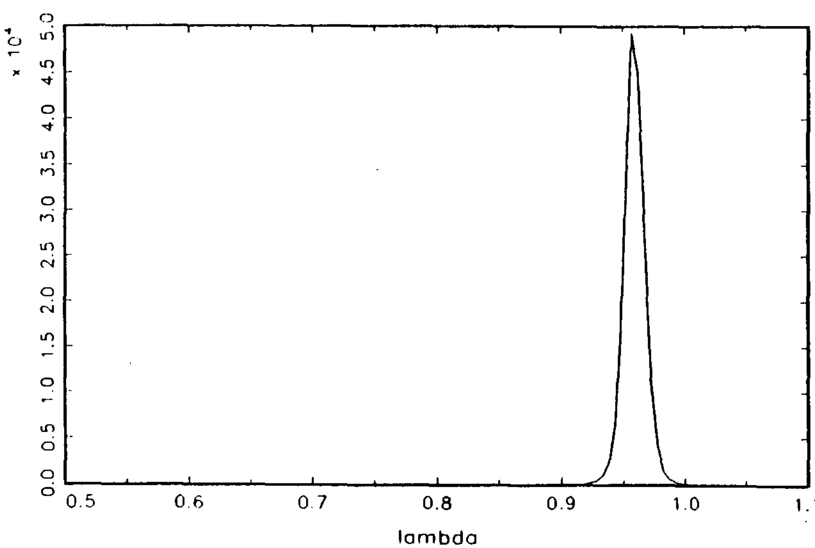
GERMANY, 1950 TO 1990.



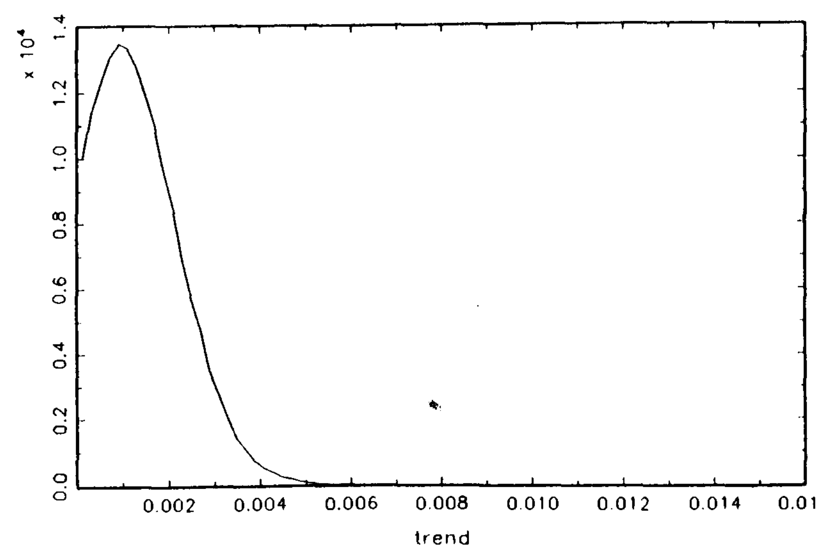
UNIT ROOTS IN GERMAN CONSUMPTION
Dens. for pairs of (t,lambda)



UNIT ROOTS IN GERMAN CONSUMPTION
Posterior density for lambda



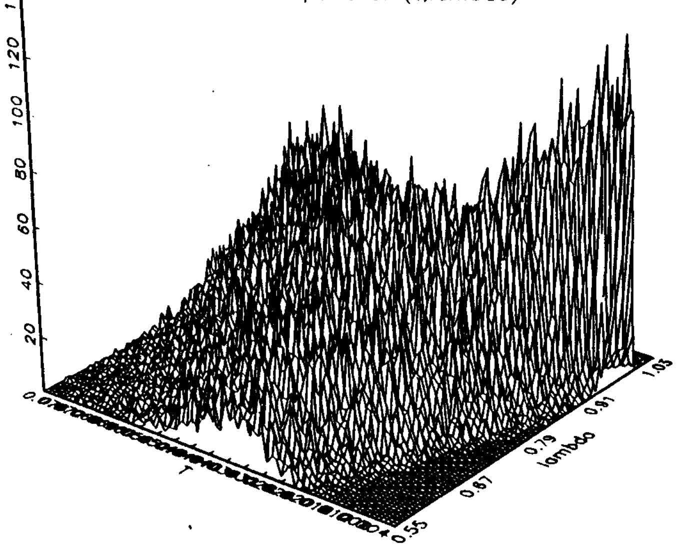
UNIT ROOTS IN GERMAN CONSUMPTION
TREND = 0.



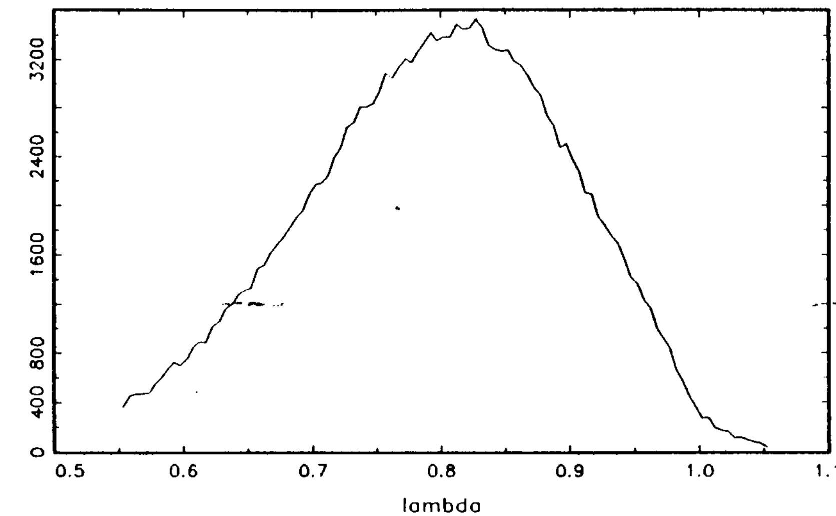
ESTIMATED TREND IN GERMAN CONSUMPTION
Post. dens. for trend coef.

NORWAY, 1946 TO 1991.

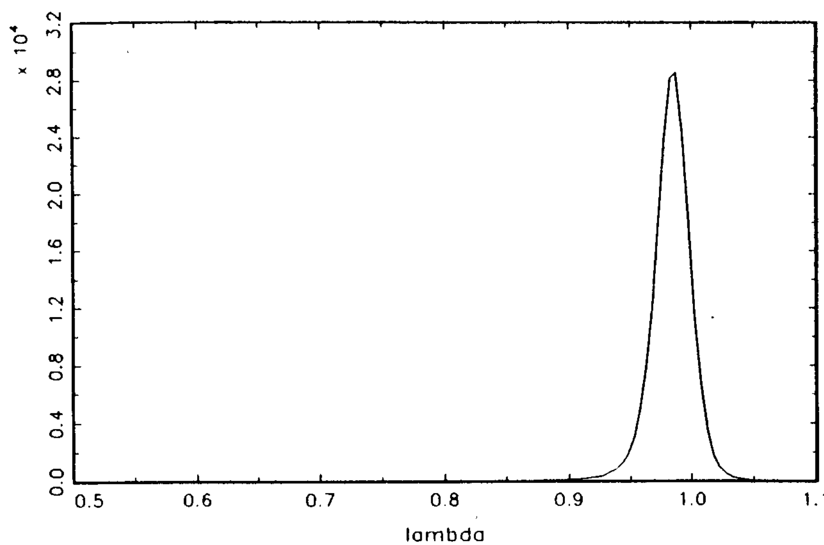
UNIT ROOTS IN NORWEGIAN CONSUMPTION
Dens. for pairs of (t,lambda)



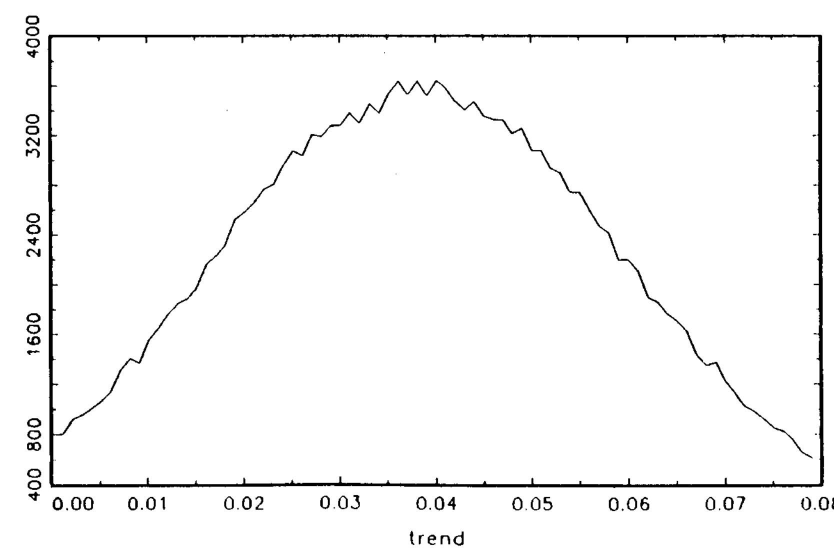
UNIT ROOTS IN NORWEGIAN CONSUMPTION
Posterior density for lambda



UNIT ROOTS IN NORWEGIAN CONSUMPTION
TREND = 0.

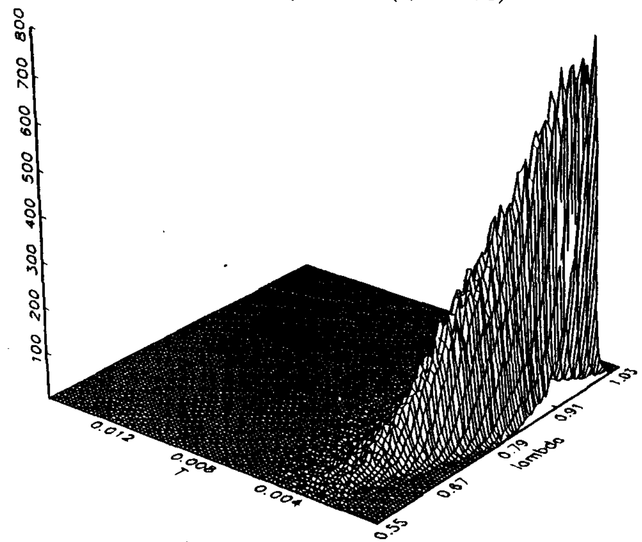


ESTIMATED TREND IN NORWEGIAN CONSUMPTION
Post. dens. for trend coef.

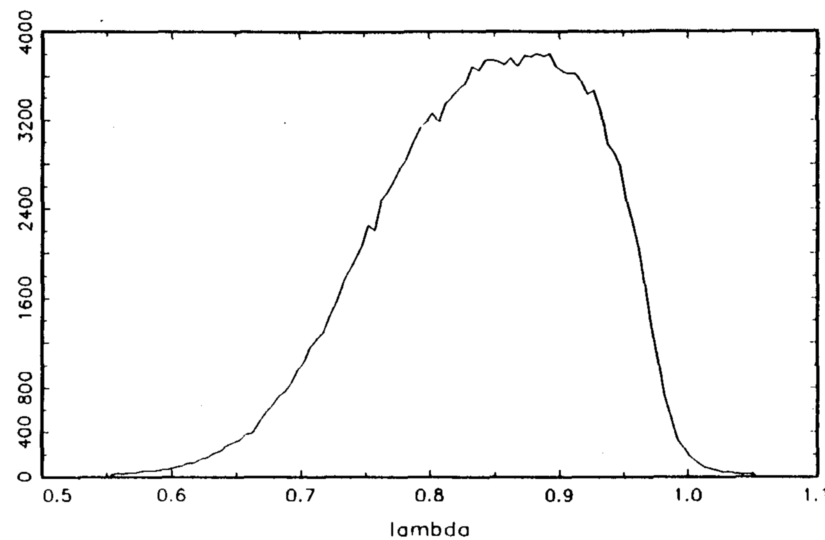


SWEDEN, 1950 TO 1990.

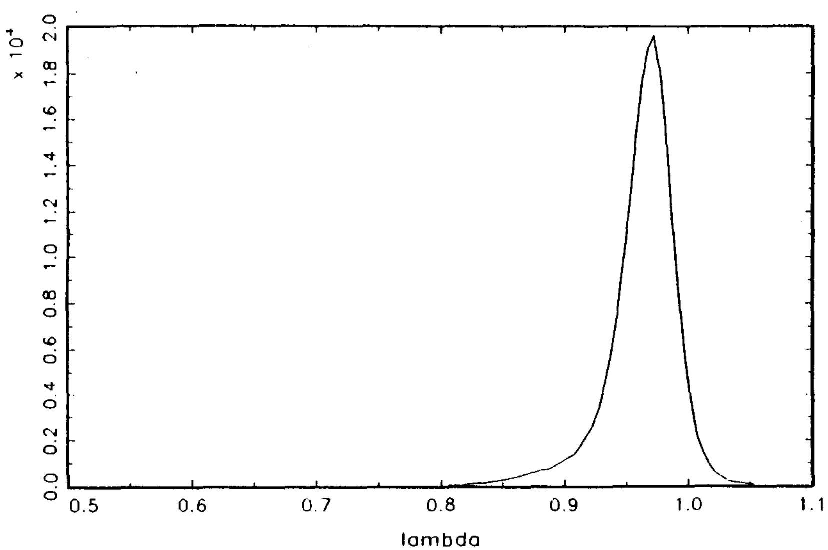
UNIT ROOTS IN SWEDISH CONSUMPTION
Dens. for pairs of (t,lambda)



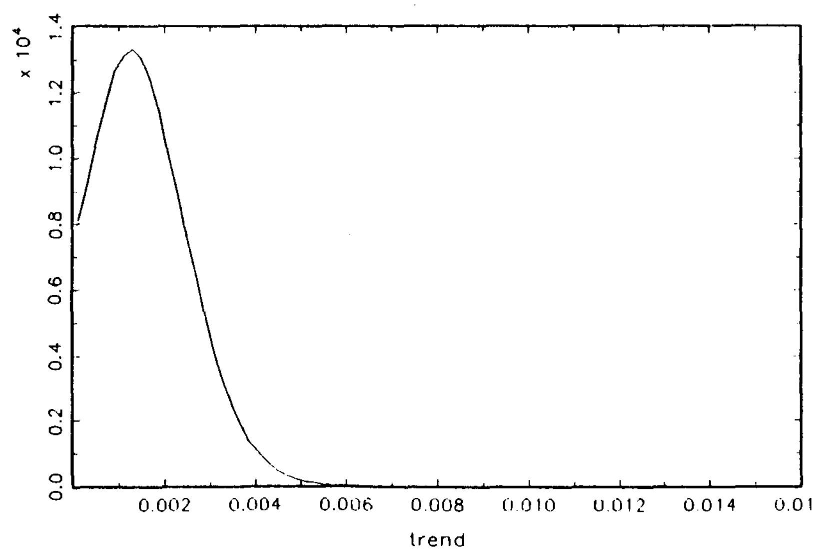
UNIT ROOTS IN SWEDISH CONSUMPTION
Posterior density for lambda



UNIT ROOTS IN SWEDISH CONSUMPTION
TREND = 0.

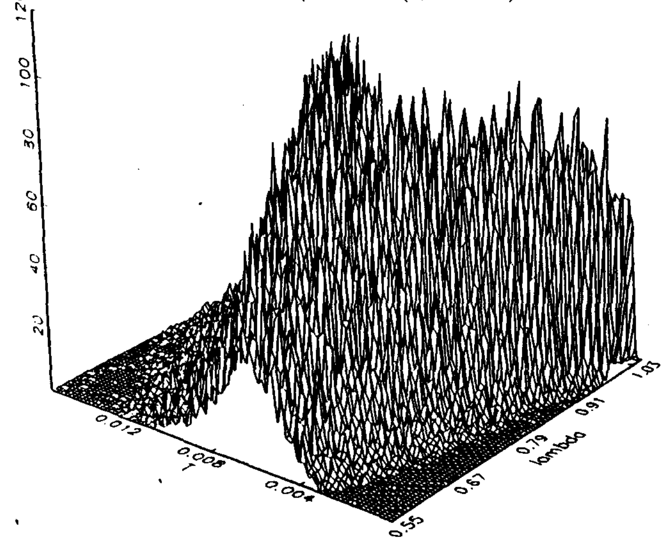


ESTIMATED TREND IN SWEDISH CONSUMPTION
Post. dens. for trend coef.

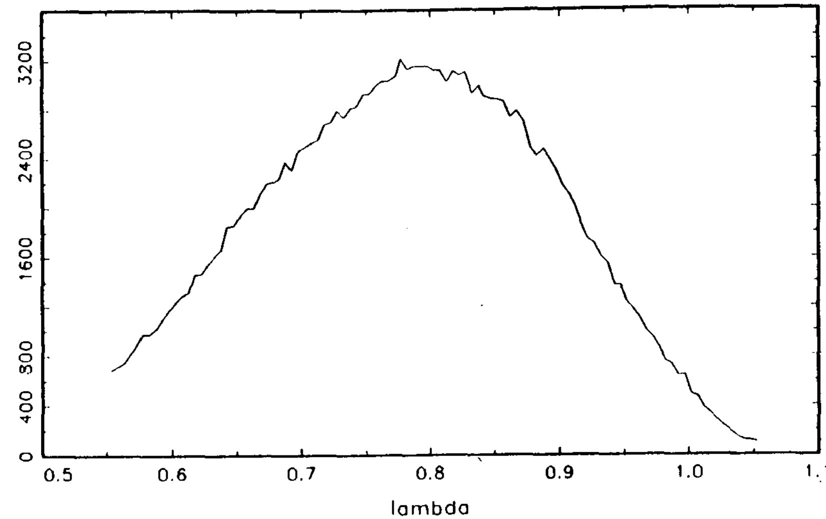


UNITED KINGDOM, 1948 TO 1990.

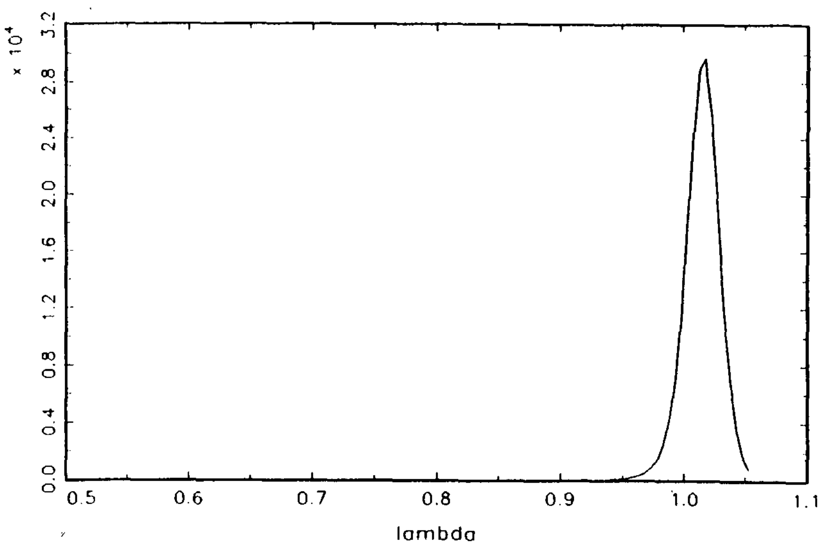
UNIT ROOTS IN UK CONSUMPTION
Dens. for pairs of (t,lambda)



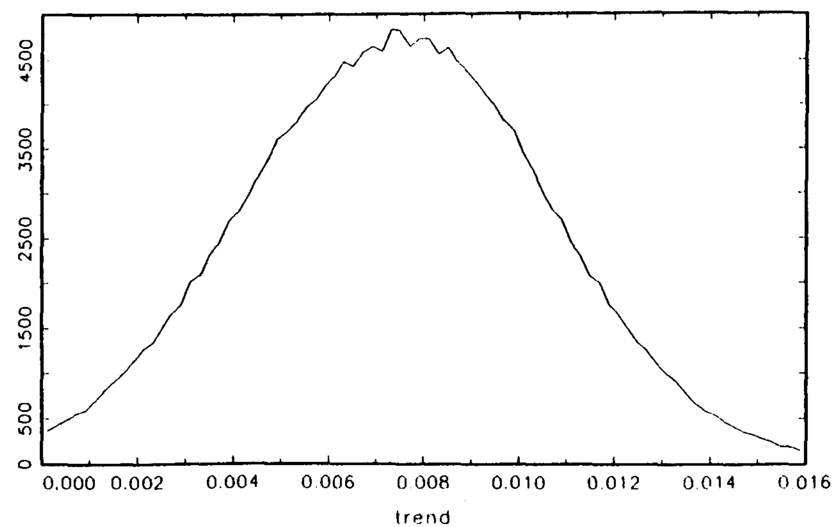
UNIT ROOTS IN UK CONSUMPTION
Posterior density for lambda



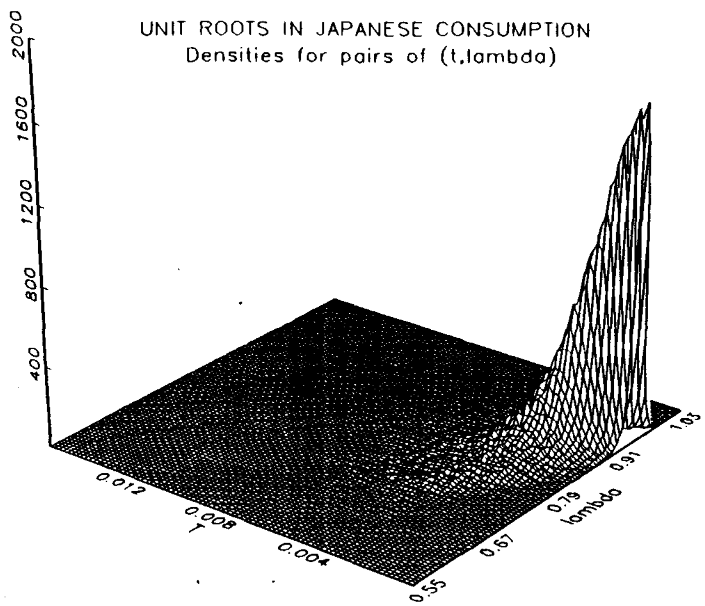
UNIT ROOTS IN UK CONSUMPTION
TREND = 0.

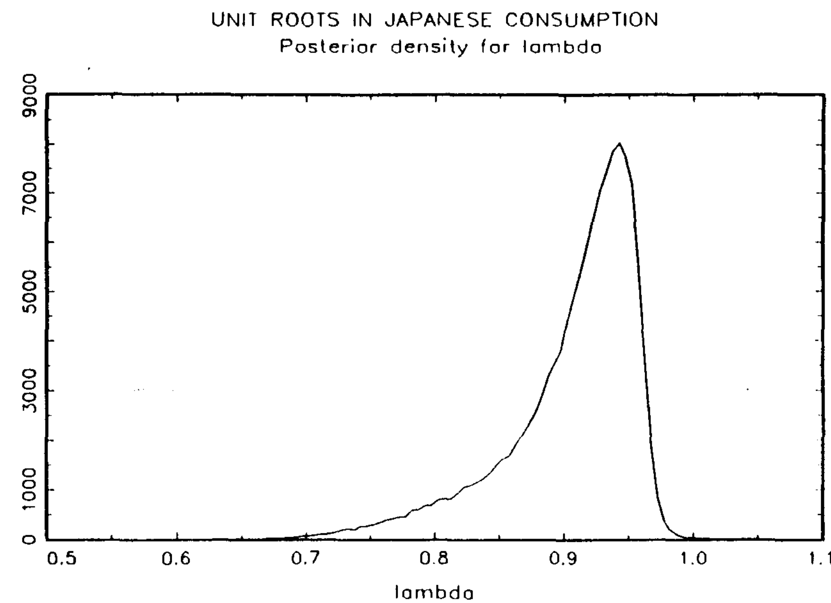


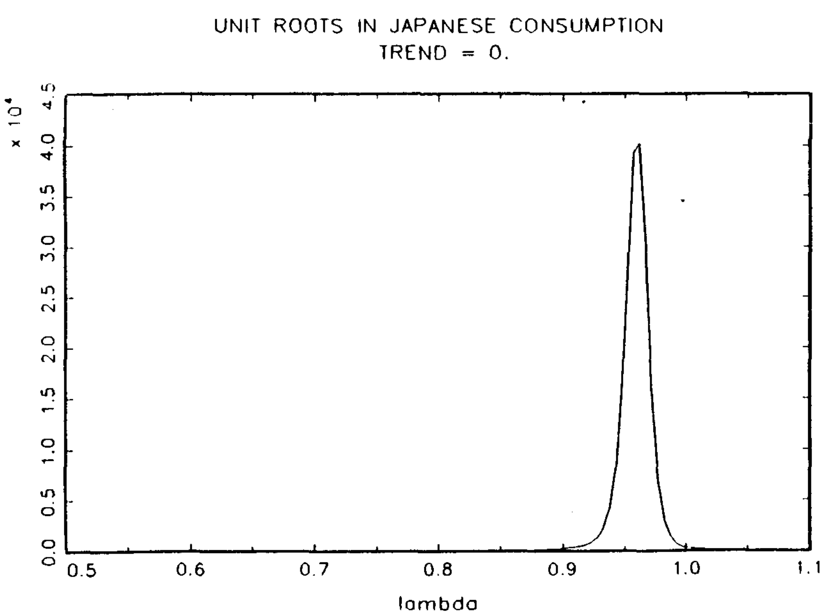
ESTIMATED TREND IN UK CONSUMPTION
Post. dens. for trend coef.

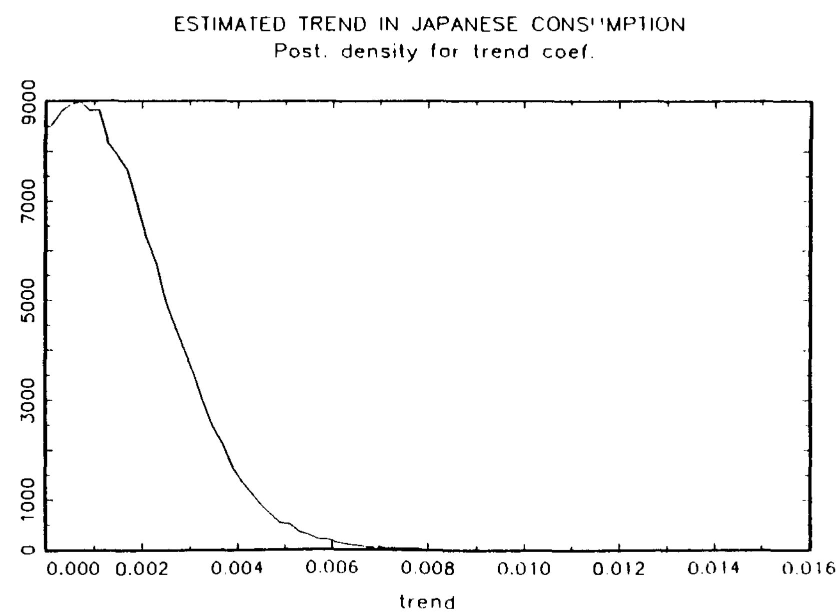


JAPAN, 1955 TO 1990.



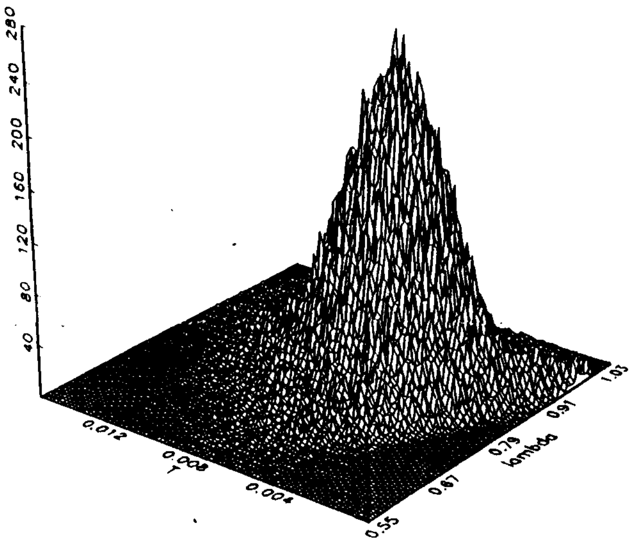




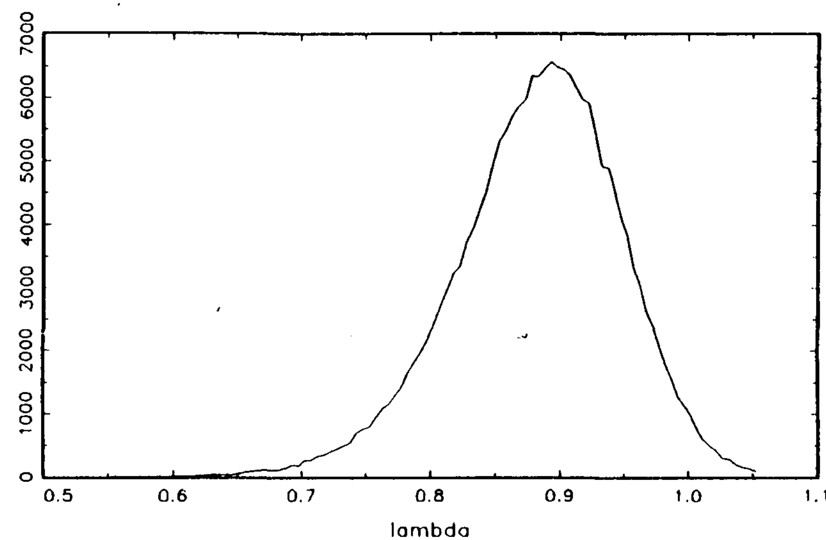


CANADA, 1946 TO 1990.

UNIT ROOTS IN CANADIAN CONSUMPTION
Dens. for pairs of (t,lambda)



UNIT ROOTS IN CANADIAN CONSUMPTION
Posterior density for lambda



UNIT ROOTS IN CANADIAN CONSUMPTION
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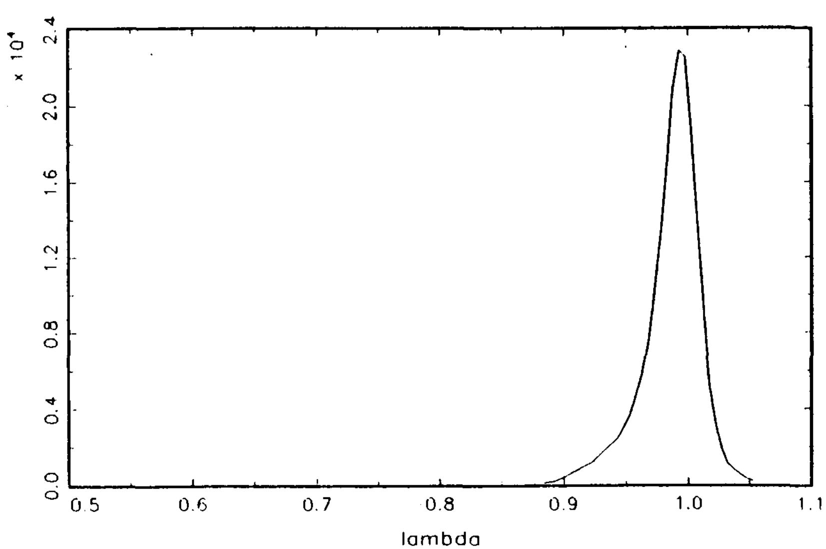
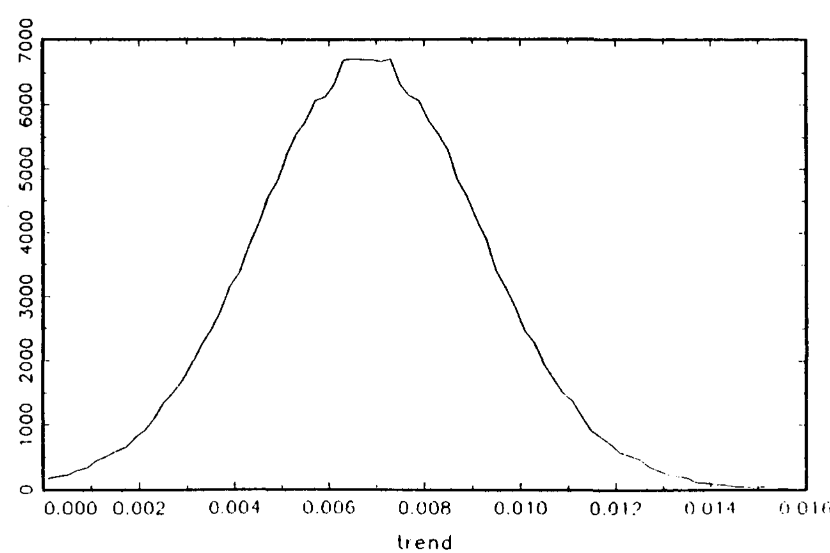
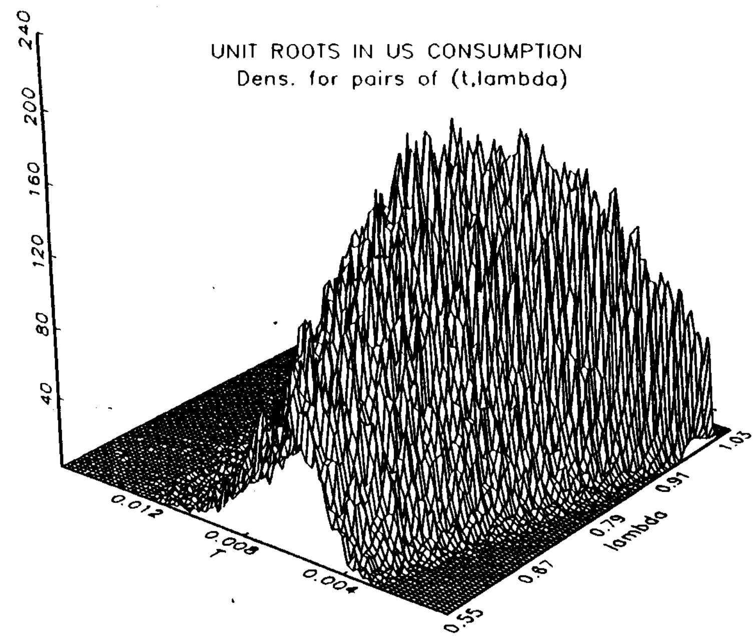


ESTIMATED TREND IN CANADIAN CONSUMPTION
Post. dens. for trend coef.

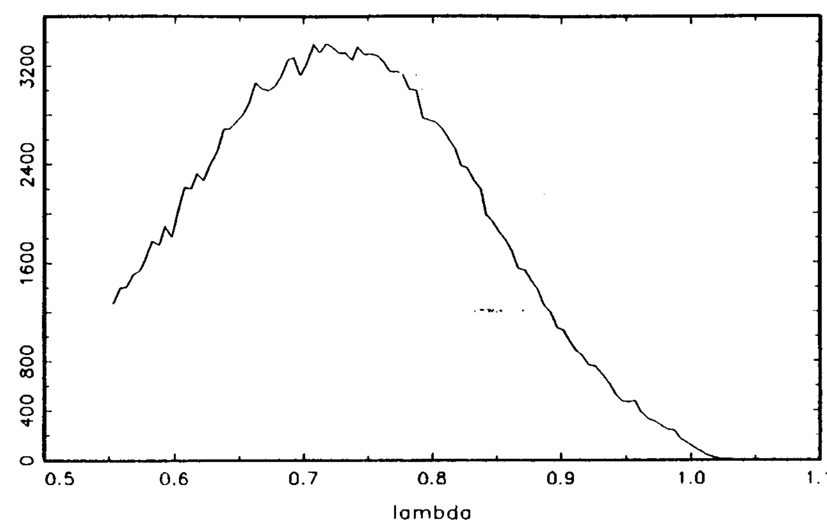


USA, 1946 TO 1990.

UNIT ROOTS IN US CONSUMPTION
Dens. for pairs of (t,lambda)



UNIT ROOTS IN US CONSUMPTION
Posterior density for lambda



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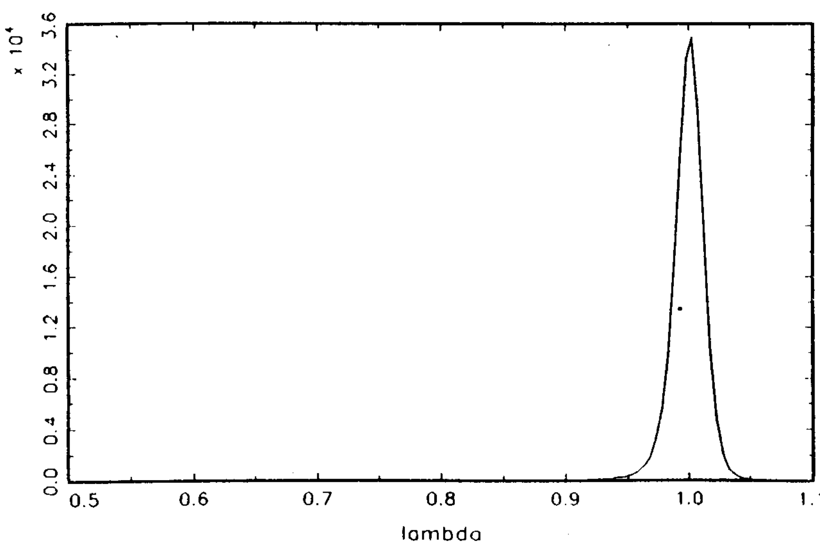
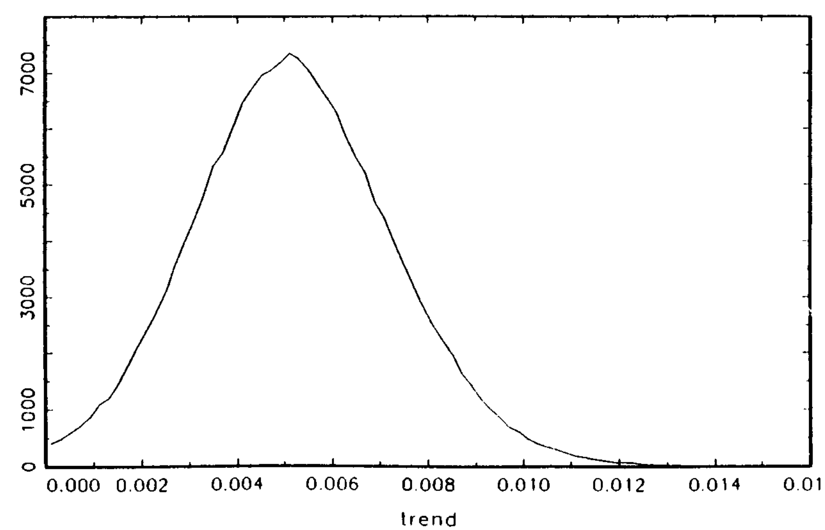
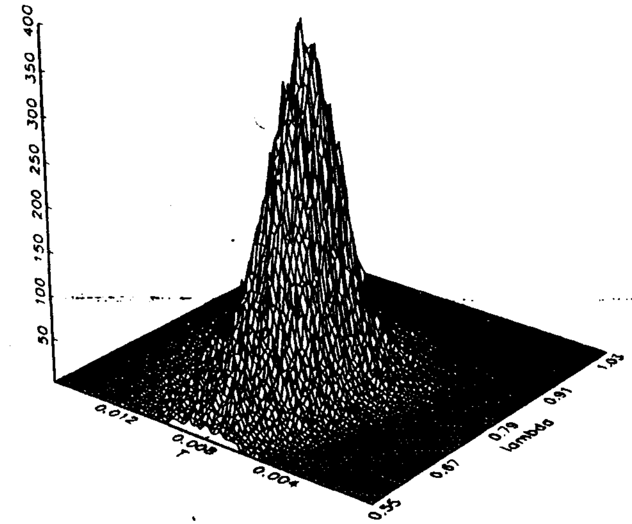


ESTIMATED TREND IN US CONSUMPTION
Post. dens. for trend coef.

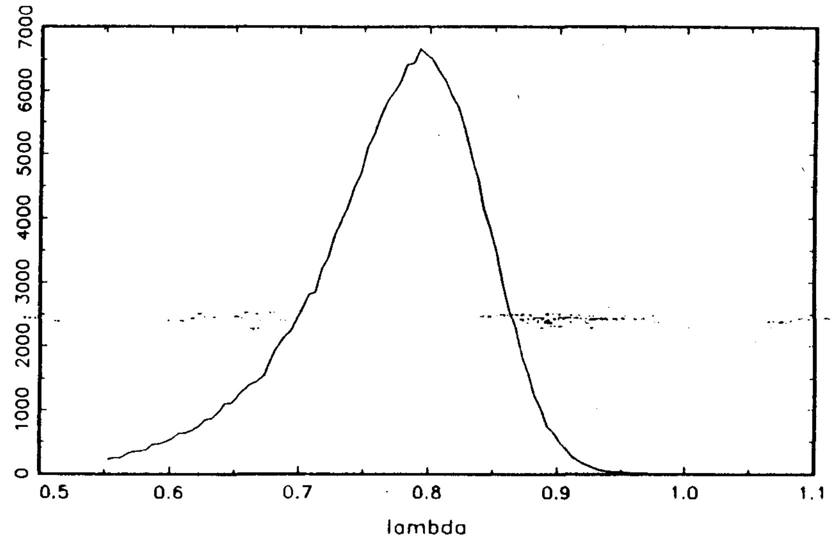


CANADA, 1926 TO 1990.

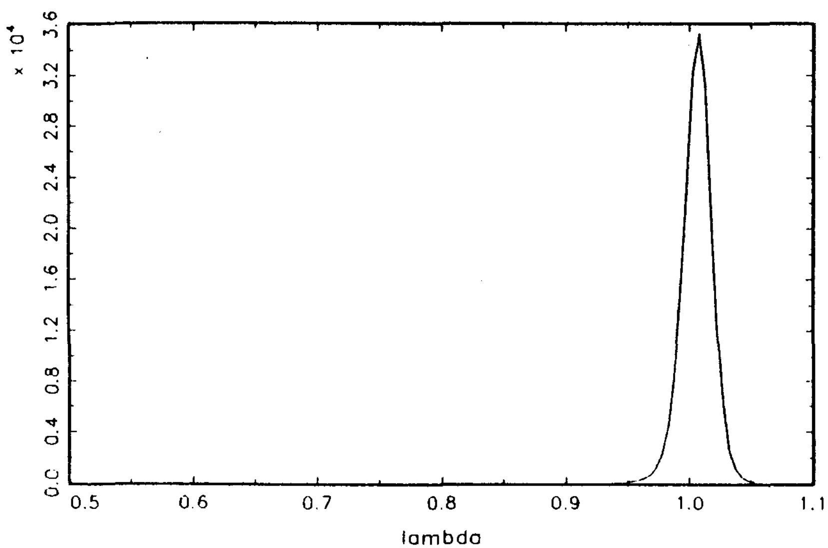
UNIT ROOTS IN CANADIAN CONSUMPTION
Dens. for pairs of (t,lambda)



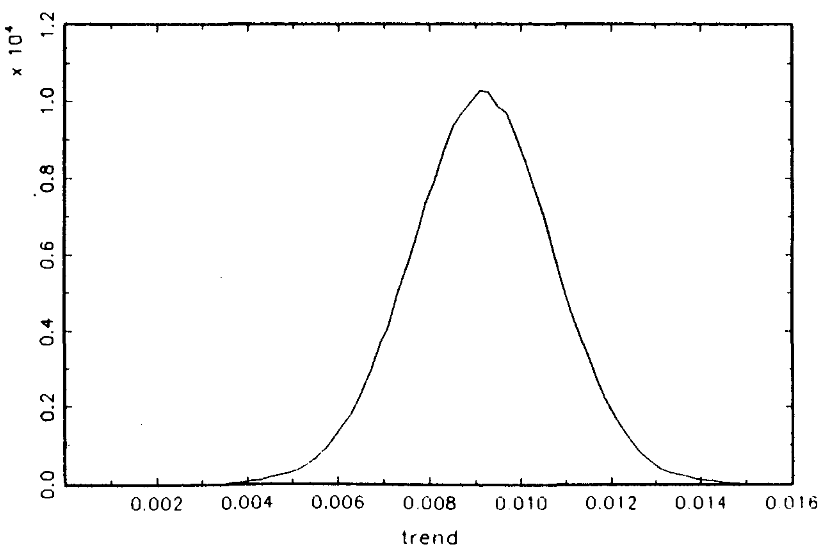
UNIT ROOTS IN CANADIAN CONSUMPTION
Posterior density for lambda



UNIT ROOTS IN CANADIAN CONSUMPTION
TREND = 0.

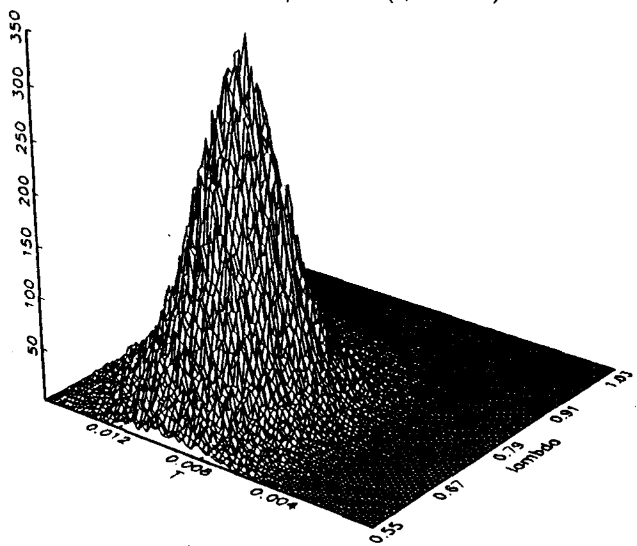


ESTIMATED TREND IN CANADIAN CONSUMPTION
Post. dens. for trend coef.

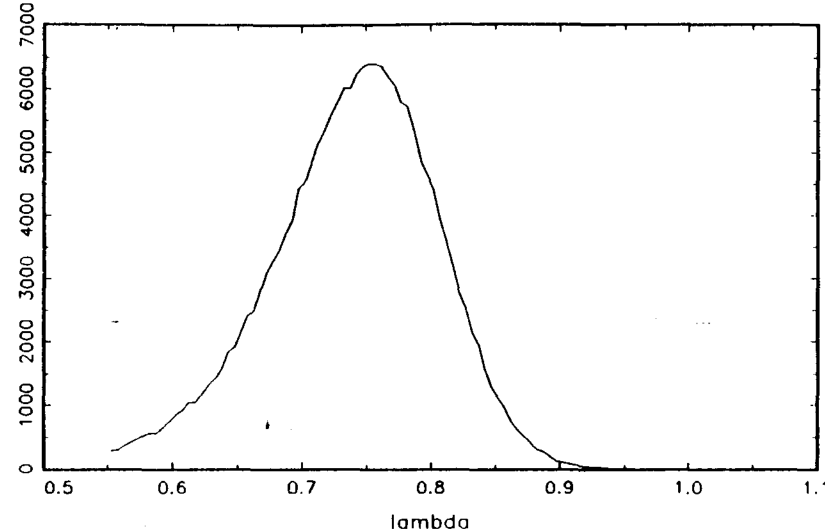


USA, 1929 TO 1990.

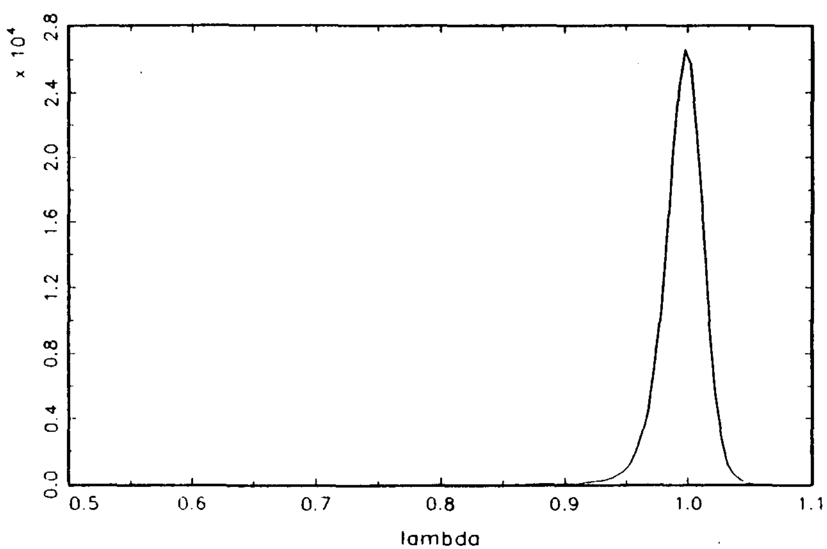
UNIT ROOTS IN US CONSUMPTION
Dens. for pairs of (t,lambda)



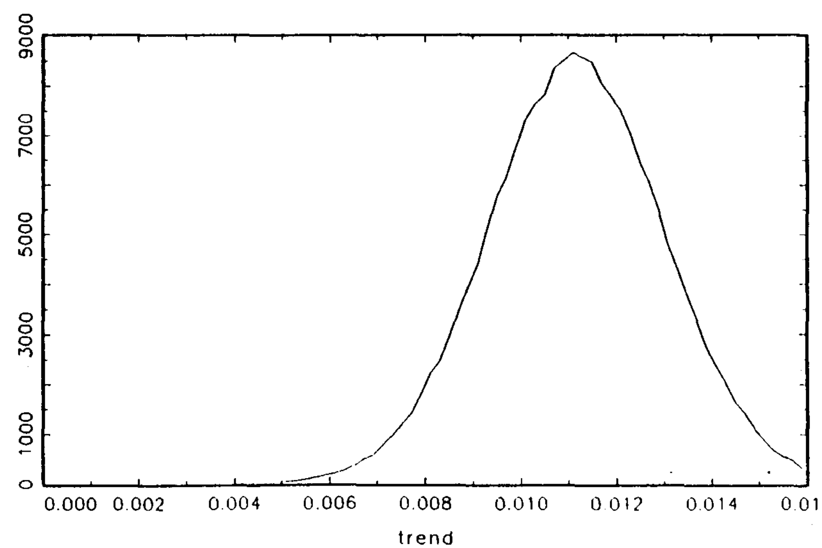
UNIT ROOTS IN US CONSUMPTION
Posterior density far lambda



UNIT ROOTS IN US CONSUMPTION
TREND = 0.

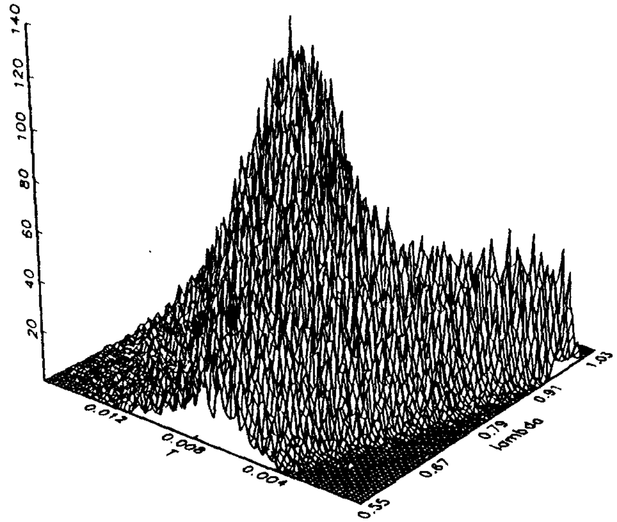


ESTIMATED TREND IN US CONSUMPTION
Post. dens. for trend coef.

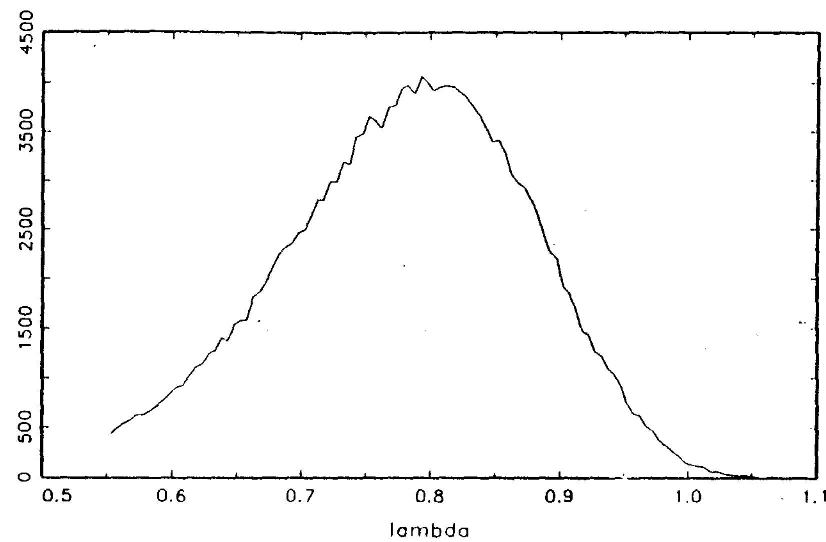


NORWAY, 1926 TO 1991.

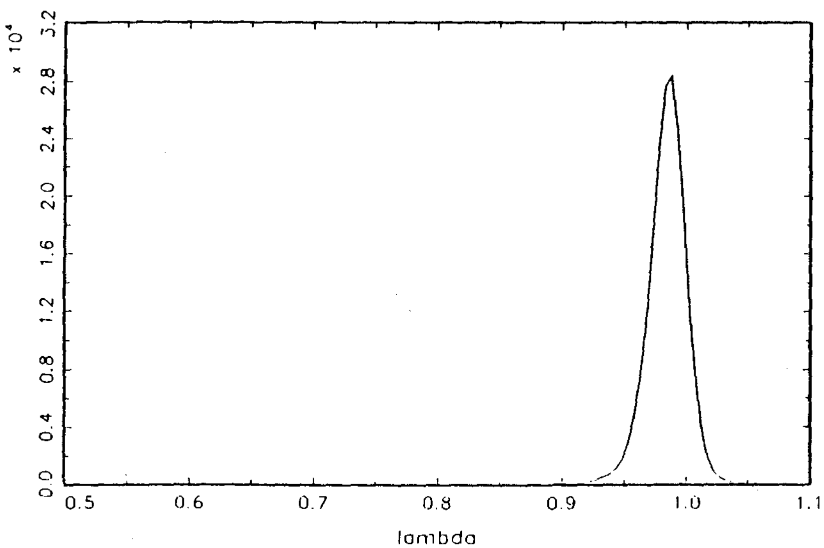
UNIT ROOTS IN NORWEGIAN CONSUMPTION
Dens. for pairs of (t,lambda)



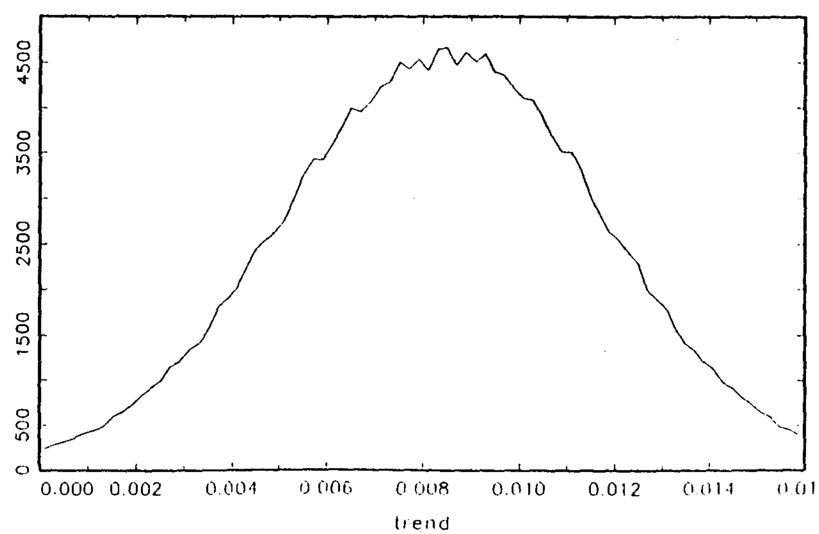
UNIT ROOTS IN NORWEGIAN CONSUMPTION
Posterior density for lambda



UNIT ROOTS IN NORWEGIAN CONSUMPTION
TREND = 0.

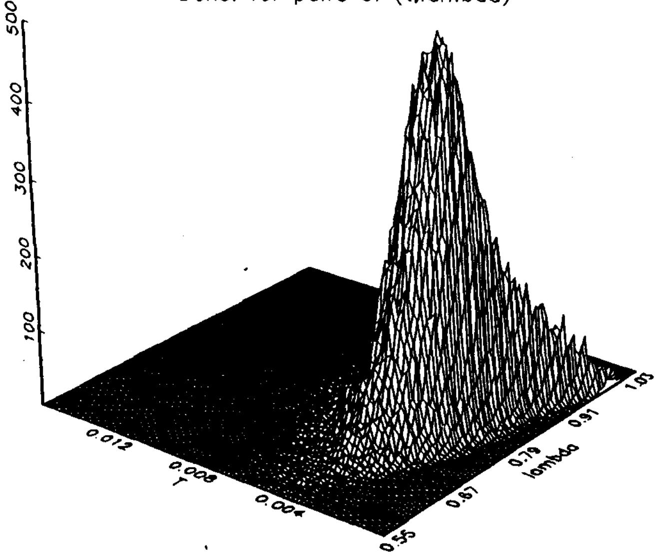


ESTIMATED TREND IN NORWEGIAN CONSUMPTION
Post. dens. for trend coef.

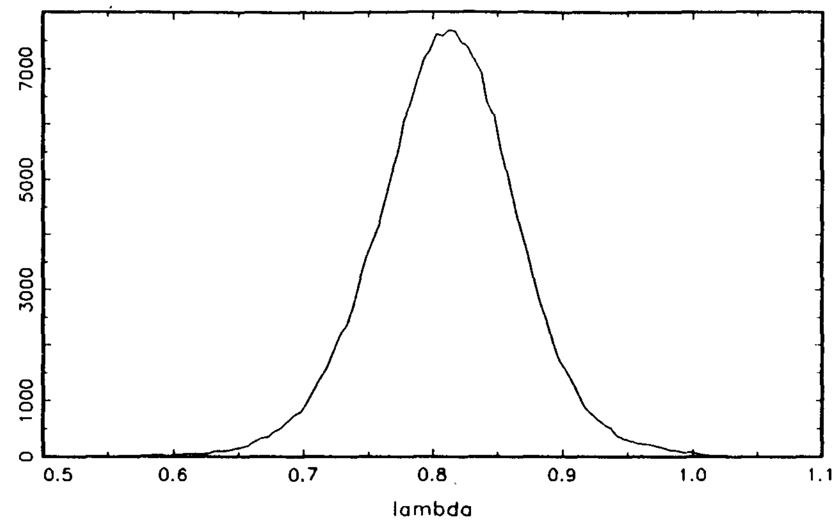


NORWAY, 1865 TO 1991.

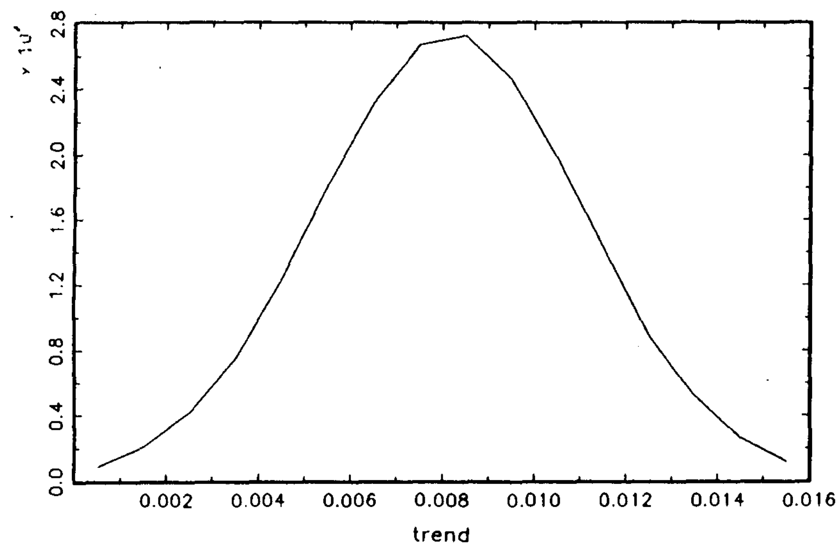
UNIT ROOTS IN NORWEGIAN CONSUMPTION
Dens. for pairs of (t,lambda)



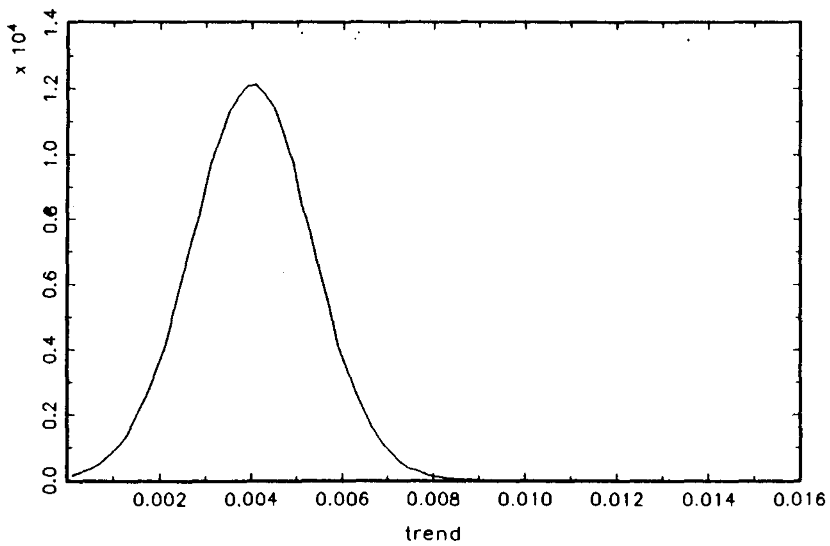
UNIT ROOTS IN NORWEGIAN CONSUMPTION
Posterior density for lambda



EST. POSTWAR TREND IN NORWEGIAN CONS.
Post. dens. for trend coeffs.

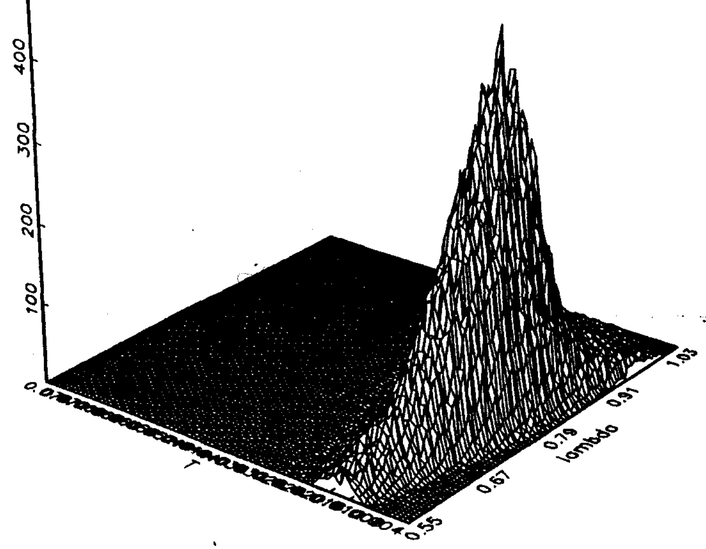


EST. PREWAR TREND IN NORWEGIAN CONS.
Post. dens. for trend coef.

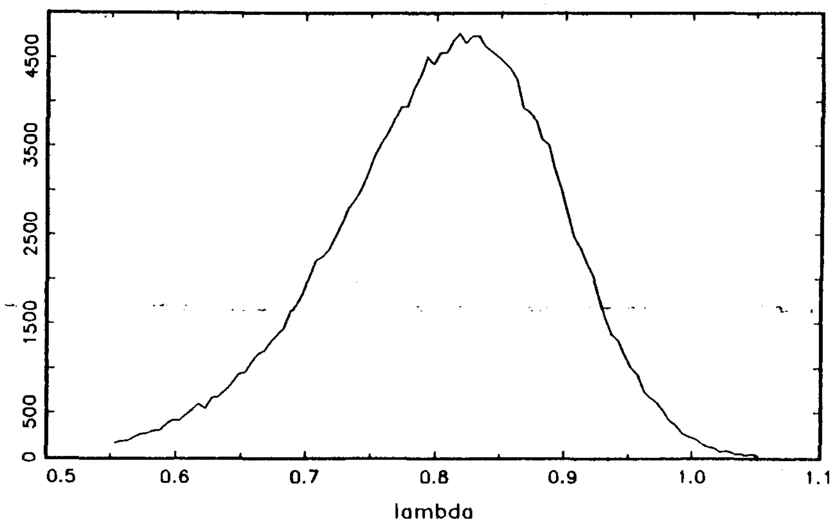


NORWAY, 1865 TO 1914.

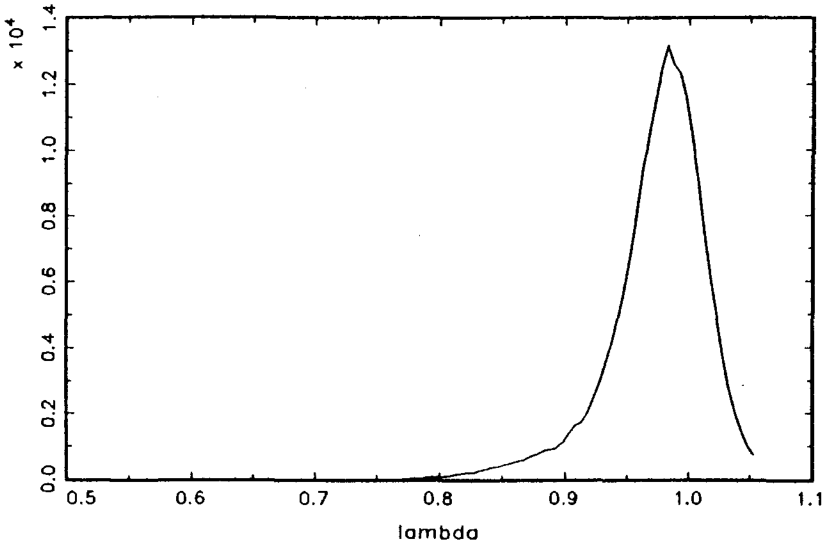
UNIT ROOTS IN NORWEGIAN CONSUMPTION
Dens. for pairs of (t,lambda)



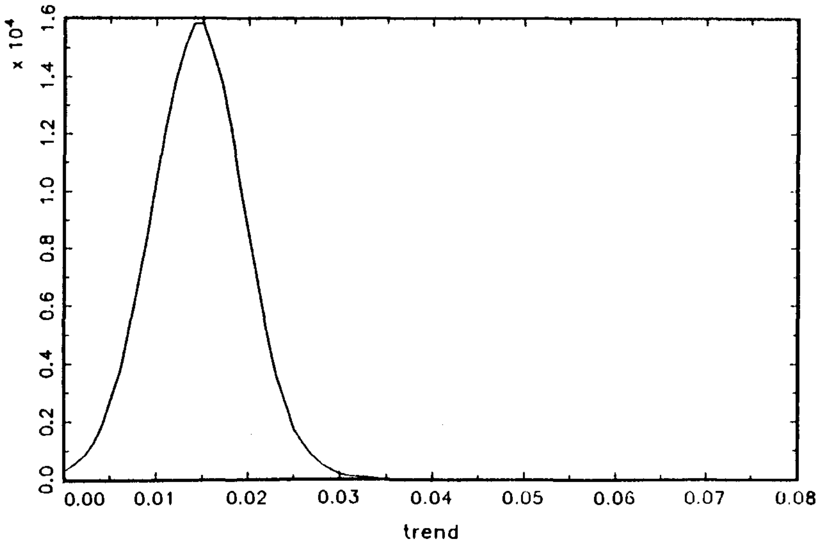
UNIT ROOTS IN NORWEGIAN CONSUMPTION
Posterior density for lambda



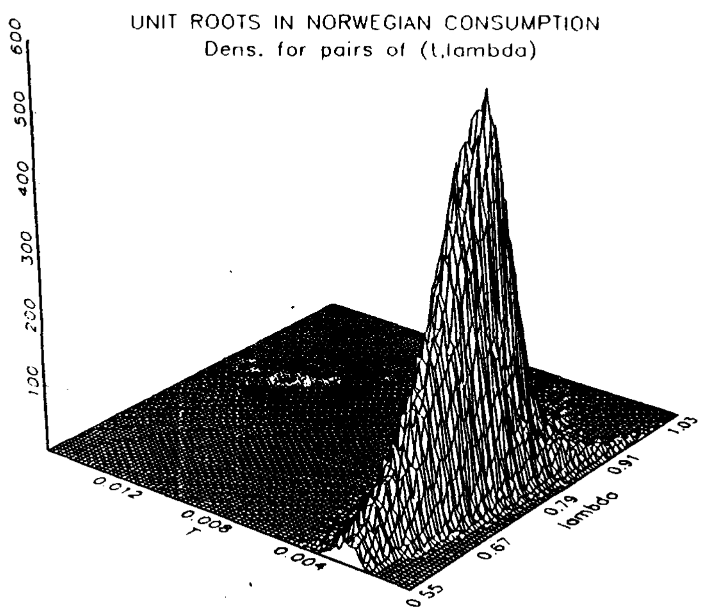
UNIT ROOTS IN NORWEGIAN CONSUMPTION
TREND = 0.

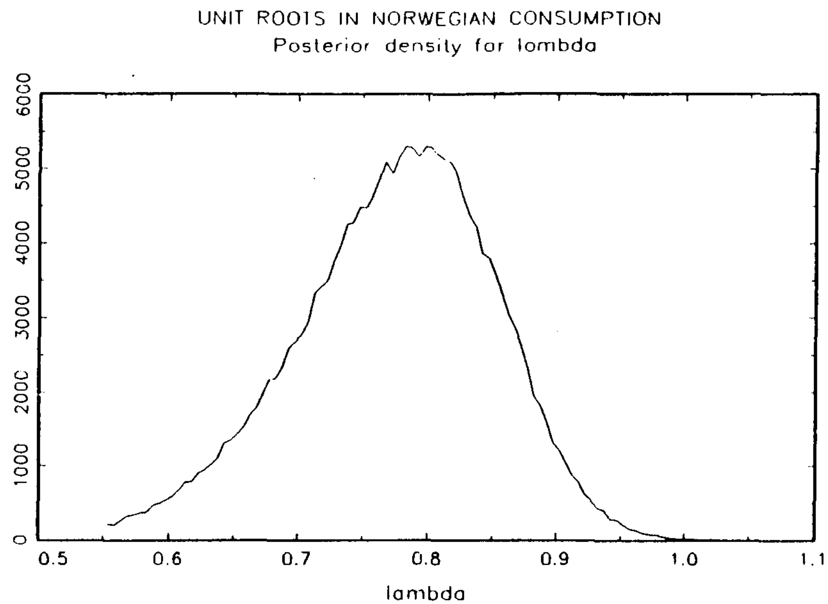


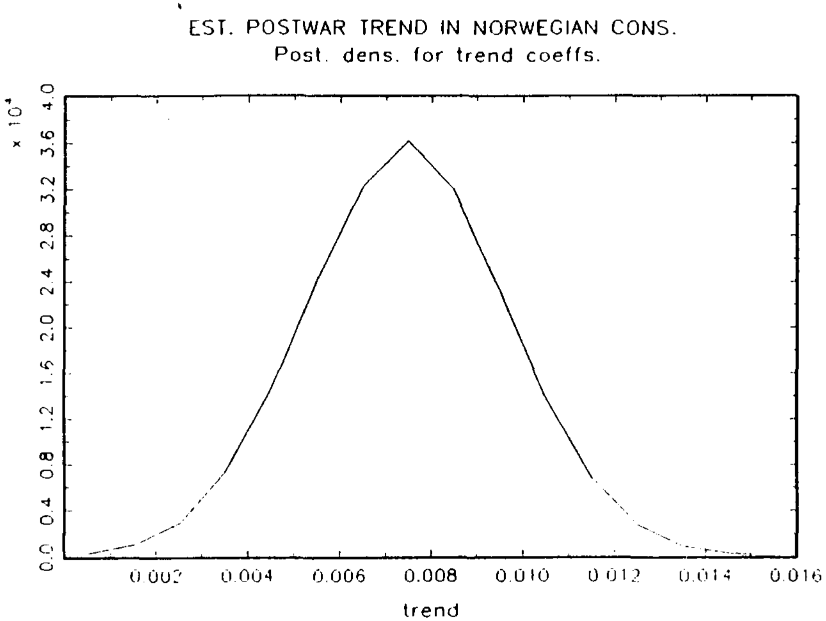
ESTIMATED TREND IN NORWEGIAN CONSUMPTION
Post. dens. for trend coef.

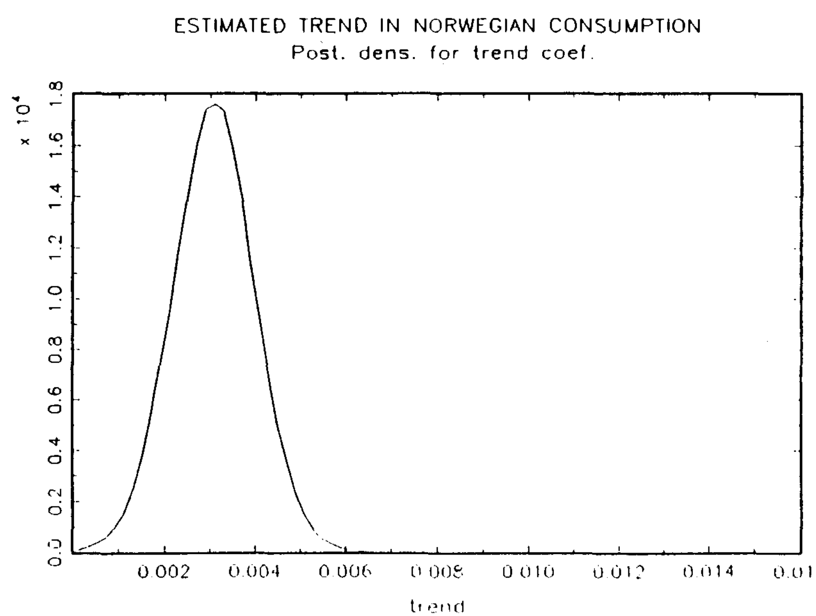


NORWAY, 1865 TO 1914 AND 1946 TO 1991.



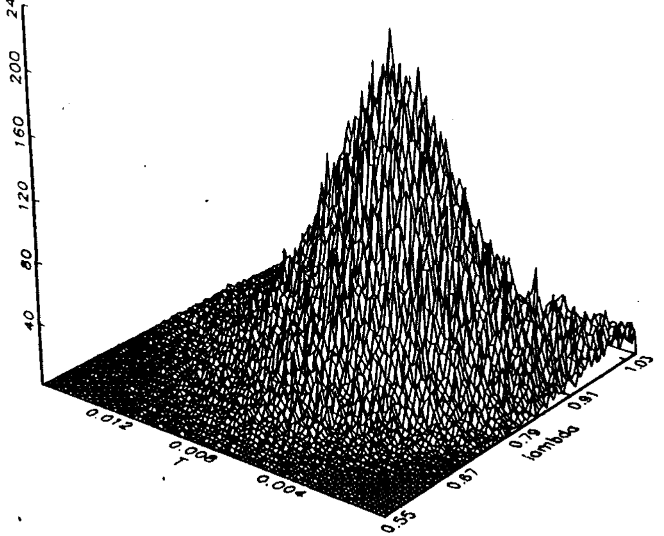




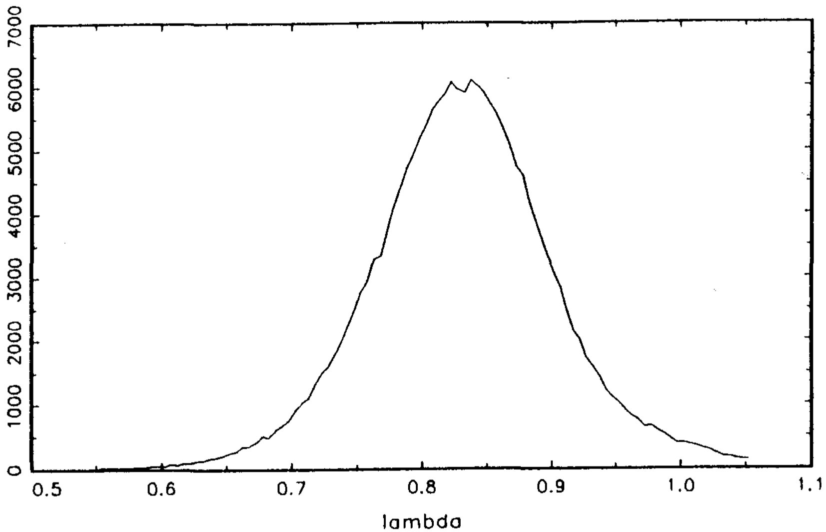


NORWAY, 1915 TO 1991.

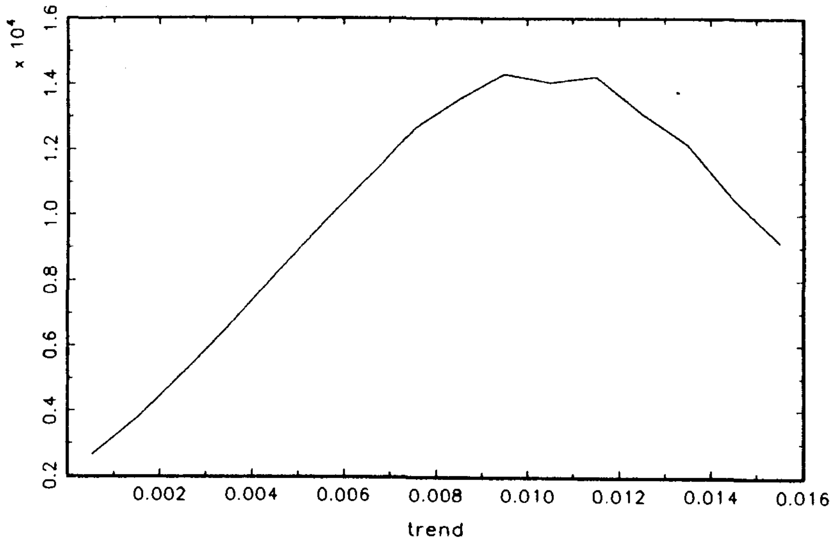
UNIT ROOTS IN NORWEGIAN CONSUMPTION
Dens. for pairs of (t,lambda)



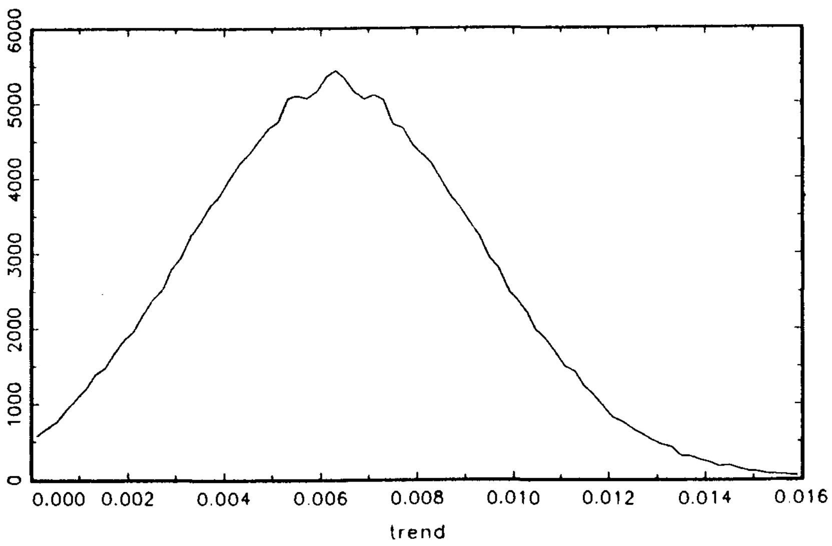
UNIT ROOTS IN NORWEGIAN CONSUMPTION
Posterior density for lambda



EST. POSTWAR TREND IN NORWEGIAN CONS.
Post. dens. for trend coeffs.



EST. PREWAR TREND IN NORWEGIAN CONS.
Post. dens. for trend coef.



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ESSAY II

CHANGES IN HOUSE PRICES AND THE PROPENSITY TO CONSUME.

A theoretical analysis and empirical study on a set of Norwegian microeconomic data.

1.0 INTRODUCTION[1]

The relationship between consumption and income has been the main topic of numerous studies. How consumption relates to the other components of household wealth has received much less attention. The purpose of this study is to test the theory based on the relationship between household consumption and non-income components of household wealth. The single most important component of this wealth is residential capital.

The theoretical link from household wealth to consumption is rather well understood and in principle there is no difference between human capital and the remaining household wealth. Normally, we assume that households have preferences for their own consumption and sometimes for the consumption of their descendants. As long as we disregard uncertainty about future prices and income streams, it should not really matter whether the source of their consumption possibilities is labour income, real estate wealth or bank accounts.

[1] I would like to thank professor Knut A. Mork for valuable comments on earlier drafts of this paper.

Based on these principles, we present a model that links the consumption of non-housing goods to house prices. It is crucial whether the household owns its dwelling or is renting it. The house-owners receive a capital gain from price increases. We would expect them to increase their consumption of other goods in response to an increase in house prices. If demand for housing is inelastic, tenants are likely to reduce their overall non-housing consumption. To them the price increase represents higher future housing costs. Although we solve our model for quadratic utility, the difference between the two types of households is related to the budget constraint. Thus, the difference should be considered a far more general property of rational consumer behaviour.

We investigate this model on a Norwegian panel data set. The data cover a period with large fluctuations in house prices and as such should be ideal for this purpose. Our results indicate that house-owners have a large propensity to consume out of price changes which is consistent with a rather short horizon. However, tenants seem to have an equally large positive estimate which questions the interpretation of this as a wealth effect. Aggregate effects that influence our results seem perhaps like a more plausible explanation. We suspect that there are common causes for the boom in house prices and consumption. In our estimates, we have t-values of about 1.5 and sometimes less [2]. Consequently, we have trouble making firm statements at a 5%-level. The results seem insensitive to outliers and influential observations.

1.1 The composition of household wealth in Norway

It is hard to obtain comprehensive data on household wealth in Norway. The available data are based on values set for taxation. These values do mostly not reflect the real value of the different assets. For instance, the taxation values of houses are set arbitrarily at about 10% to 40% of market value. But also for other assets like non-public company shares, the values do not reflect the market value.

[2] The errors are based on White correction for general heteroscedasticity.

We have to apply other sources in order to assess the value of the residential capital for an average household. In the statistics we have data on the number of dwellings of different kinds like detached, semidetached, undetached and apartments. We can acquire some information about the market prices by assuming that the prices achieved through sales by real estate agents reflect the market prices of the average house[3]. When we combine these two sources, we find a total value of the national residential capital stock of approximately NOK 1200 billion[4] in 1991. This is almost twice the value of GNP and constitutes about NOK 700' (' denotes thousands) per household.

Not all of this capital constitutes household wealth. Some of the residential capital is owned by the government and local authorities. Mostly, these are inexpensive apartments that are alotted to people as part of the social policy. Firms own a minor share of the residential capital while the remaining and the greater part is privately owned by households directly or through housing co-operations[5]. In the data set that we shall use, just over 89% of the households own their own residence. If we apply this share on the total we get an average residential wealth for households of approximately NOK 620'.

Official statistics for the year 1991 estimate household gross financial capital to be NOK 177' on average of which bank deposits sum to NOK 126'. More risky assets like bonds and shares make up only NOK 25'[6]. In the data set we shall apply, only 31% of the households are shareholders. For the average household price movements in the stock market are of much less importance than similar effects due to price fluctuations in the market for residential housing.

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[3] Most likely, the average house will be somewhat cheaper for two reasons. First, the propensity to move is higher in more highly populated areas where prices are somewhat higher. The high price markets will be over-represented in the data based on averages. Second, there will probably be a tendency that the cheaper the house the more likely it is that sales are completed without the use of a real estate agent.
[4] approximately NOK 7 = USD 1.
[5] A considerable share of apartments and some of the undetached houses have been organised as companies. For the purpose of this paper, it seems reasonable to treat these as house owners.
[6] Even though these numbers are based on tax values, it is obvious that shares constitute a much less important component of household wealth. In several countries it is common that households make private pension savings. These savings often depend strongly on the stock market. In Norway private pension savings are small in comparison.

Still, the major source of household wealth is the expected future labour income. In 1991, average household salaries amounted to NOK 208', while income for self-employed was about NOK 27.5'. Tax liabilities were NOK 65'. This results in a level of disposable income of NOK 169.5'. These numbers are averages for all households including students and retired persons. Let us assume that a household has a life expectancy of 50 years and disregard uneven distribution of labour income over the life cycle. Then, with a real growth in income of 1% and a discount factor of 4%, a representative household with 25 years left of expected life would have a human capital of about NOK 3 000' or about 5 times the level of the residential capital. Accordingly, human capital followed by the residential capital stock should be the far most important sources for household consumption. Other factors should be of much less importance.

1.2 The data set

The consumption data are collected by the Statistics Norway in their annual consumer survey for Norwegian households. In this survey, there is a sub-sample of about 200 households which reappears two years in a row. This constitutes a set of panel data. Although the time dimension is short, it enables us to study the change in consumption from one year to the next.

Our data contain three such sets and cover the years from 1986 - 1989 (that is 1986-1987, 1987-1988 and 1988-1989). This gives us observations for about 600 households. The years covered were a period of very erratic house prices and should as such suit our purpose well. From 1985 to 1987, real house prices rose by a yearly average of about 6% while from 1988 to 1989 relative prices fell by almost 12 %. Prices peaked in the beginning of 1987.

The participating households keep a full and detailed record of all their expenses for a period of two weeks during the year. The accounting period is random and evenly distributed across the year among the participating households. However, the household is surveyed at the same two-week period in both years. The expenses are multiplied by 26 in order to «annualize» the amounts.

At the end of the year an interview is done in order to supply further information. The households are asked how much they have spent on important groups of durables like vehicles, boats, furniture, household machines, electronic equipment etc. during the year. So we get a good picture of the pattern of expenditure for the household.

All households that have moved house between the two years have been removed from the survey by the Bureau. Thus, it is not possible to study the decision to move. When we investigated the data prior to estimation, we found 11 additional households that apparently had changed their residence[7] between the first and the second survey period. In order to avoid systematic biases in the data, these were removed.

The survey includes important information about the dwelling. We know whether the household lives in an apartment, semidetached or detached house. Further, we have information about the size of the house and geographical region in which the household lives. We know whether it is a rural or urban area and especially we know whether they live in one of the three largest cities (Oslo, Bergen and Trondheim). An important piece of information is whether the household owns its own house or is renting it. In the analysis we shall make extensive use of this information.

However, we do not have specific information about the price of the house or the price that the household expects it to have. In order to estimate the wealth gain/loss that the household has had, we need to know the change in value of the house from one year to the next. We combine price information obtained from the statistics of the Norwegian Association of Real Estate Agents with the information we have about each house. As a result, we obtain a price index for each of four regions and three different types of

[7] The selection criteria were discrepancies in building year, type and size of the resident between the two years. The reason why the Bureau has not discovered these problems is because households may have

houses[8]. The price index is based on the floor space of the house and is computed on a quarterly basis[9].

Income data are collected from official taxable values. We have data for the two years that the household participates and for the two previous years. Consequently, we have no information about income not stated for taxation. More important, the data contain the income that defines pension rights. In particular, this does not include pension itself. This will cause a dramatic drop in our income measure for households where someone has retired during the year. We have tried to identify those households for which this might be the case[10]. For these households we have simply assumed that their income is proportional to their paid taxes.

1.3 Aggregate versus disaggregate data

There can be several reasons for the negligence of other sources of wealth in the income/ consumption relation. Most studies have used aggregate data and normally it is difficult to obtain reliable data for the development of different wealth variables over time. Furthermore, if we want to draw inference about household behaviour, there seems to be problems with how to interpret the causal relationships in the data. This is sometimes a problem in the income/consumption relation but seems to become more severe when we consider several measures of wealth. If prices of housing and

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changed residence without changing the postal address. In rural Norway, it is quite common that houses within a smaller district have the same postal address.

[8] The regions are 1) Oslo, 2) Bergen and Trondheim, 3) Akershus county and 4) the rest of Norway. This certainly is too aggregate to capture all the price differences between different regions but it does capture the most important differences. Based on the information in the survey data from the Bureau and the level of detail in the price statistics, this is the best we can do. The house types are flats, detached and semi-detached houses.

[9] In an earlier version of this paper, we computed a variable for gains and losses in the stock market. However, the variable was constructed in an indirect manner and apparently contained too much noise. All estimates based on this variable appeared highly insignificant.

[10] The selection has been based on age and sudden changes in income not reflected in the taxes paid by the household.

consumption both rise at the same time, this opens for several different explanations. Of course, similar problems arise when we consider price movements in the stock market.

We want to trace the behavioural link between house prices and household consumption. It is important to perform such a study on disaggregate data. The change in value that a household experiences on its dwelling is closely related to changes in the general market price level for housing and to future aggregate income prospects. Even though income changes at the household level will be correlated with movements in aggregate income, it seems plausible that this correlation is moderate and is dominated by individual changes. Furthermore, in disaggregate data we remove some of the time series effects by adding yearly dummy variables. Consequently, the use of disaggregated data should enable us to study effects from different wealth variables with less serious problems of simultaneity.

2.0 LITERATURE ON THE RELATION BETWEEN THE RESIDENTIAL MARKET AND THE MACRO ECONOMY

Most of the literature concerning the housing market deals with explanations for price movements in the market and not with their consequences. The main explanations pointed out by analysts are changes in the user cost, changes in construction costs, changes in aggregate (expected) income and demographical changes. For an overview of this literature, Poterba (1991) is recommended. His paper does contain a regression of changes in GNP explained by changes in house prices in the previous period; however, his interpretation is that the housing market discounts future changes in income.

Bover et al. (1989) discuss the relationship between the labour market and housing market. They focus on the lack of regional mobility in England as a consequence of regional differences in the income/house price ratio. Further, Muellbauer (1992) compares the housing markets in UK and Germany in order to study reasons of price movements especially based on differences in policy.

Muellbauer and Murphy (1990) provide one of the few studies on macroeconomic implications of the housing market. They tell a story of how different factors have contributed to a high growth rate for the house prices, mostly in the south-east parts of England towards the end of the eighties. They emphasise the role of the liberalisation of the credit markets some years before. In addition, they claim that households seem to have adaptive expectations. Growth in prices gives rise to expectations of further growth. This led to a «bubble» in the housing market. In their opinion, households increased their consumption partly due to a wealth effect from rising house prices and partly as a consequence of increased collateral. However, they lack firm testing of their hypothesis and as the discussants following their paper show, there are other plausible interpretations of the data that Muellbauer and Murphy present.

Koskela, Loikkanen and Virén (1992) have a twofold task. First, they consider variables that explain house prices over time. Second, they study the relationship between the savings rate and different explanatory variables, i. a. house prices[11]. They use data for the Finnish economy and find a negative correlation between house prices and the saving rate. Dependent on the set of other included variables they get t-values between 1.4 and 2.3 so the effect appears to be significant. Their unit of measurement is somewhat unclear so it is hard to assess how important house prices are in explaining the saving rate. More important, they apply aggregate data and as such their results are not clear if we want to understand household behaviour. Their results may arise from an endogenous house price variable. Changes in the saving rate may be an important explanatory factor of house prices. The saving rate is not included in the first house price equation.

In a series of papers, Brodin (1989) and Brodin and Nymoen (1989, 1991) have studied the macroeconomic consequences of changes in private housing wealth in Norway.

They have reestimated a Norwegian consumption function in order to explain the boom in consumption expenditure using a measure of household wealth. The variation in this wealth variable follows to a large extent the movements in house prices. They show that earlier versions of a consumption function excluding measures of wealth could not explain the consumption history after 1985. In stead of a genuine structural break, they argue that this is due to an omitted wealth variable. When they apply their measure of wealth, they are able to reconstruct a consumption function with stable parameters both before and after 1985. Their model is able to account for the movements in consumption during both the rise and fall of consumption and house prices. They show that the time series for house prices lead on the consumption path. The model satisfies several criteria for a well specified model. They interpret their results as a wealth effect from the residential market.

2.1 Studies performed on disaggregate data

We have not found any studies that address housing wealth and consumption at a disaggregate data level. However, there are several papers that analyse the relationship between income and consumption. A pioneering study was done by Hall and Mishkin (1982). They used data from the Panel Study of Income Dynamics to study how household consumption depends on permanent and transitory movements in income. They assume a stochastic process for household income and solve for the covariances of different lags of changes in income and the change in household consumption. They demonstrate that the parameters in their model can be solved from the covariance matrix of different lags of consumption and income changes. Their main conclusion is that most of the consumption level can be explained by the PIHRE model. However, some fraction of about 20% seem to be related to the transitory part of income. They approximate consumption with food consumption since these data do not contain much data on consumption.

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[11] The other variables are lagged saving rate, real growth rate in aggregate income, the general price level

Several papers have followed in their footsteps and extended their work. Bernanke (1984) considers household expenditure on automobiles instead of food. He extends the model of Hall and Mishkin to include durable goods. He finds little evidence against the PIHRE model. Altonji and Siow (1987) redo the Hall and Mishkin study in the case where there are measurement errors in the income data. They believe that this may have caused their results. In this work they estimate ordinary regressions instead of covariance matrices. They conclude that when the model allows for measurement errors the PIHRE model can no longer be rejected. There have been several more, and we will not try to make a complete survey.

Mork and Smith (1989) make use of a Norwegian data set similar to ours but from an earlier period. They exploit the richness of the consumption specification in these data and solve the model for all household consumption. They also involve the timing aspects of the different variables in the data set. They conclude that their data are consistent with the basic PIHRE assumption.

3.0 MICROECONOMICS

Our model is based on the standard assumptions of rational agents maximising their utility within some given horizon (possibly infinite). In order to keep the model tractable and simple, we use the time additive quadratic utility function. Along with an assumption of a constant real rate of interest equal to the household rate of time preferences we obtain the random walk behaviour of household consumption expenditure (Hall (1978))[12],

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and average marginal tax rate.

[12] As mentioned in the introduction, we do not believe that the basic results of this model hinge on the exact assumptions about the preferences and the interest rate. However, the model becomes easier to handle and certainly more transparent to the reader.

$$(3.1) \qquad \Delta c_t = \varepsilon_t .$$

This implies that a household plans a constant consumption stream over its horizon. Revisions in this plan are only made as a result of new information about future income possibilities and the revisions equal the innovation in permanent income. The linearity of the model is, of course, convenient. The use of a quadratic objective function implies that certainty equivalence holds true (Theil (1954, 1957) and Simon (1956)). As a consequence, as long as the degree of uncertainty is unrelated to the agent's decisions, we may treat the optimisation problem as one under certainty[13].

Permanent income is usually defined as the annuity of the present value of future labour income and current wealth from other sources.

$$(3.2) \qquad PI_t = A_{T-t} \cdot \Omega_t^{(t)} = A_{T-t} \cdot \left\{ \sum_{s=t}^{T} \left(\frac{1}{1+r} \right)^{s-t} y_s^{(t)} + p_t \cdot h_t + W_t \right\}$$

where W is the value of non-housing wealth, h is the amount of housing capital and p is its relative price. Superscript (t) indicates that this is a planned variable based on the information set the household has at time t. A_{T-t} is an annuity factor given by

$$(3.3) \qquad A_{T-t} = \frac{r \cdot (1+r)^{T-t-1}}{(1+r)^{T-t} - 1}$$

[13] A problem may arise if future yield of capital is unknown, since uncertainty will be related to the amount of capital that the household has, and consequently to previous saving/consumption decisions.

and in the case of an infinite horizon we see that $\lim_{T\to\infty}(A_{T-t}) = r/(1+r)$. Obviously, in order to keep a constant level of consumption we have that the consumption level must equal the level of permanent income (in order to make wealth at time T equal 0),

$$(3.4) \qquad c_t = PI_t$$

Changes in consumption will be a result of innovations in permanent income so that $\varepsilon_t = \Delta PI_t$. In this paper, we are primarily concerned about how innovations in house prices affect the perceived permanent income of households and what consequences they have on household consumption.

3.1 The equilibrium condition in the housing market

One of the main results of this paper is that price changes in the housing market have quite opposite effects on household wealth dependent on whether they own their house or not. In order to comprehend this effect, it is important to understand the basic equilibrium conditions in the housing market. Further, the conditions establish the connection between different price concepts (buying prices, rental rates, user cost) and enables us to discuss price changes in more general terms. However, the basic effects at work here would apply also under more modest assumptions about the housing market. The main issue is that a rise in buying prices is related to an increase in future rental costs.

Further, the housing market is subject to different types of taxes. As long as taxes do not differ between households we might think of prices as being measured after taxes[14].

[14] We do not try to correct prices of tax liabillity in the empirical section.

We assume that households are indifferent between renting and owning their own dwelling[15]. In order to obtain this equilibrium, we must assume that a sufficient amount of perfectly well-informed households face the decision of renting versus owning within each period. We further disregard problems due to future price uncertainty[16] and transaction costs. If we assume a constant rate of depreciation of housing capital δ, we have that[17]

$$
(3.5) \qquad \begin{aligned} p_t h_t &= \sum_{i=0}^{\infty} \frac{1}{[1+r]^i} v_{t+i}^{(t)} h_{t+i} \\ &= \sum_{i=0}^{\infty} \frac{1}{[1+r]^i} v_{t+i}^{(t)} \frac{1}{[1+\delta]^i} h_t = \left\{ \sum_{i=0}^{\infty} \frac{1}{[1+r]^i [1+\delta]^i} v_{t+i}^{(t)} \right\} \cdot h_t \end{aligned}
$$

p and v are the buying and rental prices for one unit of housing capital, while h denotes the amount of housing capital units. The value of a house today must equal the present value of market rents that the house (potentially) obtains in the market. Further, this implies the following relation between buying prices today and rental prices in the long run

[15] We will disregard the possibility that households may own a house different from the one they occupy.
[16] Introducing uncertainty would, apart from violating certainty equivalence, obviously obscure the matters. However, it is an unanswered question whether a risk adverse household should prefer to own or rent its house since uncertainty will apply both to renting and to buying prices.
[17] We let the house depreciate in the formula. Actually, households maintain their houses so that the depreciation is halted or reduced. However, introducing a market and price for maintenance would complicate the presentation and we can not see that it would add much substance to the model. Again, in the empirical section we do not consider depreciation due to age of the house.

$$p_t = \sum_{i=0}^{\infty} \frac{1}{[1+r]^i [1+\delta]^i} v_{t+i}$$

$$(3.6) \qquad = \sum_{i=t}^{T-1} \frac{1}{[1+r]^{i-t} [1+\delta]^{i-t}} v_i^{(t)} + \sum_{i=T}^{\infty} \frac{1}{[1+r]^{i-t} [1+\delta]^{i-t}} v_i^{(t)}$$

$$= \sum_{i=t}^{T-1} \frac{1}{[1+r]^{i-t} [1+\delta]^{i-t}} v_i^{(t)} + \frac{1}{[1+r]^{T-t} [1+\delta]^{T-t}} p_T^{(t)}$$

Buying prices today are the present value of future rental rates before some horizon T and the buying price at this time. Time T can be any future time period. Consequently, we see that if T is set to be t+1 the rental market will reveal the expectations about the next period buying prices held in the market.

$$(3.7) \qquad p_t = \frac{1}{[1+r][1+\delta]} p_{t+1}^{(t)} + v_t$$

$$p_{t+1}^{(t)} = [1+r][1+\delta] \cdot (p_t - v_t)$$

or

$$v_t = p_t - \frac{p_{t+1}^{(t)}}{[1+r][1+\delta]}$$

We see that in order for rental prices to stay positive, $r + \delta + r\delta$ is an upper bound on the growth rate in expected buying prices. A large expected increase in the buying price will be reflected by a very low rental rate so the two will be negatively correlated.

Further, unexpected changes in buying prices will be related to similar changes in the expected future rental rates and changes in the expected buying price at time T. We have that

(3.8) $$p_t - p_t^{(t-1)} = \sum_{i=t}^{T} \frac{1}{[1+r]^{i-t}[1+\delta]^{i-t}} \left(v_i^{(t)} - v_i^{(t-1)}\right) + \frac{1}{(1+r)^{T-t}[1+\delta]^{T-t}} \left(p_T^{(t)} - p_T^{(t-1)}\right)$$

The unexpected increase in buying prices at time t reflects the present value of a similar unexpected rise in the expected path of rental prices until some time T and the buying price level at time T. We will denote the share of the unexpected price change due to increases/decreases in expected rental cost before time T as φ_T. Thus

(3.9) $$\varphi_T \equiv \frac{\sum_{i=t}^{T} \frac{1}{[1+r]^{i-t}[1+\delta]^{i-t}} \left(v^{(t)}(t) - v^{(t-1)}(t)\right)}{p_t - p_t^{(t-1)}}$$

We assume that $\varphi_T \in [0,1]$. The exact value of φ_T depends on how distant the time horizon is and on the price dynamics in the housing market. If a household has an infinite horizon, it equals 1. Further, for a given T, the higher the discount rate r and δ the closer to 1 we get[18].

[18] In order to illustrate possible values for φ, we may assume that households expect a constant rental rate in all future. This will imply the following relationship between the discount rate r + δ, the distance to the time horizon and φ:

The share of an increase in house value that is due to an increase in rental rates before the time horizon T (at time t=0).

r+δ =	.08	.12	.2
φ(T=5) =	.32	.43	.60
φ(T=10) =	.54	.68	.84
φ(T=25) =	.85	.94	.99
φ(T→∞) =	1	1	1

3.2 Permanent income, house prices and non-housing consumption

There are some complicating issues that must be considered before we can relate consumption of non-housing goods to house price movements. First, we have to realise that only surprising changes in house prices are able to change permanent income. Second, in our model we consider only non-housing consumption since none of our households move during the survey period. In addition to the direct change in permanent income, we have a change in relative prices. We have to account for this when we interpret our model. Finally, households adapt their housing consumption at rather infrequent intervals. Our model must somehow take this into consideration.

We will discuss the extent of surprise in the price movements later and first concentrate on the effects price changes are likely to have on household plans. We intend to argue that if households assess their wealth changes in the housing market and change their consumption accordingly, we expect the effects on consumption of non-housing goods to be quite sensitive to whether the household owns a house or not. House-owners receive a capital gain from house price increases[19] that tenants do not benefit from. Although we apply some strong assumptions to obtain a closed form model and testable implications, we believe this difference to be quite substantial even when these assumptions are relaxed.

First we will consider the effects of price changes if housing demand is fixed and unchangeable. Then we will turn to the opposite assumption that households are free to re-optimise housing consumption in every period.

[19] The effects from price increases and decreases are completely symmetric. Thus, we will only discuss price increases.

FIGURE 1.

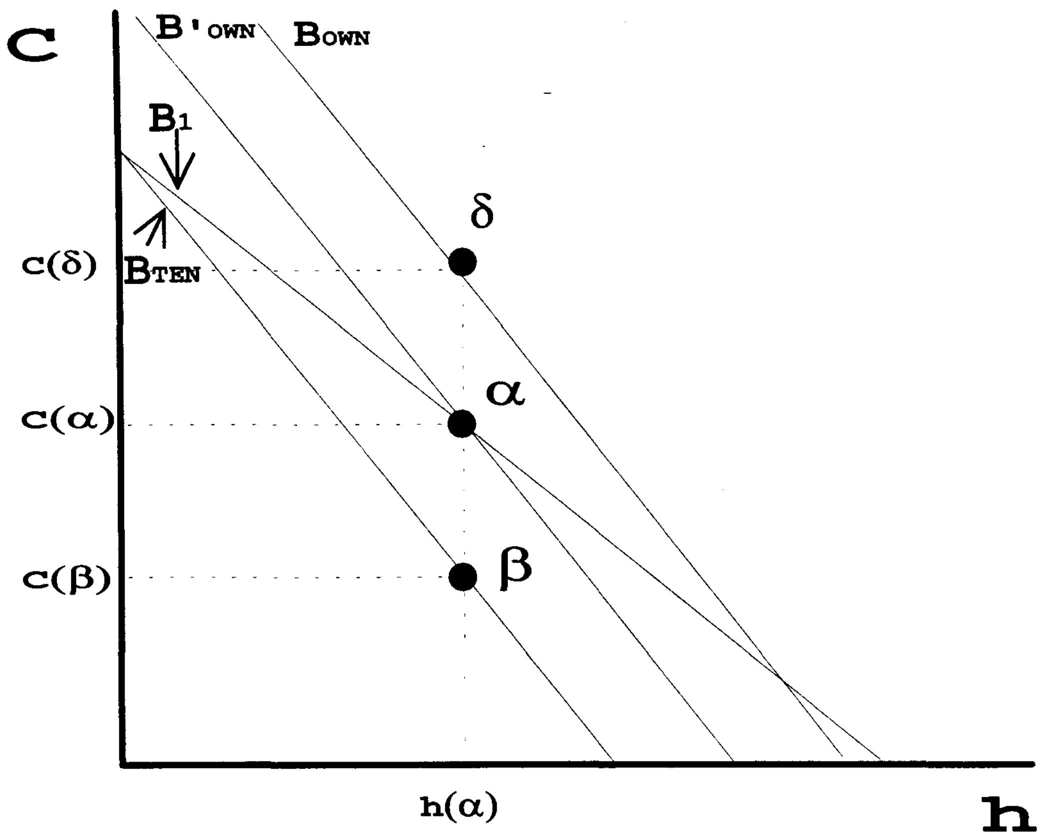
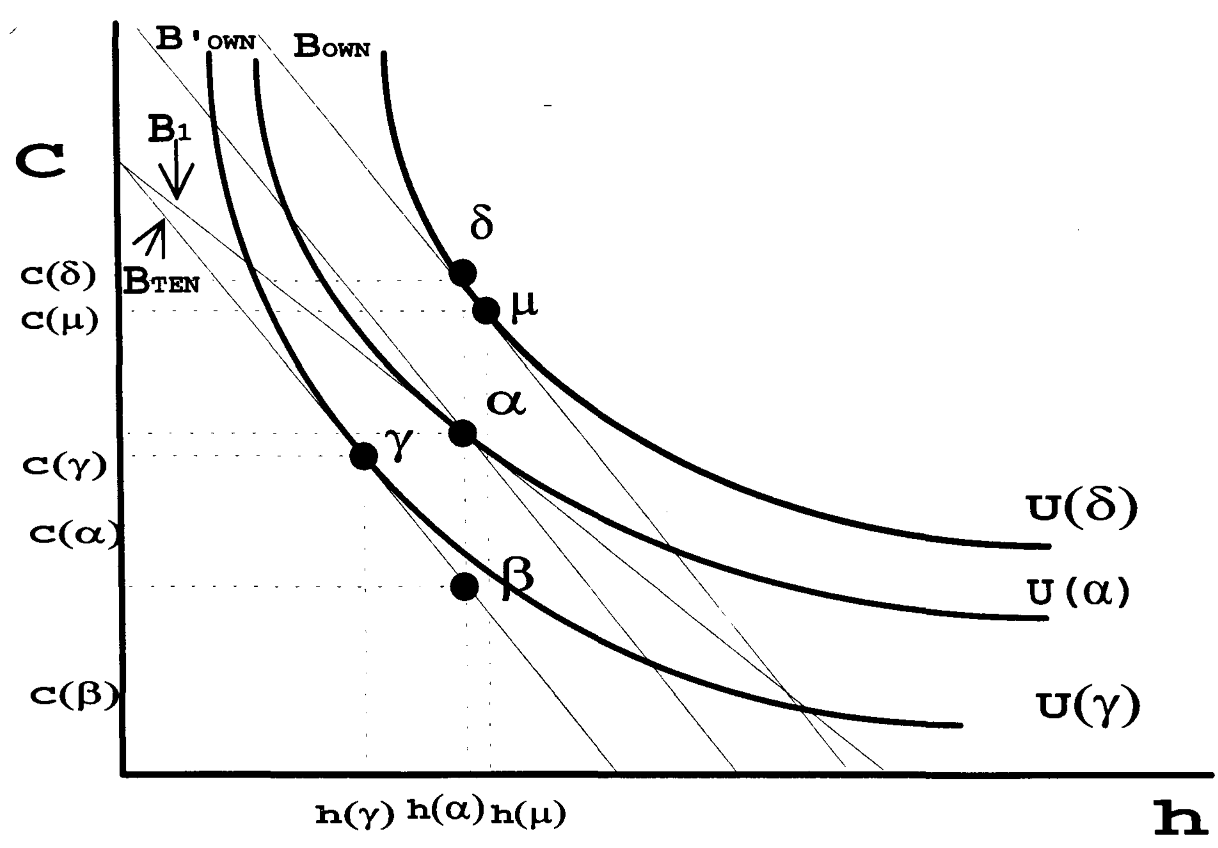


FIGURE 2.



3.2.1 *Assumption 1: housing consumption is fully rigid.*

Let us assume that households are unable to adjust their housing consumption. We may then focus on the wealth effects and disregard the substitution effects. The household continues to live in the same house until its time horizon, T. A possible argument for this, would be that transactions costs are very high[20].

In figure 1, we show the consequences for renting and self-owning households. On the first axis, we have the level of housing consumption while on the other we have the level of consumption of other goods that the household plans. Since planned consumption levels for the two goods are constant, the figures represent the plans that the household makes for all future periods until its horizon. Housing consumption is fixed so that we have no substitution effects and the consequences can be described entirely by budget constraints.

We start at some point α and the budget B_1. For tenants, an increase in house prices will shift the slope of the budget constraint from B_1 to B_{TEN}. In response to the price increase, the household will move from the point α to β. The level of housing consumption will remain unchanged by assumption. The reduction ($c^{\alpha} - c^{\beta}$) in the consumption of other goods will be a result of increased housing expenses. It will equal the permanent income equivalent of the cost increases so that

(3.10) $$c^{\beta} - c^{\alpha} = -A_T \cdot \varphi_T \cdot \left(p_t - p_t^{(t-1)}\right) \cdot h_t \; .$$

[20] Taken rigorously, this assumption is not compatible with the equilibrium assumption for the housing market. Besides direct expenses to a removal company, real-estate agents and governmental fees, transaction costs comprise considerable social costs as well. Consequently, decisions to move are often related to change of job location or major changes in the composition of the family and rarely occur. When they do occur, however, the household may re-optimise their consumption level.

A owner will, of course, face the same cost increases as the tenant. However, he will add a gain to his permanent income from the price increase. Obviously, if $\varphi_T = 1$, he will be able to consume the same amount as before. The budget line will move from B_{TEN} to B'_{OWN} passing through α. The increases in housing costs will exactly match the capital gain from the price rise. The household will consume the same amount of non-housing goods as before. If $\varphi_T < 1$, the wealth increase from the price rise will surpass the increase in housing costs and the budget line will shift further as to B_{OWN}. The household may increase its consumption of non-housing goods from c^α to c^δ as a result of the price increase. The level c^δ is characterised by

$$(3.11) \qquad c^\delta - c^\alpha = A_T \cdot (1-\varphi_T) \cdot \left(p_t - p_t^{(t-1)}\right) \cdot h_t$$

Note that independent of the value of φ_T, the difference between the change in consumption for an owner and a tenant will be

$$(3.12) \qquad c^\delta - c^\beta = A_T \cdot \left(p_t - p_t^{(t-1)}\right) \cdot h_t .$$

It equals the annuity of the change in value of the house. The difference is equal to the vertical distance between the B_{OWN} and B_{TEN} in the figure. When households are unable to adjust their housing consumption to changes in relative prices, we have quite different implications for consumption of other goods whether the household owns its house or not.

3.2.2 *Housing consumption is fully flexible*

Occasionally, households do move. When they do, they are likely to re-optimise their consumption level in light of their permanent income and expected future relative prices. We will conduct the discussion assuming that households are free to adjust consumption in every period (which is probably no more accurate than the assumption of no adjustment possibilities).

The relevant price for the household deciding on the next period's level of housing consumption is the rental rate. In order to keep things simple and tractable, we will assume that the price change is perceived as a once-and-for-all change in relative prices between housing and non-housing goods[21]. As long as households plan for a constant level of consumption, we may continue to analyse the effects in a similar figure.

We see (figure 2) that the substitution effect will imply that the tenant household will prefer the optimal point $\gamma = (c^{\gamma}, h^{\gamma})$ to β. They want to substitute some of their housing consumption by consumption of other goods. Whether or not the new consumption level c^{γ} is higher or lower than the previous level c^{α} depends on whether the own price elasticity of housing demand is more or less than 1. If the elasticity is less than 1, which is a plausible assumption[22], the new consumption level will still be less than before. The tenant will reduce the consumption of other goods facing the new set of prices.

A house-owner will have the same shift in the budget line as before. If housing has an income elasticity of zero, the home-owner wishes to reduce his housing consumption to h^{γ}, the same as the tenant. The difference in consumption of non-housing goods between the two groups of consumers will be the same as in the case where the households were unable to change their housing consumption. More realistically, the income elasticity is positive, and the house-owner wishes to have a housing

[21] If the price change should change the expected time path of relative prices as well, this will add complicating intertemporal effects. However, it is likely to affect home-owners and tenants in a similar manner and should not affect the difference between these two types.

consumption above h^{γ}. This can only happen at the expense of consumption of non-housing goods. This will reduce the difference of non-housing consumption between the two types of households. However, it is highly unlikely that this difference approaches zero. This would imply that non-housing was an inferior good with an income elasticity equal to or less than zero. In general, the new consumption level will be at some point μ where the owner has higher consumption of both goods compared to the tenant.

A reasonable assumption is that housing has approximately the same share of the budget of low and high income households. A constant share implies that the gap between the consumption of non-housing goods becomes

$$c^{\mu} - c^{\gamma} = S_n \cdot A \cdot \left(p_t - p_t^{(t-1)}\right) \cdot h_t$$

where S_n is the share of non-housing consumption of the total budget, at least 60 - 70%. Thus, most of the difference remains even in this case.

As we see, there is no reason to believe that the differences between owners and tenants should disappear even if they are free to adjust their housing consumption within each period. Probably, housing consumption is neither completely rigid in the long run nor is it flexible in the short. The true story should be some mixture of the two situations described above[23].

One possible objection is that tenants and house-owners do not have the same propensity to move. We might believe that there is a self-selection mechanism so that households that expect to move more often will tend to rent their house instead of

[22] A price elasticity less than 1 implies that a tenant facing a price increase on housing will plan to spend a larger share of his life time wealth on housing and less on other goods.

[23] We disregard complementary problems like while the household is waiting for an opportunity to increase their housing consumption level they will prefer to spend extra on restaurants or cars. We cannot see that this will add much substance to the discussion.

owning it. The reason is that some of the transaction costs in the housing market are related to the change of owners, such as governmental fees and costs connected to the use of a real estate agent. In the extreme case, we could think of tenants as being able to adjust the housing consumption level without any costs, while owners are completely stuck with their residences. Yet, we would still expect the changes to be of opposite sign for the two types. Some difference will still remain.

When we consider residential consumption, life cycle considerations might be of some importance. Young households may expect to increase their housing consumption in the future while older households plan a reduction. Consequently, young households will give a greater weight to future housing costs than in the simple constant consumption plan. One might think that φ_T will be relatively larger and possibly greater than 1 for this group. Older households will similarly have a smaller value of φ_T. We will test for different values dependant on age.

3.3 Wealth effects of house price changes on consumption of other goods: conclusions

The reaction to price changes in the housing market depends crucially on the position that the household has in the housing market. The general conclusions are:

1. When housing consumption is fixed, we expect that owners will make a non-negative change in the consumption of other goods. Tenants will make a negative shift.

2. The difference in value between the two types of households in this case is equal to the annuity of the price change.

3. If housing consumption is fully flexible, we still expect the difference between the owners' and tenants' reactions to remain. The difference may be reduced somewhat,

since owners receiving a capital gain will want to keep a higher level of housing consumption than a tenant not receiving it. Tenants are likely to reduce their consumption of non-housing goods. Houseowners will increase this level.

3.4 Were the house price changes unexpected?

As long as price changes are not unexpected, they cannot be assumed to change the household's permanent income. We are then left with a problem of assessing the expectations that households held prior to the price change.

The ideal way to do this in a model assuming rational behaviour on behalf of the households, would be to deduce the expectations implicit in the rental market prices. Unfortunately, the necessary price data are unavailable[24].

As a rule of thumb, annual rents are approximately 10% of the buying price of a house. However, during the sample period, prices rose by almost this amount at several occasions. From 1988 to 1989 it fell with more than 10%. When prices rose, house-owners had almost no costs on housing. If this was expected in the market, the rental market should have disappeared and rental rates gone to zero. If the market is not too imperfect, this reveals that households could only partially have expected the price changes that occurred.

We use the actual real price change as the explanatory variable in our estimations. In some senses it probably does not matter whether we know the exact size of the surprise to the household. If we exaggerate the actual surprise, this will tend to reduce the estimated propensity to consume out of wealth changes.

[24] As far as I know of there were no erratic movements of rental rates indicating the actual price movements that we observed.

To unfold the argument, let us assume that the market did form price expectations different from just stable prices so that $p_{t+1}^{(t)} \neq p_t$. Further, let us assume that all households share the same expectations at time t about house prices in the next period, $p_{t+1}^{(t)}$. What will be the consequences for our results of using our assumption? Instead of using the true unexpected price change $\pi_{t+1}^{true} = p_{t+1} - p_{t+1}^{(t)}$, we make use of $\pi_{t+1}^{used} = p_{t+1} - p_t$ in our empirical work. In general, we have that $\pi_{t+1}^{true} = \pi_{t+1}^{used}(1-\gamma_{t+1})$ where γ_{t+1} measures how well the households guess at time t the real price change that occurs at time t+1. If $\gamma_{t+1} = 1$, the household had perfect foresight while $\gamma_{t+1} = 0$ is the assumption that we use. We want to draw inference about the true propensity to consume out of unexpected changes in housing wealth. Let us call this parameter α. Further, let us assume that our ignorance about household expectations is the only source of bias. It enters the estimation equation as

$$
\begin{aligned}
\alpha^{used} \cdot \left(W_{t+1}^h - W_t^h\right) &= \alpha^{used} \cdot \pi_{t+1}^{used} \cdot h_{t+1} \\
&= \alpha^{used} \cdot \frac{1}{\left(1-\gamma_{t+1}\right)} \cdot \pi_{t+1}^{true} \cdot h_{t+1} \\
&= \alpha^{true} \cdot \left(W_{t+1}^h - W_{t+1}^{(t)h}\right) \\
\Rightarrow \quad \alpha^{used} &= \alpha^{true} \cdot \left(1-\gamma_{t+1}\right)
\end{aligned}
\tag{3.14}
$$

As we can see, this leads to a rescaling of the estimate. As long as the plausible[25] assumption $\gamma < 1$ holds true, our estimated parameter will have the correct sign. Remember that our testing to a large extent is based on different signs of the wealth effect for house-owners and tenants. The rental rates at the time indicate as well that γ could not have been too far away from zero.

[25] This is a plausible assumption. Otherwise, households must have been surprised that the changes in prices were too small. This does not seem reasonable since the price movements in these periods were the most erratic occurring since prior to the Second World War.

If households have had a partial foresight in that they have guessed the correct sign of the change, but not the size ($\gamma \in \langle 0,1 \rangle$) our estimated parameter will be biased towards zero. As we shall see later on, we are more astonished at the size of the parameters than at how close they are to zero. Finally, since we are considering different time periods, the γ that biases our estimate will be a weighted average of γs over time. For instance, if households hold adaptive expectations, we expect that they to some extent made a correct guess of the price change as prices rose but overshot when prices started to fall making the average γ not to far from zero. Presumably, we are able to perform the testing procedure along the lines outlined above[26].

3.4 Permanent income effects from news about labour income

We will use the available data in order to measure the news the households have received about their human capital. As we saw, there were some problems with assessing the expectations the households had about the path of future house prices. These problems are not less when we consider labour income since the information will be more closely related to each household.

We assume that the innovation in permanent income is proportional to the innovation in current income. In order to do so, we will assume that labour income follows a univariate AR(p)-model[27].

[26] It is possible that we have a natural selection mechanism that households expecting a relatively low price increase relative to the market will tend to rent their house while more «optimistic» households will tend to be houseowners. Then tenants will have a larger surprise change in price rising periods and a smaller wealth effect in periods where prices fall. However, it is difficult to see that this might account for the results that we obtain later on.

[27] This is a common procedure. In aggregate data it is used by Hall (1978) and several studies following his. Also previous studies of permanent income apply this assumption. In disaggregate data it is used by Hall and Mishkin (1982) and most of the studies that follow from this work. One might possibly assume an ARMA(p',q')-model (Flavin (1981)). This can be given an AR(∞) representation. Assuming that parameters after some lag p are insignificantly small, this may be approximated by the model above.

(3.15) $$y_t^i = \mu + \rho_1 \cdot y_{t-1}^i + \rho_2 \cdot y_{t-2}^i + \ldots + \rho_p \cdot y_{t-p}^i + \varepsilon_t^y$$

or if we use a lag operator

(3.16) $$y_t^i = \mu + \rho(L) \cdot y_t^i + \varepsilon_t^y$$

where $\rho(L) = \rho_1 \cdot L + \rho_2 \cdot L^2 + \ldots + \rho_p \cdot L^p$.

Let ψ_τ be the parameters defined by the moving average representation of this autoregressive model according to $1/[1-\rho(L)] = 1 - \psi_1 \cdot L - \psi_2 \cdot L^2 - \ldots - \psi_\infty \cdot L^\infty$. Then ψ_τ describes the effect that a shock to the labour income process today has on income τ periods ahead. The change in permanent income from the innovation at time t will be

(3.17) $$\Delta PI_t = A_{T-t} \cdot e_t^y \cdot \left(1 + \sum_{\tau=1}^{T} \frac{\psi_\tau}{(1+r)^\tau}\right).$$

Flavin (1981) shows that this relates to the AR(p)-model according to

(3.18) $$\begin{aligned} \Delta PI_t &= A_{T-t} \cdot \varphi_T^y \cdot \frac{1}{1 - \sum_{j=1}^{p} \left(\frac{1}{1+r}\right)^j \cdot \rho_j} \cdot e_t^y \\ &\equiv A_{T-t} \cdot \Psi_T \cdot e_t^y \end{aligned}$$

where $\varphi_T^y \equiv \frac{\left(1+\sum_{\tau=1}^{T}\frac{\psi_\tau}{(1+r)^\tau}\right)}{\left(1+\sum_{\tau=1}^{\infty}\frac{\psi_\tau}{(1+r)^\tau}\right)}$ is the present value share of the income changes that accrue before time T. So if $T \to \infty \Rightarrow \varphi_T^y \to 1$. The increase in total consumption possibilities is proportional to the innovation in income. The factor depends on the time horizon, the autoregressive parameters for labour income and the discount rate. The change in permanent income is the annuity of this wealth change. Consequences of changes in the budget constraint were discussed in the previous sections.

So far, we have said nothing about the stationarity properties of income. It is quite possible that the time process of individual labour income contains a unit root. In our estimates we shall see that $\sum \rho > 0.8$ in all versions of the model and even above 1 in some. In time series modelling we know that a unit root tends to bias the estimated parameters towards zero and they will be non-normal. Quah (1993) shows that when data are combined time-series and cross-section, the estimated parameters will still be asymptotically normal. However, they will contain a bias. He further shows that when the cross-sectional variation is large, the bias will be small. Thus, the problem should not be important in our analysis and we disregard the possibility of unit roots here.

Finally, in our model we add lagged consumption and wealth $\mu_c \cdot c_{t-1} + \mu_h \cdot p_{t-1} \cdot h_t^i$. The lagged consumption level is included since it reflects the households' expectations about permanent income at time t-1 and as such should be an ideal predictor of future income levels. Further, since we do not have individual information about the constant term, we include the level of housing wealth/consumption since it might reflect some general level of permanent income. It may serve the same purpose as c_{t-1} as well.

4.0 THE BASIC MODEL

Our model consists of two equations that will be estimated simultaneously. The main equation describes the changes in household consumption. The second is the autoregressive formulation of income. The residual from this equation is the "unexpected" income change proportional to the change in permanent income.

The consumption equation will be

(4.1)
$$\begin{aligned}\Delta c_t &= k_1 \cdot D_1 + k_2 \cdot D_2 + k_3 \cdot D_3 + \alpha_1 \Delta CHILD_t + \alpha_2 \Delta MEMBER_t \\ &+ \alpha_3 \cdot \varepsilon_t^y + \alpha_4 h(t)(p_t - p_{t-1}) \cdot D_e + \alpha_5 \cdot h(t)(p_t - p_{t-1})(1 - D_e) + \varepsilon_t^k\end{aligned}$$

Here, the D's are yearly dummy-variables, one for each of the three sub-panels[28] (1986 - 87, 1987-88 and 1988-89). CHILD and MEMBER are the number of children and members in the household. The change in these are assumed to change the level of consumption. Further, ε_t^y is the residual from the income equation while D_e is a dummy-variable assuming the value 1 if the household is a owner and 0 otherwise. Thus, the two last terms are the reactions of the two types of households to changes in the value of their houses.

The income equation is described by

(4.2)
$$y_t = \mu_1 \cdot D_1 + \mu_2 \cdot D_2 + \mu_3 \cdot D_3 + \rho_1 \cdot y_{t-1} + \rho_1 \cdot y_{t-2} + \rho_1 \cdot y_{t-3} + \mu_h \cdot p_{t-1} h^i(t) + \mu_c \cdot c_{t-1}^i + e_t^y$$

where y is private disposable income at different points in time.

[28] We did consider dummy variables for each quarter as well. This did apparently not add much to the results and are left out here. Further, we tried to make the dummies in the consumption function linear in income two years back. This in order to capture differences in household consumption level. However, the results did not make much sense and did not alter much on the point estimates for our main parameters.

The model is estimated simultaneously using the FIML (Full Information Maximum Likelihood) procedure in TSP 4.2. When we want to compare different models we apply likelihood ratio tests when simple t-tests are insufficient. The parameters are expected to be $\alpha_3 = A_{T-t} \cdot \Psi_T$ (from equation 3.18), $\alpha_4 = A_T \cdot (1-\varphi_T)$ (eq. 3.11) and $\alpha_5 = -A_T \cdot \varphi_T$ (eq.3.10). The formulas of α_4 and α_5 are for the case when households may not adjust their housing consumption level. We do expect the two parameters to be of opposite signs and differ considerably also under more general assumptions.

All money denominated variables are deflated by the consumer price index. We consider two different measures of consumption. First, we have total consumption expenditure except for housing, electricity and fuel. This contains almost all consumption apart from housing and as such should be close to our theoretical variable. However, this measure contains investments in durable goods and might be too volatile. Our second measure leaves out investment in durables and avoids this problem. In order to interpret the results, we assume that this is a constant fraction of total consumption of non-housing goods and services. On average, this is a fraction 0.78 of total consumption of non-housing goods and services and consequently we expect the estimates to be smaller in magnitude by a similar factor. We disregard changes in relative prices between durables and other non-housing consumption. Such effects are likely to be caught by the yearly dummy variables.

Since we operate at a disaggregate level, we quite naturally have problems with outliers and influential observations. In order to assess to what extent our results are sensitive to these observations we estimate sub-samples where they are left out. Influential observations are identified through the matrix $H = X(X'X)^{-1}X'$ where X of course is the set of explanatory variables. If h_{ii} is the diagonal element of H belonging to observation i, this observation is influential if $h_{ii} > k / n$, where k is the number of rows in X while n is the number of observations (Kmenta (1986) page 425). An observation is assumed to be an outlier if (Green 1990, page 304)

$$\hat{e}_i = \left| \frac{e_i^k}{\sqrt{\left(s^2 \cdot (1 - h_{ii})\right)}} \right| > 2.5 \qquad (4.3)$$

The process is carried out for both measures of consumption but not for different versions of the model in order to assure that results are comparable across different model specifications. The criteria are quite liberal so that approximately 10% of the sample is removed in these cases. This is done to show that our results appear throughout the sample and are not merely a consequence of some extreme observations.

In our models, we find both skewness and kurtosis in the error term. Thus, our results are reported along with heteroscedastic consistent standard errors. Together with the procedure against outliers and influential observation, we hope that this should be sufficient to assure that this is not a major cause for our results.

4.1 Expected parameter values

Our main hypothesis is for the parameters α_4 and α_5. These are the estimates of the propensity to consume from house price changes. The value of these parameters will be a result of the annuity factor and the parameter φ_T. Consequently, they will be closely related to the distance to the time horizon. In order to give an idea of the expected level of these parameters, we can assume that price changes are uncorrelated over time. We obtain the following values depending on different assumptions on the time horizon and the values of the real interest rate and the rate of depreciation.

Theoretical values of A, α_4 and α_5.

	r =	r = 0	r = 0.04 δ = 0.04	r = 0.04 δ = 0.08	r = 0.07 δ = 0.05	r = 0.1 δ = 0.1
A	T = 5	.2	.225	.225	.264	.334
	T = 10	.1	.123	.123	.163	.239
	T = 25	.04	.064	.064	.110	.202
	T→∞	0	.04	.04	.1	.2
α_4	T = 5		.153	.128	.150	.134
	T = 10		.057	.039	.052	.038
	T = 25		.010	.004	.007	.002
	T→∞		0	0	0	0
α_5	T = 5		-.072	-.097	-.114	-.200
	T = 10		-.066	-.084	-.111	-.201
	T = 25		-.054	-.060	-.103	-.200
	T→∞		-.040	-.040	-.100	-.200

The table applies when households are unable to adjust their housing consumption level. The parameter α_3 may be influenced by general noise. Its theoretical value depends crucially on the values of the autoregressive parameters and the stationarity properties of income. However, it should be positive.

5.0 RESULTS

In table 1, we have given results for the basic model. In the top half of the table we have results for the income equation. Then we have the results for the consumption equation and at the bottom we have some statistics. The results are for the full sample and for a sample where outliers and influential observations are removed. Further, we present results for a model where annual dummies are left out.

Our income relation shows a stable behaviour across time and different sub-samples. The sum of the autoregressive coefficients is stable in the area of 0.8 to 0.9 and all parameters have sensible signs and interpretations. Notably, last year's consumption level appears to be a good predictor of the income level the following year. The ratio of explained variation is stable above 70%. The income equation behaves very similarly in all the remaining estimations and will not appear in the later tables.

The consumption equation shows low explanatory power. This probably indicates that we have a considerable amount of noise in the dependent variable. Still, several of the parameters show sensible values. α_1 and α_2, which are the parameters for changes in the number of children (1) and number of household members (2), seem reasonable. The two parameters show a high degree of multicolinearity. The cost of a new member in the household is positive in all models, but we are not able to single out whether children are more or less expensive than adults. When durables are included, children appear to be more costly but less so when this is excluded. The parameters are simultaneously significant at a 90% level[29].

Owners seem to have a positive propensity to consume (α_4) capital gains in the housing market at about 10% to 15%. This is quite in line with what we expect from our theoretical model. However, it does imply a rather short horizon in the household behaviour. The t-values are 1.3 and 1.4 in the full sample cases[30], so we cannot exclude the possibility of $\alpha_4 = 0$. This is the theoretical value in case of an infinite horizon.

However, the results for tenants, α_5, do question the interpretation of this as a wealth effect. In all cases but one, this estimate is positive and larger in magnitude than α_4. The only exception is the case where we have left out consumption of durables and use

[29] The LR-test observator is 5.4 for total non-housing consumption and 9.6 when durables are left out.

a reduced sample. In this case the parameter estimate drops below zero, saving some hope for the model. The parameter does in no case pass a t-value of 1.21; so at this level firm statements are not easy to make.

α_3, the propensity to consume out of "unexpected" income changes, is positive and quite in line with what we could expect for all groups of consumers. However, it has t-values of less than 1 and is not significantly different from zero.

In table 2, we have split the sample into sub-samples based on the age of the oldest person in the household[31]. If the horizon comes closer as the household head ages, we would expect the parameter values for owners to increase with age. We would expect the difference between tenants and owners to increase as well.

In these sub-samples several of the parameters become significantly different from zero. However, from the point of view of our theoretical model, the results are not easy to interpret. The household with a head less than 60 years old has a positive propensity to consume housing wealth changes regardless of whether it is renting or owning its house. For owners the parameter has a t-value of 1.95 in the full sample (1.46 in the reduced sample) and the point value is very high. For tenants the point values are even higher and the t-value is 1.36 (1.94 in the reduced sample). This clearly does not support the theory.

A possible explanation would be that some common factors drive both house prices and aggregate consumption. Then households having the most expensive houses are the ones that change their consumption level most. Our dummies are apparently not capable of removing this effect.

[30] When we remove influential observations and outliers, the sharpness of the parameters quite naturally drops.

[31] In an earlier version of this paper we estimated this model restraining all the parameters except for α_3 to α_5 to be equal across different age groups. The essential results were almost the same as here.

The results for households aged above 60 do not help us. The point values are significant and negative. The sign of α_5 is consistent with the predictions for the tenant group although the point estimate is very high. The result for the house-owners does not abide the theory nor does it follow the explanation of aggregate effects. However, the results for this group do support the commonly held view that the consumption boom in the period was mainly a phenomenon caused by the young and middle-aged.

The most likely explanation to our results is that our model somehow is misspecified. For example, it could be that there are dynamic effects in how households form their expectations. However, I find it hard to make a story consistent with rational behaviour that would explain the results we get.

The propensity to consume out of income surprises is positive and large and in several cases significantly different from zero. This is less so for the youngest group and higher in older groups. This may, or may not, be consistent with the theory. It depends mainly on the persistence in income innovations over time.

In table 3, we split the sample according to different years (86-87, 87-88, 88-89). We see that there are significant differences between the different time periods. An hypothesis that all the three panels can be merged can be rejected at a high level. It gives a likelihood ratio statistics of 220 compared to a critical value of 31.4 (20 d.o.f) at 5% level. The hypothesis that panels 1 and 2 show a similar pattern can be rejected as well (LR = 45 compared to 18.3 with 10 d.o.f.), but appears somewhat more plausible. Still, the three panels show completely different behaviour according to the model.

Probably, this should not surprise us. The data cover a period with a very shifting environment. It starts in a period with high optimism in consumer sentiments and ends in a period of depression, where many households were unable to repay their debts. Relaxing our assumption that households expect house prices to remain stable, is insufficient to explain the parameter values we see in these models. It would demand very odd expectation schemes on behalf of the consumers and especially between tenants and owners.

In tables 4, 5, and 6, we consider some extensions of the model. In these tables we focus only on the parameters that concern our hypothesis, α_3, α_4, and α_5. The rest of the model shows similar behaviour as above. First, it is possible that the housing market in several rural areas is quite thin. It is not unlikely that in these areas there was a considerable uncertainty about how prices actually had changed. We might think that the wealth effect is less pronounced in these areas. In order to investigate this possibility, we split the sample in more or less populated areas[32]. As can be seen from table 4, no systematic differences apply to these sub-samples. If any, rural areas are the ones that fit the hypothesis best. For total consumption the model is quite consistent with the theory in these areas. This does not fit an assumption that the house price information was less in these areas.

Further, we consider the impact that liquidity constraints may have on our results. In table 5 the propensities to consume are made linearly dependant on a variable we have called "degree of debts". The variable is calculated as

$$(5.1) \qquad DD = \frac{debts - finacial\ wealth - housing\ wealth}{income}$$

The variables in the formula are measured as averages of periods t and t-1. «Degree of debts» enters the equation as $(\alpha + \beta * DD) \cdot x_t$ where α and β are parameter vectors replacing α_3, α_4, and α_5 and $x'_t = \left[\varepsilon_t^y,\ \ h(t)(p_t - p_{t-1}) \cdot D,\ \ h(t)(p_t - p_{t-1})(1 - D_\varepsilon)\right]$ represents both the income surprise and wealth change the household has experienced. As we can see from the table, all β's are positive. This implies that households having a high degree of debt react more to wealth and income changes. For income the parameter has a t-value of 1.7 for total consumption. We would expect house-owners to consume more of increases in housing wealth when they have a large debt burden since it serves as collateral[33]. The estimated parameter confirms this assumption. However, the largest and most significant parameter appears for tenants. Since the wealth price changes affect tenants negatively, this is hard to interpret within the model.

In table 6, we consider the possibility that house prices may affect consumption differently when they go up and when they go down. When prices go up, the house may serve as increased collateral. However, the bank does not call for an immediate repayment as prices go down. On the other hand, if households had adaptive expectations about house prices, previous increases would make the households expect further increases. The drop in prices may have been a big surprise to them. This would imply major revisions of their consumption level. We see that as prices rose the signs of the effects were much in line with what we would expect from the theoretical model. The size of the effect is very high, indicating a short horizon on behalf of houseowners and a long horizon for tenants. When prices drop, house owners apparently do not adjust their consumption level in response. This is consistent with households being liquidity constrained, but not with the wealth model.

On the other hand, tenants reduce their consumption level by a considerable amount as prices go down. This contradicts our hypothesis strongly since they are assumed to face reductions in future housing costs. Consequently, the restrictions on the budget constraint for non-housing goods become less tight and should enable them to increase their consumption in response. Basically, this behaviour of tenants is what violates the model predictions of rational behaviour. Thus, based on this group, we can not claim that consumers in general respond rationally to house price changes.

[32] This is based on the classification given by the Statistics Norway.

[33] It is a question of whether banks increase the pressure on the household when prices fall.

6.0 Conclusions

In this paper we have investigated the impact that changes in house prices have on household consumption expenditure. A theoretical model is elaborated. One of the main conclusions is that rational behaviour implies that households should react quite differently depending on whether they are house-owners or not. This is based on the house owners receiving capital gains/losses from house price changes. Tenants do not. Although, we apply some restricting assumptions to obtain this result, it seems quite safe to assume that this difference is a much more general implication of rational behaviour.

In general, our empirical section does not support the model. House owners comply reasonably well to the model predictions although the implied horizon appears to be short. Also, the responses to price increases and decreases are not symmetric. Our main problem is the behaviour of tenants. Apparently, they have large positive propensities to consume out of price changes in the housing market. This seems to violate rational behaviour since price increases imply increased future housing costs for this group. The main problem seems to be that they reduced their consumption level as the housing market cracked. A possible explanation would be that the wealth effect for tenants is more difficult to observe since it relates to future housing costs. However, it would not explain the significant estimates that we find in some sub-samples.

In our models, we do have problems with lack of explanatory power. Also the estimated parameters are not stable over time. This may raise the question of whether our results somehow are spurious or perhaps a matter of coincidence. However, in large groups (households with head less than 60 years/above 60 years) the propositions from the theory are rejected at a satisfactory level. All together, we think that the evidence presented here does not support the hypothesis that households in general relate to their housing wealth in an economically rational manner. Thus, we tend to reject this possibility.

8. APPENDIX: The data

The data are constructed from three sources. The consumption data come from «Forbruksundersøkelsen» (a Norwegian consumer survey) for the years 1986 to 1989. Expenditure on non-durables and services is the aggregate of food, beaverages and tabacco, clothes and shoes, equipment, health services, travel and transport, education and entertainment, and other goods and services. Housing, light and fuel are left out. The total expenditure measure includes vehicles, boats, recreation equipment, furniture, linen goods, and other durables in addition to the non-durables and services. The expences that enter the narrow consumption definition are measured during a two-week period. The durable part is measured at an annual basis.

The survey is also the source of household location, type and size of house. The price data for houses are published by the Norwegian Association of Real Estate Agents. The price statistics are aggregated to fit our different regions.

The income data are taken from the official values set for taxation. The data are what is called «pensjonsgivende inntekt». We add income across members of the household so that we obtain one household income variable. This information is available for the survey years and the two preceding years.

All price denominated variables are deflated by the retail sales price index.

TABLE 1: BASIC MODEL.

	Total consumption except housing, light and fuel.			Consumption of non-durables and services except housing, light and fuel.		
	Full sample	Full sampl. no annual dummies.	Influential obs. and outliers left out.	Full sample	Full sampl. no annual dummies.	Influential obs. and outliers left out.
Income equation.						
ρ_1	.599 (.069)	.552 (.073)	.633 (.055)	.661 (.071)	.581 (.075)	.672 (.060)
ρ_2	.007 (.094)	.032 (.102)	.235 (.083)	-.000 (.095)	.031 (.106)	.207 (.082)
ρ_3	.201 (.079)	.198 (.087)	.009 (.075)	.203 (.081)	.218 (.092)	.006 (.074)
μ_1	14980 (6606)	19521 (6069)	20345 (6314)	16913 (6497)	22973 (6250)	23040 (6515)
μ_2	31572 (6897)		34856 (6462)	34004 (6818)		37327 (6727)
μ_3	-3175 (7493)		1112 (7075)	-2737 (7521)		3115 (7274)
μ_c	.179 (.027)	.183 (.028)	.123 (.031)	.209 (.081)	.090 (.029)	.146 (.040)
μ_h	.006 (.011)	.006 (.010)	-.007 (.011)	.005 (.011)	.008 (.011)	-.012 (.011)
Cons. equation.						
k_1	-8373 (7109)	-9511 (3892)	-6084 (5029)	-2395 (4721)	-3492 (2491)	2176 (3520)
k_2	-11883 (5860)		-7680 (4891)	-5353 (3611)		-5988 (3278)
k_3	-4252 (9273)		-5685 (7275)	-2715 (6779)		-7234 (5067)
α_1	12603 (12080)	8692 (11750)	-247 (11118)	2312 (9118)	-3614 (9191)	-2130 (7637)
α_2	7041 (9376)	11364 (8949)	16760 (9341)	9839 (6822)	16442 (6768)	12081 (6884)
α_3	.106 (.117)	.073 (.107)	.063 (.095)	.056 (.077)	.020 (.074)	.045 (.060)
α_4	.152 (.115)	.073 (.079)	.096 (.098)	.116 (.081)	.105 (.054)	.041 (.064)
α_5	.264 (.259)	.168 (.233)	.197 (.165)	.199 (.165)	.166 (.152)	-.061 (.113)
R^2(innt)	.755	.728	.818	.764	.737	.807
R^2(kons)	.078	.074	.059	.064	.053	.067
Log lik.	-13817.16	-13850.91	-12165.4	-13606.6	-13645.9	-12296.3
nobs	559	559	502	559	559	514

TABLE 2 : ESTIMATES OF THE CONSUMPTION EQUATION FROM DIFFERENT SUB-SAMPLES BASED ON AGE OF HEAD OF HOUSEHOLD.

	Total consumption except from housing, light and fuel.					Consumption of non-durables and services except housing, light and fuel.				
	AGE [20-40>	AGE [40-60>	AGE [20-60>	AGE [20-60> Reduced sample	AGE 60 +	AGE [20-40>	AGE [40-60>	AGE [20-60>	AGE [20-60> Reduced sample	AGE 60 +
Consumption equation										
k_1	-20670 (15339)	-6975 (11611)	-12705 (9406)	-7647 (6804)	5729 (8269)	-3676 (9104)	-4398 (8578)	-4712 (6360)	835 (4624)	7344 (4640)
k_2	-33151 (12065)	-6844 (8963)	-16899 (7390)	-10939 (6086)	-1426 (7828)	-11087 (6321)	-2137 (6548)	-5104 (4562)	-5027 (4177)	-7646 (4791)
k_3	1760 (16713)	-12722 (15291)	1630 (11539)	-4541 (9297)	-29240 (12398)	10512 (9454)	-16522 (13881)	1723 (8821)	-2925 (6407)	-20111 (7262)
α_1	-25702 (25335)	-6755 (13077)	9477 (13702)	-8946 (12850)	-22669 (27086)	-23000 (15341)	-14053 (12137)	-5871 (10340)	-5906 (9065)	34758 (31554)
α_2	62583 (22173)	5609 (12270)	11635 (12083)	27496 (12785)	6422 (12421)	36111 (12333)	13890 (10379)	18060 (8878)	16382 (9468)	12137 (8808)
α_3	.137 (.184)	.421 (.177)	.395 (.143)	.326 (.117)	.497 (.153)	.069 (.085)	.216 (.152)	.224 (.096)	.185 (.075)	.281 (.101)
α_4	.128 (.229)	.310 (.171)	.269 (.138)	.174 (.119)	-.325 (.150)	.054 (.115)	.195 (.144)	.183 (.098)	.101 (.076)	-.186 (.105)
α_5	.093 (.310)	1.640 (.708)	.436 (.320)	.381 (.196)	-.432 (.238)	.234 (.150)	.940 (.656)	.339 (.207)	.094 (.120)	-.435 (.200)
R^2(kons)	.138	.220	.160	.089	.078	.088	.103	.064	.073	.137
Log lik.	-4829	-5034	-9884.01	-8519	-3879.67	-4730.75	-4996.32	-9751.80	-8608.93	-3794.00
nobs	195	203	398	350	161	195	203	398	358	161

TABLE 3 : ESTIMATES FOR THE CONSUMPTION EQUATION FROM THE DIFFERENT SUB-PANELS.

	Total consumption except from housing, light and fuel.				Consumption of non-durables and services except housing, light and fuel.			
	PANEL 1 (1986-87)	PANEL 2 (1987-88)	PANEL 1 AND 2	PANEL 3 (1988-89)	PANEL 1 (1986-87)	PANEL 2 (1987-88)	PANEL 1 AND 2	PANEL 3 (1988-89)
Consumption equation.								
k	-9897 (7890)	-7083 (7051)	-13350 (7578)	-17116 (13048)	-2354 (5021)	-1200 (4042)	-4867 (4808)	-15854 (10002)
k_2			-9219 (5999)				-3381 (3636)	
α_1	51674 (30669)	8245 (19933)	12540 (17141)	26823 (17780)	13215 (26731)	-13036 (11763)	-6503 (11871)	10956 (15799)
α_2	-5631 (15336)	-18172 (23955)	1241 (11769)	1444 (17731)	13612 (10616)	13739 (10651)	16108 (7463)	8389 (15019)
α_3	.216 (.198)	.879 (.381)	1.295 (.165)	.783 (.183)	-.002 (.150)	.212 (.163)	.451 (.104)	.309 (.149)
α_4	.221 (.196)	.363 (.194)	.344 (.140)	-.076 (.194)	.167 (.135)	.271 (.135)	.235 (.096)	-.083 (.141)
α_5	-.363 (.209)	.868 (.980)	.167 (.569)	.111 (.195)	-.227 (.208)	.893 (.634)	.256 (.377)	-.017 (.153)
R^2(kons)	.124	.217	.167	.151	.054	.114	.066	.081
Log lik.	-4783	-4688	-9493.65	-4236.63	-4709.	-4600.07	-9331.58	-4229.88
nobs	193	192	385	174	193	192	385	174

TABLE 4. ESTIMATE FOR CITIES AND RURAL AREAS, KEY PARAMETERS.

	Total consumption less housing, light and fuel.			Total consumption less housing, light and fuel.		
	Rural areas	Central areas	The three major cities	Rural areas	Central areas	The three major cities
α_3	.394 (.127)	.146 (.194)	.232 (.244)	.284 (.103)	.176 (.121)	.235 (.194)
α_4	.167 (.145)	.100 (.191)	.159 (.265)	.090 (.096)	.123 (.146)	.099 (.205)
α_5	-.102 (.306)	.396 (.359)	.332 (.245)	.067 (.208)	.195 (.290)	.028 (.242)
R^2(kons)	.134	.046	.151	.042	.113	.188
nobs	330	229	89	330	229	89

TABLE 5. LIQUIDITY CONSTRAINTS AND PROPENSITY TO CONSUME.

	Total consumption except housing, fuel and light.	Non durable consumption and services except housing, fuel and light.
α_3	.411 (.181)	.093 (.089)
β_3	.052 (.031)	.008 (.016)
α_4	.269 (.151)	.030 (.094)
β_4	.023 (.015)	.002 (.009)
α_5	.002 (.185)	.031 (.120)
β_5	.275 (.120)	.020 (.040)
Nobs	522	522

TABLE 6. ESTIMATES FOR RISING AND FALLING HOUSE PRICES.

	Total consumption	Consumption of non-durables and services
UP α_4	.382 (.203)	.273 (.140)
UP α_5	-.430 (.300)	-.155 (.196)
DOWN α_4	.045 (.171)	.032 (.120)
DOWNα_5	.451 (.310)	.254 (.229)

Table 7. CARACTERISTICS OF THE DATA.

	SAMPLE	NOBS	MEAN	STD. DEV.	MIN.	MAX.
Panel 1						
$c_t - c_{t-1}$	All	193	-4395	90377	-348189	319271
(Total)	Owners	178	-1908	90577	-348189	319271
	Tenants	15	-33903	85331	-193387	108463
$c_t - c_{t-1}$	All		189	58821	-236160	319640
(Non-dur.and serv.)	Owners		2180	58373	-236160	319640
	Tenants		-23444	61001	-172298	80744
$y_t - y_{t-1}$	All		3838	43468	-153190	158419
	Owners		4584	44528	-153190	158419
	Tenants		-5020	27460	-66267	43411
$(p_t - p_{t-1}) \cdot h$	All		27266	31654	-46365	157043
	Owners		26747	30984	-46365	157043
	Tenants		33423	39507	-3761	157043
Panel 2						
$c_t - c_{t-1}$	All	192	-14625	78312	-436128	187832
(Total)	Owners	167	-11429	70060	-196215	187832
	Tenants	25	-35976	119631	-436128	84162
$c_t - c_{t-1}$	All		-7493	49361	-249955	175103
(Non-dur.and serv.)	Owners		-5885	47141	-142210	175103
	Tenants		-18232	62305	-249955	86964
$y_t - y_{t-1}$	All		19155	34876	-105323	157810
	Owners		18415	33221	-105323	104777
	Tenants		24095	44915	-19495	157810
$(p_t - p_{t-1}) \cdot h$	All		-17358	28500	-91087	56119
	Owners		-17142	28731	-83382	56119
	Tenants		-18804	27424	-91087	38709
Panel 3						
$c_t - c_{t-1}$	All	174	-15680	73251	-289905	239496
(Total)	Owners	154	-15491	74924	-289905	239496
	Tenants	20	-17138	60413	-109568	137320
$c_t - c_{t-1}$	All		-11431	57762	-326533	227621
(Non-dur.and serv.)	Owners		-11528	60325	-326533	227621
	Tenants		-10680	32781	-118868	49338
$y_t - y_{t-1}$	All		-23568	55303	-301158	142432
	Owners		-26697	54695	-301158	142432
	Tenants		519	55386	-174783	121455
$(p_t - p_{t-1}) \cdot h$	All		-66307	37109	-265181	-14825
	Owners		-66992	37145	-265181	-14825
	Tenants		-61030	37353	-140678	-19122

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ESSAY III

EXCESS SMOOTHNESS AND MEASUREMENT ERRORS IN NORWEGIAN QUARTERLY INCOME AND CONSUMPTION DATA.

1.0 INTRODUCTION

When Friedman introduced the permanent income hypothesis (PIH) some 40 years ago, one of its main virtues was its ability to explain why the path of aggregate consumption was smooth relatively to aggregate income. The smoothness of consumption was due to the smoothness of permanent income relatively to current income. Since temporary changes in income were assumed to have only small impacts on life time resources, consumers would save in high income years and dissave in low income years, thus, smoothing the consumption stream. As a matter of fact, the smoothness of consumption had been considered one of the prime "raisons d'etre" for the PIH/LCH.

Deaton (1986), West (1988) and Campbell and Deaton (1989) demonstrated, however, that permanent income was not necessarily less volatile than current income. They showed that whether or not permanent income was more or less smooth than current income depended on the stochastic dynamics of income. Testing on US data, they concluded that there was reason to believe that permanent income was actually less smooth than current income. So the permanent income/life cycle theories were not able to explain the observed smoothness of consumption. Consumption showed *excess smoothness*.

Here our task is two-fold. First, Campbell and Deaton based their model on a pure permanent income theory. We shall investigate a model introduced by Galí (1990) and elaborated by Lønning (1994) where Life Cycle considerations are accounted for. We intend to show that this would smooth consumption relatively to the pure PIH model used by Campbell and Deaton. Further, this model is in accordance with the excess sensitivity properties showed by Flavin.

Second, we intend to redo the study of Campbell and Deaton in order to establish the behaviour of Norwegian data. Do Norwegian data show excess smoothness/sensitivity of any sort? Theory ought to have some general validity across demographical differences and across time. It certainly is of interest to consider the overall validity of Campbell and Deatons' results when the model is applied to other sets of data. Further, we have updated the US data to compare the results to the Norwegian results.

When we started this work, we had reason to believe that consumption in Norway is far less smooth than in the US compared to data for income[1]. This is in fact true. However, we conclude that this is probably due to an insufficient quality of Norwegian seasonally adjusted data. When we consider annual data we see that the Norwegian data behave much like in the US.

2.0 THEORY

2.1 The Permanent Income model tested by Campbell and Deaton

Hall (1978) and Flavin (1981) combine the PIH with rational expectations (RE) and are able to establish a firm mathematical representation of the implicit ideas of the PIH. Hall uses a quadratic utility function:

[1] I would like to thank Professor Knut A. Mork, who first made me aware that this might be the case for Norwegian data. This idea first motivated this paper.

1. $$U = -\frac{1}{2}\sum_{i=t}^{\infty}\frac{1}{(1+\rho)}(c^* - c_i)^2$$

to represent the household preferences. Here, c denotes the consumption level and * marks the «bliss level». ρ is the conventional rate of time preference. The use of a quadratic utility function assures that certainty equivalence holds and that marginal utility is linear in the consumption level. This implies a convenient form of the Euler equation:

2. $$c_{t+1}^{(t)} = c^* \cdot \left(\frac{r-\rho}{1+r}\right) + c_t \cdot \left(\frac{1+\rho}{1+r}\right) = (c^* - c_t) \cdot \left(\frac{r-\rho}{1+r}\right) + c_t$$

Superscript (i) denotes that this is an expected/planned variable based on the time i information set. r is the universal real rate of interest assumed to be constant and deterministic. If we apply the common assumption that r = ρ, and add uncertainty of future income streams, we obtain the simple and convenient form that household consumption should be a random walk.

3. $$\Delta c_t = \varepsilon_t$$

The household plan for a constant level of consumption and only news about future income possibilities may change this path to a new level. ε_t is related to the change in life time resources.

The present value of future income prospects can be written as

4. $$\Omega_t^{(t)} = W_t + \sum_{i=t}^{\infty} \left(\frac{1}{1+r} \right)^{i-t} y_i^{(t)}$$

W represents current non-human wealth[2], y is labour income and Ω is the present value of future expected income streams and represents the life time resources available to the household. The permanent income is defined as the annuity of Ω given by the annuity factor A. As long as we have an infinite horizon $A = r/(1+r)$.

5. $$PI_t^{(t)} = \frac{r}{(1+r)} \cdot \Omega_t^{(t)}$$

A plan for a constant level of consumption implies that $c_t = PI_t^{(t)}$. The news that arrives to the household about future income prospects, the innovation in Ω, in a period is given according to

6. $$\xi_t = \sum_{i=t}^{\infty} \frac{\left(y_i^{(t)} - y_i^{(t-1)} \right)}{(1+r)^{i-t}}$$

and in order for the intertemporal budget constraint to hold we have that

[2]If efficient capital markets exist, we have that W is the «best» estimate of the present value of the future rents and interest earned by this capital. We will assume that future interest rate r is known with certainty so that news that accrues to the household about future income possibillities concerns the value of human capital. If we have uncertainty about future yield of capital we can no longer trust certainty equivalence to hold. The reason for this is that the degree of future uncertainty becomes a decision variable for the households through their savings decisions. Of course, if we include investments in human capital, we must assume an uncertain return to this investment as well. By convension, these problems are disregarded.

7. $\varepsilon_t = \frac{r}{(1+r)} \cdot \xi_t$.

The shift in the consumption path equals the innovation in permanent income. Furthermore, the consumption decision has consequences for savings as well (Campbell (1987)). Savings are defined as capital income plus labour income less consumption,

8. $s_t = r \cdot W_t + y_t - c_t$.

If we replace $W_t = W_{t-1} + s_{t-1}$ we get that

9. $\Delta s_t = r[W_{t-1} + s_{t-1}] - r \cdot W_{t-1} + \Delta y_t - \Delta c_t$

so that

10. $s_t - \Delta y_t - (1+r) \cdot s_{t-1} = -\Delta c_t = -\frac{r}{(1+r)} \varepsilon_t$

The above relationships have several testable implications. Campbell and Deaton make use of relations 7 and 10 above. A weak implication of these equalities is that the volatility of Δc_t and the expression $s_t - \Delta y_t - (1+r) \cdot s_{t-1}$ should equal that of permanent income. We can rewrite 7 as

11. $\operatorname{var}(\varepsilon_t) = \operatorname{var}\left(\frac{r}{(1+r)} \cdot \xi\right)_t$.

The left hand side of this equation may be evaluated either through the volatility of consumption (eq.7) or by the fluctuations in the saving expression on the left hand side of equation 10. Campbell and Deaton demonstrate that US consumption is less volatile than permanent income evaluated from the disposable labour income and label it "excess smoothness". Later, it has also become known as the "Deaton paradox". Flavin (1981) investigates an implication of expression 3. The random walk property of consumption implies that consumption is orthogonal to lagged information. She demonstrates that consumption is, in fact, sensitive to predictable changes in income and thus violates the orthogonality condition.

Flavin developes a formula that links an ARMA-representation of the income process to its permanent income equivalent. If income is represented by

12. $$y_t = \rho_0 + \rho_1 \cdot y_{t-1} + \rho_2 \cdot y_{t-2} + \ldots + \rho_j \cdot y_{t-j} + e_t + \delta_1 \cdot e_{t-1} + \delta_2 \cdot e_{t-2} + \ldots + \delta_i \cdot e_{t-i} +$$

where e is i.i.d. then the innovation in permanent income at time t will be

13. $$\Delta PI_t = \left(\frac{r}{(1+r)}\right) \cdot \frac{1 + \sum_{i=1}^{\infty} \left(\frac{1}{1+r}\right)^i \delta_i}{1 - \sum_{i=1}^{\infty} \left(\frac{1}{1+r}\right)^i \rho_i} \cdot e_t$$

When we formulate an AR(1)-model in first differences, $\Delta y_t = \beta_0 + \beta_1 \cdot \Delta y_{t-1} + e_t$, we have that $\rho_1 = 1 + \beta_1$, $\rho_2 = -\beta_1$ so that $\Delta PI_t = [(1+r)/(1+r-\beta_1)] \cdot e_t$. Volatility can then be measured by

14. $$\sigma_{PI} = \frac{(1+r)}{(1+r-\beta_1)} \cdot \sigma$$

where σ is the standard error of innovations in current income, e.

This procedure may be criticised for not including all relevant information that the households may use to form their expectations about future income and exaggerating the income "news" that they receive. Campbell (1987) shows that the household saving summarises the household income expectations and, thus, is an ideal predictor for this variable. Campbell and Deaton apply this in a VAR-model that exploits this property of saving. We performed a similar test on Norwegian data but it did not give much new insight into the data behavior. In order to simplify the exposition this is left out of the final version

2.2 Life cycle considerations and excess smoothness

Life-Cycle effects may be a possible (partial) explanation for the excess smoothness results of Campbell and Deaton and the excess sensitivity of Flavin. In Lønning (1994), we demonstrated how life cycle considerations may affect the dynamic stochastic properties of aggregate consumption. We intend to show how the same model may affect the excess smoothness problems discussed here. Galí (1990) discusses excess smoothness in his paper and finds that excess smoothness follows from his model. However, he finds that the effects must be small. We show how the degree of excess smoothness relates to central parameters in the model and argue that Galí probably undervalues the effects due to too small parameter values.

In the first essay we show that a model involving limited horizon on part of consumers and a life-cycle saving motive through decaying income over the life and an AR(1)-model in income, implies that aggregate consumption has an ARMA(2,1)-representation. This may be transformed to an MA(∞)-representation.

15. $$c_t = \Psi_0 \cdot e_t + \Psi_1 \cdot e_{t-1} + \Psi_2 \cdot e_{t-2} + \ldots + \Psi_q \cdot e_{t-q} + \ldots$$

where e_t is the innovation in the income process and is assumed non-autocorrelated[3]. The income level is assumed to be an AR(1)-model. The parameters are related according to

16. $$\Psi_1 \equiv (1-p)\cdot\Psi_0 + \Gamma\cdot\rho\cdot\Psi_0, \qquad \Psi_2 \equiv (1-p)\cdot\Psi_1 + \Gamma\cdot\rho^2\cdot\Psi_0,$$
$$\Psi_3 \equiv (1-p)\cdot\Psi_2 + \Gamma\cdot\rho^3\cdot\Psi_0, etc...$$

ρ is the AR(1)-parameter of the income process. In the Campbell and Deaton framework income contains a unit root which implies that $\rho = 1$. Ψ_0 and Γ are transformations of the basic parameters in the model. $\Gamma = \alpha + p - \alpha\cdot p$, where p is the probability of dying within a certain period. α is a parameter that describes the reduction in income over the life cycle. This parameter describes the importance of the life cycle savings motive. α is far less than 1[4] while 0 is the extreme case where no reduction occurs and the life cycle productivity change and the savings motive disappear. If households are infinitely lived p = 0, while more realistically p is a small number about 0.005-0.01 (a quarterly mortality rate) or 0.02-0.04 (an annual mortality rate)[5]. In Lønning (1994), we show that $\Psi_0 = z/(z+\alpha)$ when ρ = 1[6]. z is the real rate of interest adjusted for the probability of death and is given by $z = (1+r)/(1-p) - 1$. If households have an infinite horizon p and α will be zero so that $\Delta c_t = e_t$. This is the usual result from the PIHRE model. Taking first differences of consumption, we have that

17. $$\Delta c_t = \Psi_0\cdot e_t + \alpha\cdot(1-p)\cdot\Psi_0\cdot e_{t-1} + \alpha\cdot(1-p)^2\cdot\Psi_0\cdot e_{t-2} + \alpha\cdot(1-p)^3\cdot\Psi_0\cdot e_{t-3} + \ldots$$

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3 That is $e_t = y_t^{(t)} - y_t^{(t-1)}$ and $\text{cov}(e_{t,i}, e_{t+j}) = 0, \forall i \neq j,\ i,j \in N$.

4 α=1 implies that the household has all its income in the first period and only dissaves in future periods.

5 From the Poisson distribution we know that the expected life time of a household is equal to the expected remaining life time of all living households and is equal to 1/p. It is probably most reasonable to scale p according to the average remaining life-time of households of perhaps 30 years. Gali applies very small values of α and p. In his work p is based on expected life time rather than expected average remaining life time. α is calibrated from aggregate savings data contingent on p and is positively related to the chosen p (see table 5 p.447).

6 In this model we treat income as an AR(1)-process in income and when we assume a unit root process, this leaves us a random walk for income.

The volatility of consumption expressed as the variance of the change in aggregate consumption will be

18.
$$\begin{aligned}\operatorname{var}(\Delta c_t) &= \Psi_0^2 \cdot \sigma_e^2 + \alpha^2 \cdot (1-p)^2 \cdot \Psi_0^2 \cdot \sigma_e^2 + \alpha^2 \cdot (1-p)^4 \cdot \Psi_0^2 \cdot \sigma_e^2 + \alpha^2 \cdot (1-p)^6 \cdot \Psi_0^2 \cdot \sigma_e^2 + \ldots \\ &= \Psi_0^2 \cdot \sigma_e^2 \cdot \left[1 - \alpha^2 + \alpha^2 \cdot \frac{1}{p \cdot (2-p)}\right]\end{aligned}$$

In the permanent income case $\alpha = 0$ so that $\Psi_0 = 1$. In this case we have that $\operatorname{var}(\Delta c_t) = \sigma_e^2$. Thus, whether or not we should expect excess smoothness from the life cycle model depends on whether the expression

19.
$$\Psi_0 \cdot \sqrt{\left[1 - \alpha^2 + \alpha^2 \cdot \frac{1}{p \cdot (2-p)}\right]}$$

is less than 1 or not.

There are two major effects that separate this model from the PIHRE-model. First, innovations in aggregate income will change the path of future wages. But since worker productivity declines over time, future changes in the wage rate will have less impact on life time income of existing generations. This is reflected in Ψ_0 depending on α. Second, the generations that die out in the period are replaced by new generations. These do not start out in direct proportion to the previous generations but will scale their consumption to the future expected income level. This will in general depend on all future innovations to the income path, and makes the current change in consumption depend on all lagged income innovations. Of course the number of households that is replaced by new-born households in each period depends on the number of new households (= probability of death), p.

In figures 1 and 2 we have plotted the degree of excess sensitivity that we might expect from this model. The degree depends on p and the size of the life-cycle motive α. We have computed values of expression 19 for different plausible values of α, the life cycle saving parameter, given reasonable values for the mortality rate. In the paper we investigate the excess smoothness both on annual and on quarterly data. Reasonable values for quarterly data can be p=0.005 and p=0.01 and interest rates r=0.015, and four times as much for annual data[7]. α is a parameter measuring the geometric decline in productivity of the labour force. On an annual basis, this should reasonably be below 0.1, and a quarter of that size for quarterly data.

We see that the model is able to explain some smoothness in the consumption path. In an area of α below 0.1 (annual frequency) and 0.025 (quarterly frequency), the model clearly indicates excess smoothness. The degree of smoothness is clearly below 1 and for several interesting values of α less than 0.8. This applies both when we use annual and quarterly data. This implicates that life cycle motives may very well be a reason for the consumption path being too smooth from the view of permanent income theory. However, as we shall see, the data indicate an excess smoothness of consumption of above 50% on both frequencies. Thus, left alone, the Life-Cycle model is insufficient for explaining the excess smoothness problem.

Further, from the MA-representation of aggregate consumption, we see that consumption is positively related to lagged innovations in income. This might be a part of the explanation of the excess sensitivity presented by Flavin. The "slow adjustment"-property of consumption that Campbell and Deaton find, is consistent with this model. It is due to the fact that some of the future wealth effects of income shocks are related to the wealth of yet unborn generations.

2.3 Problems with aggregating from a single household model

Excess smoothness is a property of aggregate data. However, the theory is based on a single household or consumer. This poses a special problem and might be yet an other reason why

[7] A more liberal interpretation of the parameters in the model would be that p is the horizon of the household. Then a myopic household may have a much shorter horizon than indicated here and p and r would be larger.

we obtain the excess smoothness result. The simple argument is that aggregate income contains an insufficient amount of information in order to assess the expectations that the household has about future income prospects.

The point may be illustrated in the following manner. Let us assume that individual households form their expectations according to an autoregressive model $y_t^i = \mu + \rho^i \cdot y_{t-1}^i + \varepsilon_t^i$[8]. However, we are estimating $\sum_i y_t^i = \mu^i + \rho \cdot \sum_i y_{t-1}^i + \mathrm{E}_t$ using aggregate data. As a consequence, we have that $\mathrm{E}_t = \sum_i \varepsilon_t^i + \sum_i (\rho^i - \rho) \cdot y_{t-1}^i$. The variances of E_t and $\sum_i \varepsilon_t^i$ will be equal only if $\rho^i = \rho,\ \forall i$, else we have that $\mathrm{var}(E) > \mathrm{var}\left(\sum_i \varepsilon^i\right)$. The use of aggregate data will tend to exaggerate the volatility of permanent income. The reason is that too much of the innovation in the income level is recorded as news to the households. The inclusion of lagged savings cannot be assumed to fully correct for this problem.

Further, a similar effect will not arise on the consumption measure since this is only related to the change in the level of consumption. We know the plan for future consumption for the household at both individual and aggregate level. This aggregation problem does not, however, account for the excess sensitivity problem of Flavin. There is nothing in this aggregation that should cause aggregate consumption to be predictable from lagged information. Consumption should still be a random walk.

2.4 Other tests of excess smoothness

West (1988) investigates the same phenomenon as Campbell and Deaton and is often referred to as a co-inventor of the excess smoothness paradox. He applies a different testing framework. However, he reaches exactly the same conclusions for US aggregate data. He

[8] The reasoning can of course be extended to more than one lag in the autoregressive model or an ARMA-model. The assumption that ρ differs between households seems reasonable since sensitivity to business cycles varies across sectors and professions

shows that private information on the part of households does smooth income. However, there are considerable limitations to how much this can smooth consumption and it cannot explain the degree of smoothness that we observe.

Some studies try to assess the generality of the findings of Campbell and Deaton. To a certain extent they apply different testing procedures and partly estimation is done for other countries. Galí (1993) performs tests on 6 different OECD-countries. He tests for both non-durable and durable consumption separately. Based on his testing framework he concludes that durable consumption shows strong excess smoothness. For non-durables which are more easy to compare with Campbell and Deaton, he finds that USA, Canada, UK, and Italy all show significant (1%-level) excess smoothness in quarterly consumption data. For Japan and France, he finds weak evidence of the opposite. France is just significant at 5% level.

Patterson (1995) tests a VAR-model similar to Campbell and Deaton on data for UK. He finds little evidence of excess smoothness and the measures of volatility in income and consumption are very close. Although the cross-country evidence is not decisive, there seems to be some general support to the excess smoothness paradox.

2.5 Different explanations for the excess smoothness results of Campbell and Deaton in the literature

Several authors have tried to modify and redo the Campbell and Deaton study in order to explain their results. Among them are Caballero (1990) (precautionary saving), Christiano (1987) (endogenous interest rates), Christiano, Eichenbaum and Marshall (1991) (time aggregation bias), Diebold and Rudebusch (1991) (uncertainty in the estimate of permanent income), Normandin (1994) (precautionary saving), Quah (1990) (permanent as well as transitory shocks in labour income), Caballero (1995) (near-rationality and heterogeneity), and Pischke (1995) (individual households versus aggregate data). In parenthesis, we have indicated their explanations.

Caballero (1990) and Normandin (1994) both emphasise the need for a precautionary savings motive in the utility function. The model used by Campbell and Deaton is based on a quadratic utility function and consequently certainty equivalence is forced onto the system. This excludes the fact that households adopt their consumption decisions to future uncertainty and behaviour like precautionary saving. Zeldes (1989) shows that introducing the constant relative risk aversion (CRRA) utility function in itself is insufficient to explain the findings of Campbell and Deaton and Flavin. However, if heteroscedastisity is present in the innovations in labour income, this might for some parameterisation of the model explain the empirical problems met in aggregate data. Normandin applies a different form of the utility function (hyperbolic absolute risk aversion). He finds that the excess sensitivity and excess smoothness problem is what we should expect from some reasonable parameterisations of this model.

Christiano (1987) analyses an integrated model of consumption, investments and total output in a closed economy. He studies the effect of a positive shock to productivity and show that consumption adjusts only gradually to a new level. The reason behind this, is that the increase in productivity increases the return to investment. Consumption is consequently smoothened by a positive correlation between interest rates and shocks in the economy. However, his model has considerable problems explaining other empirical facts of aggregate data.

Christiano, Eichenbaum and Marshall (1991) are concerned with the way the data are measured. Testing on data that are time aggregates implies that consumers are assumed to receive news and reoptimise at the same intervals as data are collected. If, however, consumers optimise on a more frequent or on a continuous basis the use of time aggregates would imply a time aggregation bias. This result was originally shown by Working (1960) and is likely to imply a positive moving average term of .25 in the limit. The authors use an assumption of continuous optimisation to show that in this case the behaviour of US aggregate data need not be far away from what we would expect given these assumptions. This may account both for the excess smoothness and for the excess sensitivity problem in the data.
Diebold and Rudebush (1991) and Quah (1990), both challenge the way permanent income is modelled. Diebold and Rudebush consider a univariate model for income that allows for fractional integration. Constructing confidence limits for the permanent income they show that these limits are wide enough to include areas where the excess sensitivity problem vanishes. A

fractional model is more likely to capture the long run dynamic, while an ARMA-model often reflects short run dynamics.

Quah (1990) makes the assumption that labour income is a result of two stochastic processes. One is the non-stationary element that Campbell and Deaton use and the other is a stationary process. In this setting he shows that consumption is likely to appear excessively smooth and it is no problem to redo the Campbell and Deaton results well within the frameworks of the PIH. Still, his model cannot account for the excess sensitivity property of Flavin.

Caballero (1995) explains the excess smoothness by assuming near-rational behaviour of the consumers. They are assumed to change their consumption level as they get too far from their optimal level. Since households tend to hit the trigger point at different times this causes a smoothing of aggregate consumption. The hypothesis can account for the observed movements in data.

Pischke (1995) assumes that individual households are mainly concerned with their private information since this is much more important to them than movements in aggregate income. When they project own future income they lack knowledge of the information in aggregate data. News at this level are therefore gradually passed on into household decisions. Thus, consumption smoothing results, because individual income information is less persistent than aggregate income news.

3.0 DATA

We have constructed data for Norway and USA for the purpose of this study. The sources of the data are the National Accounts data for Norway and the National Income and Product Accounts (NIPA) for the US. The data are constructed along the lines of Blinder and Deaton (1985) for the US. Their data were the same data that were applied by Campbell and Deaton in the excess smoothness study. The American data have been expanded to the end of 1994. They are used as comparison to the Norwegian results.

Blinder and Deaton's main series were total aggregate consumption, non-durables and services consumption, and disposable labour income. Consumption of semidurables like shoes and clothes are left out of the nondurables and services consumption expenditure measure, along with the consumption expenditure on durables.

The separation of capital and labour income is, of course, important when we think of how the theory was formulated. Capital income is assumed to be non-stochastic and consequently, cannot change permanent income. A natural way to think of this assumption is to consider it a lower bound on the change in permanent income. Since the authors were primarily concerned with the excess smoothness of consumption, this was a natural procedure. All of their data series have been seasonally adjusted prior to any analysis.

3.1 The US data

Blinder and Deaton (1985) construct their income and consumption data from the NIPA-data. In their paper they describe the changes and elaborations they do to the official data. We have tried to reconstruct their series. Our data show a very similar pattern to theirs and the main difference is probably that they have removed the tax rebate in the second quarter of 1975 and the similar increase in the third. We had no information on how they did this and in our study we have used dummies for these two quarters. The data series are labour income, total consumption expenditure and expenditure on non-durables and services, shoes and clothes excluded. Our data date from 1953:1 to 1994:4. To make the results comparable to the Norwegian data, we shall restrict ourselves to the period 1970:1 to 1994:4. Testing the period 1953:1 to 1985:4 gave results very close to those of Campbell and Deaton. Annual data are just the sum of the four quarters.

The quarterly US data are plotted in figures 3, 4, and 5 at the end of the paper. The figures show data at levels and in first differences. We can see that the income data appear to be more volatile than the consumption data. The volatility in the data appears to be rather constant over time.

3.2 The Norwegian data

Our Norwegian data are constructed from the national income accounts. Statistics Norway has since the beginning of 1978 published quarterly data for disposable income and its components. However, for internal use they have constructed data dating back to 1970 for all the necessary elements. At the other end, we have data until 1994:4 making a total of 25 years and 100 quarterly observations. Our measure of disposable income is the same as that of the Bureau. Labour income is defined as disposable income less interest received on bank deposits and dividends, adding interest paid on loans. Taxes are not divided into taxes on capital income and labour income in the statistics. Most of the taxes are related to labour income and we have allocated all of the taxes to this variable.

Our consumption data cover the same period. We apply two measures of consumption similar to the measures used by Campbell and Deaton. The first is total consumption expenditure including durables. The second is a measure of the expenditure on nondurables and services. In the second series, we have removed durables and semidurables including clothing and footwear. The reason for using two measures is that none of them equals exactly the theoretical variable. We want to have data for total actual consumption. However, total consumption expenditure includes investments in consumer durables and is likely to be more volatile than actual consumption. The consumption of non-durables and services does not cover the whole consumption spectrum. All series have been deflated by an appropriate consumer price index. To make the narrow consumption measure representative for the overall consumption, it is rescaled by a factor (mean of total consumption expenditure)/(mean of nondurables and services expenditure) 1.351. Implicitly, we assume that this narrow consumption measure is proportional to total actual consumption and is equally volatile in percentage terms.

Our data have been constructed from seasonally unadjusted data. These series have been seasonally adjusted by the Statistics Norway according to their standard procedures. The method is based on ARIMA-X11.

Plots of the seasonally adjusted data series are in figures 6, 7, and 8. The consumption data show a boom beginning in 1985:1 and peaking in 1986:2 but staying high until fall 1987. The income data show no similar movement at the time resulting in an extraordinary negative private saving rate from 1985:4 to 1987:4[9]. Consumption appears to be more volatile in 1980 and 1981 and in the last years, and less volatile prior to 1980. Before 1978, the seasonal pattern in the consumption data was constructed in a more mechanical way and shows a much more stable pattern. We investigate the influence this might have on our results. Income does not seem to reveal any clear pattern. Possibly, there is an increase in volatility over time. Data are analysed in real terms to avoid any kind of non-stationarity inherit in the price data.

We will compare the results from Norwegian annual and quarterly data with similar results for the US. Originally, the annual data were intended as a confirmation of our results for quarterly data, but as we shall see there are considerable discrepancies in the results and conclusions when the frequencies are changed. Our annual data are just the annual sum of the quarterly nonadjusted series. We did try to extend our annual series back to 1963. In doing this, we had to apply a somewhat different measure of disposable income. These income data showed odd behaviour, however, and led us to doubt the quality of these data[10]. Thus, our effective sample is 1970 to 1994 at both frequencies. We have only 25 annual observations, but our models are very simple and parsimonious with few estimated parameters so that this should be sufficient.

4.0 RESULTS

We use a very simple testing procedure equal to the first analysis done in Campbell and Deaton. They did extend their analysis to a more sophisticated logarithmic VAR-model in income and saving based on equation 10. As said before, we have conducted a similar analysis on our Norwegian data. However, the results did not add new insight and are dropped from this final version of the paper. Consequently, we compare the volatility of consumption with that of the permanent income equivalent of labour income in simple univariate models.

[9] Net capital income is negative in Norway so that total disposable income is some 3% lower than labour income on average.

[10] The inclusion of this period did not in itself change the results to any substantial degree. However, the correlation pattern was quite different.

As did Campbell and Deaton, we have removed the general trend in the data by adding a constant term to the regression. That is, volatility of consumption is measured by the standard deviation of ΔC around its mean.

In order to make US and Norwegian results comparable, we have divided all our consumption series by the mean of total consumption and multiplied by 100. Consequently, the mean of our consumption series is 100 (since the narrow consumption measure is rescaled to total consumption). Income is rescaled accordingly. The volatility may then be interpreted in percentage terms of total consumption expenditure.

Finally, we should mention that the use of total disposable income raises the volatility in Norwegian income by about 10-15% and will not enter the analysis. The saving expression in eq. 10 is approximately the same as the change in consumption and the results are almost identical. None of these variables are listed in the results tables and graphs. In the figures section at the end of the paper, we have included several recursive plots of estimates. First, we show the stability in the estimated AR(1)-parameter in the univariate representation of income. Then we have similar figures for consumption, and finally we include some plots of the time volatility measures over time.

4.1 Results concerning excess smoothness in US data

The excess smoothness property of US quarterly data was thoroughly demonstrated by Campbell and Deaton. In order to obtain estimates of the volatility in permanent income based on the labour income data, a univariate specification of income is needed. Campbell and Deaton test their data and conclude that the data can be assumed stationary in first differences. Taking a unit root for given, they studied different ARMA(p,q)-representations of labour income changes but concluded that an AR(1)-model seemed to do as well as any. Our results for US are based on an AR(1) model in first differences of the data.

For the simple framework applied here, Campbell and Deaton estimated that permanent income measured by disposable labour income was 2.85 times as volatile as changes in the nondurables and services consumption, and the factor was 1.65 compared to overall consumption expenditure. In Table 1 we have replicated these results for the period 1970:3 to 1994:4. The factors are for this sample 2.78 and 1.75. Annual data very much confirm these results and conclusions. The equivalent factors become 2.37 and 1.53 and the autocorrelation pattern in the data appears to be much the same.

Consistent with Campbell and Deaton, the tax rebate in 1975:2 and the following increase in 1975:3 are handled through dummy variables in the autoregressive model for income. There is a similar fluctuation in the beginning of 1993 and this is treated the same way. Finally, in the annual data the boom in 1973 and the consecutive bust in 1974 tend to dominate the estimate of the AR(1)-parameter and turn this negative. In the data, this appears to be a one-time occurrence and is neutralised by a dummy in 1974. This can be defended by reference to the exceptional increase in oil prices in the fall of 1973. OPEC II did not have a similar effect on the data.

The results seem to be quite stable over time. In figure 9, we have plotted recursive estimates for the autoregressive parameter in labour income changes. The AR(1)-parameter starts just above 0.3 and ends near 0.25. This is somewhat less than for Campbell and Deaton's sample where they obtained 0.44.

In figure 17, treating quarterly US data, we see the volatility in labour permanent income[11] at top and somewhat above the estimate for volatility in current income. The difference is due to the positive estimate of the AR(1) parameter. At the bottom, we have the volatility estimate based on changes in consumption of non-durables and services. We have included the standard deviation together with its 95 % confidence limits. The confidence limits are based on $(n-1)\cdot S^2/\sigma^2 \sim \chi^2_{n-1}$ where S is the estimated standard deviation of consumption changes and σ is the "true" volatility. As we can see, the estimates show a large degree of stability over time.

The estimates for annual data repeat much of the story from the more frequent data (figures 10. and 18.). The estimates seem to be stable over time. The difference between current

[11] Based on equation 14.

income fluctuations and consumption volatility is somewhat reduced, but a higher estimated value of the AR(1)-parameter for labour income compensates for the lower volatility in income. The AR(1)-parameter is somewhat reduced over time and leads to a reduction in the volatility in permanent income from labour income data over time.

4.2 Analysis of time series properties of the consumption and income data for Norway.

One of the basic assumptions of the Campbell and Deaton paper was that labour income is stationary in first differences instead of trend stationary in level. A natural way for us to start would be to investigate the stationarity properties of disposable labour income in Norway in order to check the validity of this assumption.

In table 2, results of an augmented Dickey-Fuller τ-test for a unit root are given for income and first differences of income[12]. The number of lags k in the augmented Dickey-Fuller regression is decided upon by starting from a high number of k and reducing until the last is significant. When data are used in levels, a trend and a constant term are included, while first-differenced data are only tested using a constant term. The results are quite clear in favour of a unit root present in the level of income. All the estimated test values are far below the 10 % critical value of -3.15. We cannot reject the null hypothesis of a unit root. Further, the series seem to be stationary in first differences since all of the estimated values are far above -3.51, which is the 1% critical value without a trend included. The unit root is consequently supported by the evidence and the income series appear to be integrated of order 1.

Ghysels and Perron (1993) have shown that preadjusting the data for seasonality may bias testing for a unit root towards accepting the null hypothesis. As a control, we have included tests on the seasonally unadjusted data as well. The tests are based on Hylleberg, Engle, Granger and Yoo (HEGY, 1990). The HEGY-test included seasonal dummy variables. The relevant significant tables are included in their papers. When dummy variables and trend are

[12] The ρ-test $(n^*(\rho-1))$ and the comparable Phillips-Perron-tests were performed as well without influencing on the conclusions drawn from the tests referred.

included the 5% critical values are -3.53 (Y1), -2.94 (Y2) and 1.94 for the F-test for the combination of the two last parameters. When no trend is included, the critical test values for Y1 decrease to -2.95 (-2.63 at the 10%-level). The results from these tests are given in the same table and confirm that a unit root seems to be present in levels at both the zero and the biannual frequency. However, we cannot reject unit roots in first differences as well. This may be due to seasonal unit roots which are not removed by first differences or may be a consequence of a lack of power on behalf of these tests. At least, a unit root in the level seems well supported by the evidence. In the seasonally adjusted data, the first differences apparently contain no unit root.

From the table we can further see that the same conclusions apply to the consumption data. However, no unit root appears on the seasonal frequencies and consequently, unit roots are rejected in the first-differenced data. Since a unit root in the consumption level was one of the implications of the theory, this is reassuring for the analysis we perform here[13].

We have tested for unit roots in the annual data as well. When a trend is included, the critical 5%-value is -3.60 (-3.24 at 10%) (25 observations) and without a trend -3.00 (-2.63). In the income data, we see that the conclusions remain the same. The test value is well below the 10% critical value in all the series measured in levels. The first differenced data give a test observer above the 5%-critical value. The results for the consumption data seem to be unfavourable to the unit root hypothesis in levels. For total consumption the unit root hypothesis is close to being rejected at the 10%-level while for non-durables and services it can be rejected at the 1%-level. This is consistent with previous results by Lønning (1994) doubting the unit root in aggregate consumption data. Note, however, that the results are sensitive to the number of included lags and the Phillips-Perron test (-2.86) is not rejecting. The first-differenced consumption data appear stationary in all tests.

In the remainder of this paper, we shall accept the assumption that all series are I(1) in levels.

[13] A test for unit root can of course never become a verification of the unit root hypothesis since apparently reasonable stationary alternatives can be infinitely close. In addition, the unit root tests are not particulary powerful and tests performed on annual data by Lønning (1994) are not quite so supportive to the unit root hypothesis in consumption.

4.3 Univariate representation of labour income.

In order to represent the permanent income equivalent of changes in current income, it is necessary to specify the news that arrives about current and future income. First, we investigate the simple univariate forecasting of labour income. This is the conventional procedure and implies the assumptions that all the household information about future income can be summarised in the history of aggregate income, and that the relationship between the income history and the future can be well approximated by a linear relationship.

When we investigate the autocorrelation functions (actual and partial) of first differenced data, a spike appears at lag one while further lags indicate little additional correlation. This indicates clearly an AR(1) or MA(1)-model. In Table 3 we have included the results from different ARMA-models fitted to the series. As we can see the results very much reflect the impression from the correlation functions. As was the case for the Blinder-Deaton data, first differences of the income series seem to be best represented by an AR(1)-model or alternatively an MA(1)-process. However, as opposed to the results of Campbell and Deaton, the correlation seems to be negative in the quarterly data. Consequently, permanent income should be less volatile than the news in current income. The estimated parameters were significant in all cases.

Annual data show a clear positive autocorrelation. This indicates that the short run data misrepresent the long run persistence of income changes. Further, it seems hard to give any good explanations for the difference that occurs based on intuition or theoretical analysis. Based on annual data, permanent income appears to be far more volatile than current income.

Annual and quarterly data differ in which model is best. While for quarterly data MA(1) and AR(1) appear to be equally good, the AR(1) model seems to be better when we consider the annual data. This emphasises the temporal aspect of the fluctuation in the quarterly data. Later, we shall argue that the difference may stem from measurement errors in the quarterly income data. That an MA(1) representation seems to be more likely in quarterly data gives some credence to this interpretation.

In the table, results for an ARMA(2,1) and ARMA(1,2)-model have been included. As we can see, the estimated parameters become much less sharp and almost no increase in explanatory power results, indicating a problem of colinearity in the model. Furthermore, we can see that for quarterly data the estimates of the standard deviations in the estimated permanent income vary only slightly between the different models, so the choice of ARMA-representation of the income process appears to be unimportant for the conclusions we shall be making. In annual data, we had problems with convergence in these extended models. However, the resulting increase in explanatory power was low. We have tried other specifications within the ARMA(2,2) framework as well. All other models seemed worse than the ones we give here. In my investigations, I shall concentrate on the AR(1)-models.

The estimate of the AR(1)-parameter is quite stable. We show recursive estimates together with +- 2 standard deviations. In the quarterly data, the estimate hovers around -.2 and most of the confidence intervals seem to be on the negative side. This is quite in contrast with the results for USA. For annual data, the estimate is close to 0.6 all the time and the confidence interval is above zero all the way. When we consider how different Norwegian and US data behaved when measured on a quarterly frequency, the similarity in annual data is quite striking.

4.4 Results of excess smoothness testing on Norwegian data.

The Norwegian quarterly data do not behave like their corresponding US series. This applies not only to the income data, but is perhaps even more apparent in the consumption data. Based on the results in table 1, Norwegian consumption as measured by the narrow definition appears to be more than 3.5 times as volatile. For total consumption the factor is 2.3. The difference in volatility between total consumption and nondurables and services is much less in Norway, almost negligible in fact. In the US the broad consumption measure is approximately 60% more volatile. Labour income in Norway is about 50 % more volatile than in the US and is approximately equally volatile as consumption. However, the negative AR(1)-parameter makes permanent income less volatile than consumption. If any, this indicates that we have excess volatility in Norwegian consumption. This is probably not significant since the AR(1)-parameter is not significantly different from zero.

In annual data, we see once again how similar these data seem to be to the American data. Norwegian data show excess smoothness by a factor of 2.04, which is seemly close to 2.37 in the US. Consumption of non-durables and services is somewhat more volatile than in the US and labour income a little less. However, a larger estimate of the AR(1)-parameter compensates for this. The main difference between the two countries is the volatility of total consumption where Norwegian consumption is almost twice as volatile. Based on this consumption measure, Norwegian data still show excess smoothness, but to much smaller degree. However, all together, the Norwegian annual data seem to be yet a convincing argument in favour of the excess smoothness paradox.

One might expect that the results have arisen from particular events in the data or in the economy. Consequently, in figures 19 and 20 we have included the recursive estimates for the volatility measures. They are quite stable over time and confirm that the results we have discussed are valid throughout most of the sample. In quarterly data we see that the volatility in nondurables and services increases in the beginning of 1980 to the present level. The reason for this increase is probably the mechanical way in which Norwegian data were constructed prior to 1978. Using a seasonal filter on these data has made them artificially little volatile. Note, however, that the correlation pattern in the data is stable also across these sub-periods so that it is mainly the overall volatility that has changed. The volatility in current income is approximately the same as for consumption, but since the autocorrelation is negative the volatility in permanent labour income is less than that of current income indicating excess volatility. The results are not significant and we cannot exclude that the results are consistent with the theory. The annual data show a pattern that is very similar to the US and indicate clear evidence in favour of excess smoothness.

4.5 Conclusions concerning excess smoothness in Norwegian data

As we have seen, Norwegian data led to quite contradicting conclusions concerning the «Deaton's Paradox». Quarterly data suggest a Norwegian paradox in that although the labour income process was assumed to be non-stationary, this does not quite enable us to explain the high volatility in the consumption data. However, at the annual frequency, this is reversed and there seem to be no discrepancies between Norwegian and US data from a theoretical point of view. A simple explanation to this could be that the long run properties of income are misrepresented in the quarterly data and consequently we underestimate the persistence in the income innovations. This could be due to short run movements in labour income that dominate the more long run underlying process in the estimation of a univariate income model. Thus, true permanent income is more volatile than it appears from quarterly data.

We believe the problem to be more profound than that. The high volatility and odd behaviour of Norwegian data seem to indicate a problem of measurement errors in the short run movements of the data. In the next chapter we investigate and document this further. We focus on the consumption data and try to rule out other possible explanations for the behaviour of these data. It follows that we have much more confidence in the results from the annual data and we conclude that the excess smoothness property is confirmed in the Norwegian data.

If we have measurement errors in both income and consumption data this will have quite different effects on the measurement of permanent income from the two variables. While measurement errors in the consumption data tend to exaggerate the volatility in permanent income, this is not necessarily the case for aggregate income. Although measurement errors will increase the volatility in current income, it will bias the AR(1)-parameter negatively. This will tend to underestimate the persistence in the income innovations and will consequently reduce the estimate of permanent income volatility.

5.0 FURTHER INVESTIGATION OF THE NORWEGIAN CONSUMPTION DATA

In this section, we investigate some possible arguments that suggest that we overestimate the variability of private consumption. First, however, we focus on further properties of Norwegian data pointing towards measurement errors. If we have measurement errors, we should expect a clear negative correlation in the first difference of the data. We have already seen that this is the case for the income series. This is in contrast with evidence from annual and US data which all show a significant positive value for this parameter.

If we estimate an AR(1)-model in consumption of non-durables and services, we get that

20.
$$\underset{}{\Delta c_t} = \underset{(.15)}{0.88} - \underset{(.093)}{.409} \cdot \Delta c_{t-1} + e_t$$

The data seem to have a strong negative correlation. Again, if we compare with results from annual data the estimate is 0.424 and in US data it is 0.338 and 0.318. According to the theory consumption should be a random walk, so the estimate violates this restriction in all cases. However, the positive values we find in annual data are quite consistent with what we find in other countries and are related to the excess sensitivity of consumption to lagged values of income, as described by Flavin (1981). The negative value in the Norwegian data must be considered odd.

Further, we cannot see that the negative value is related to specific events or data points. The recursive estimates show that the estimated value is stable over time (figure 15). When we talk to people at the Statistics, they often point out the last half of the 1980's as a period when they had severe problems with the underlying statistics, primarily the retail sales index. However, our results do not point at this period as a particularly problematic period. The degree of measurement errors, if we interpret the AR(1) estimate that way, shows a stable pattern in this period. Also the more mechanic data construction prior to 1978 apparently does not change this picture of stability.

One obvious argument in favour of the negative correlation is that our consumption data contain investment in durable goods. From a theoretical point of view, durables cause a negative correlation in household consumption expenditure[14]. Mankiw (1982) elaborates this idea. Total consumption expenditure obviously contains durables. Furthermore, our data are quarterly rather than annual making goods «more durable» in the sense that they are expected to last for four times as many periods as for yearly data[15]. However, the way we have constructed the data for non-durables and services should reduce this problem to a minimum in these data. Mankiw tests the hypothesis on data for US durable consumption expenditure and rejects it decisively. If this was to be the explanation for some of the results that we obtain here, it is reasonable to assume that durables would show an even clearer negative value. However, if we make the same autoregression as for nondurables and services, we get an AR(1)-parameter of 0.134[16]. This evidence clearly is against this explanation.

When we test our basic hypothesis, we apply consumption of nondurables and services as a proxy for the overall consumption level. Thus, we implicitly assume that this is some constant fraction of total consumption. However, a subgroup of consumption is likely to be sensitive to changes in relative prices. Consequently, the variation in non-durables and services consumption expenditure may not result from varition in the overall consumption level only. Changes in relative prices could have caused additional movements and variation. The correlation between changes in durables and changes in nondurables and services consumption is only 0.157, and adds to our suspicion that this might be the case. On the other hand, this low correlation could perfectly well be a result of independent measurement errors in the two types of consumption goods.

In order to investigate the price assumption, we have estimated a regression with Δc_t as the dependant variable and $\Delta\left(P_{DUR} \cdot c_{DUR} / P_{ND}\right)_t$ as the explanatory variable along with a constant term[17] (DUR is the part of consumption not comprised by non-durables and services). We

[14] Aggregation from individual households to aggregate data implies an extreme coordination of information across households and is perhaps very critical to this hypothesis.

[15] Changes in the stocks of nondurable goods are likely to have similar effects as durable expenditure.

[16] The interpretation of this would be that the problem of measurement errors in durable consumption is less severe. This is perhaps natural since the data for some durables like cars are not based on surveys but rather on more complete sales numbers.

[17] An argument for such a regression could be the assumption of a Cobb-Douglas utility function between the two subgroups of consumption. This implies constant budget shares and leads to the regression above. However,

may then repeat the previous testing procedure using the residuals as the proximating variable for the change in total consumption level. We obtain a regression like (standard deviation in parenthesis)

21. $$\underset{}{\Delta c_t} = \underset{(.15)}{0.55} + \underset{(.11)}{0.40} \cdot \Delta\left(\frac{P_{DUR} \cdot c_{DUR}}{P_{ND}}\right)_t, \qquad \sigma = 1.44, \quad DW = 2.97.$$

Additional lags of the explanatory variable were unable to contribute to the explanation of Δc_t and reduce the estimated standard deviation of the equation. The estimated effect of the relative prices has the correct sign and is significant with a t-value of 3.64. The standard deviation is reduced from 1.54 when no relative prices are included, to a value of 1.44. Although we find a reduction, 6.5% is not sufficient to explain the discrepancies that we have found earlier. The Durbin-Watson indicates that the negative correlation persists and if we include a first order lag of the dependant variable we get an estimated coefficient of -.43.

Without further investigation it is difficult to tell where the errors arise. The seasonally adjusted retail sales data show a negative correlation pattern of the same type as seen here, but smaller in magnitude. The two series do not cover exactly the same subset of consumption and we cannot compare directly. However, the quarterly data are adjusted in order to be consistent with annual data. This is done as to preserve the overall seasonal pattern. As we saw in Lønning and Mork (1994), this was quite successful for the long run seasonal pattern. However, it is not unlikely that it has caused additional errors in the short run information of the data. It is possible that the long run consistency takes place at the expence of the short run information.

---

this is not completely consistent with the assumption we have made for the intertemporal utility function. Alternatively, the regression can be considered a linear approximation. Measurement errors in the price level would have a similar effect. We tried using $\Delta\left(\frac{P_{DUR}}{P_{ND}}\right)$ as explanatory variable as well giving approximately the same results.

All together, we have problems making up a plausible story for the behaviour of Norwegian quarterly consumption data. There seem to be little besides measurement errors, that can account for the observations we have made here. But also the disposable labour income data show a similar sign of measurement errors. Although we have not discussed these data in any detail, the negative autocorrelation at lag 1 does not increase our confidence in these data. Probably, we have a similar problem here.

6.0 CONCLUSIONS

When households have an infinite horizon and behave according to the PIHRE-model, this has implications for how volatile consumption should be relatively to income. If households have a limited horizon and life cycle saving is important, we should expect aggregate consumption to be excessively smooth compared to this model. The Life-Cycle model we present indicates an effect of about 20% decrease in the volatility. However, this is not quite sufficient to explain the actual degree of excess smoothness observed in the data, so the explanation of the paradox cannot rely entirely on this model.

Further, we have shown that quarterly Norwegian data behave quite differently from similar US. data. The difference is particularly suspicious since annual data show a rather similar behaviour. The first order correlation in the data seems unreasonable and we have attributed this difference mainly to measurement errors in quarterly data in Norway. The existence of durable goods and variations in relative prices cannot account for the observed behaviour. The measurement error problem explains why quarterly Norwegian data seem to indicate a reverse Deatons Paradox of excess volatility.

It is worth noting that we cannot exclude measurement errors in US data. Wilcox (1992) has an extensive discussion on this subject. For that reason, and since we do not know that Norwegian households behave in exactly the same way as US households, we cannot treat the US data as the true pattern that Norwegian consumption should mimic. However, when we compare with the behaviour of annual data, the US quarterly data give more confidence. Further, there seem to be an abundance of theoretical explanation for a positive autocorrelation in consumption changes. The opposite assumption seems much harder to explain.

We conclude that the results from quarterly data are unreliable and base our conclusions on annual data. The overall conclusions are that Norwegian data show a similar pattern to the US. We find clear evidence of excess smoothness in aggregate consumption.

7.0 THE DATA

Norwegian consumption expenditure data are taken from the National Accounts data before the latest revision. The income data come from Husholdningenes Inntektsregnskap. Labour income is defined as: Lønn + Driftsresultat + Stønader fra utlandet + Stønader fra offentlig - direkte skatt - andre utgifter. Total consumption is taken directly from the statistics while consumption of non-durables and services is defined as: Matvarer (00) + Drikkevarer og tobakk (11) + Elektrisitet (12) + Brensel (13) + Driftsutgifter (14) + Helsepleie (62)+ Offentlig transport (61) + Andre tjenester (60). Entry codes at two-digit level are in parenthesis. The series have been seasonally adjusted by Statistics Norway. Labour income is deflated by the implicit price index from the seasonally adjusted data for total consumption expenditure.

The US data are taken from the National Income and Products Account and are made along the lines of Blinder and Deaton (1985). The labour income data are:

Wage and Salary Disburse (A) + Other Labour Income (B) + Transfers paid to persons - Interest paid by persons to business - Personal Contributions for Social Insurance + share of (Proprietors income with inventory valuation and capital consumption adjustments - (Personal taxes and nontax payments - Personal tax and nontax receipts, federal and local)). The share is (A+B)/(A + B + Rental income of persons with capital consumption adjustment + personal dividend income + Personal interest income).

The consumption data are the same as the definition in the statistics with two adjustments. Nontax payments from persons to government were not deducted from disposable income and is added to the consumption data. Further, Clothes and shoes are reclassified as durables. The income data and total consumption data are deflated by a consumer price index for total consumption. The narrow measure of consumption is deflated by a similar price index.

All data have been normalized so that the mean of the consumption series is 100. The income series are scaled according to total consumption.

Table 1: Volatility in the basic series, sample is 1970.3 to 1994:4.

		σ ΔC(nd)	σ ΔC(tc)	σ ΔI	σ ΔPI
Norway	Quarterly	1.54	1.57	1.45	1.18
	Annually	1.72	3.28	1.54	3.51
US	Quarterly	.431	.686	.926	1.20
	Annually	1.13	1.69	1.91	2.67

Table 2. Tests for unit roots in the Norwegian series.

	DF τ-test	k	HEGY Y1	Y2	F(Y3∪Y4)	k	DF τ-test annual data.	k
Total consumption expenditure.	-2.78	3	-.089	-3.00	5.41	5	-2.46	2
First difference of tot. consumption exp.	-4.09	2	-3.44	-3.01	5.37	4	-3.70	1
Expenditure on non-durables and services.	-2.93	4	.04	-4.24	5.64	5	-3.91	2
First difference of exp. on non-durables and services.	-9.69	1	-3.67	-3.32	4.26	4	-3.39	1
Labour income.	-1.68	1	-1.07	-1.42	2.65	5	-2.51	1
First differences of labour income.	-12.1	0	-2.41	-1.45	2.59	4	-3.35	2
Disposable income.	-1.74	1	-0.85	-1.42	2.99	5	-2.33	1
First differences of disp. income.	-11.8	0	-2.15	-1.45	2.94	4	-3.25	2

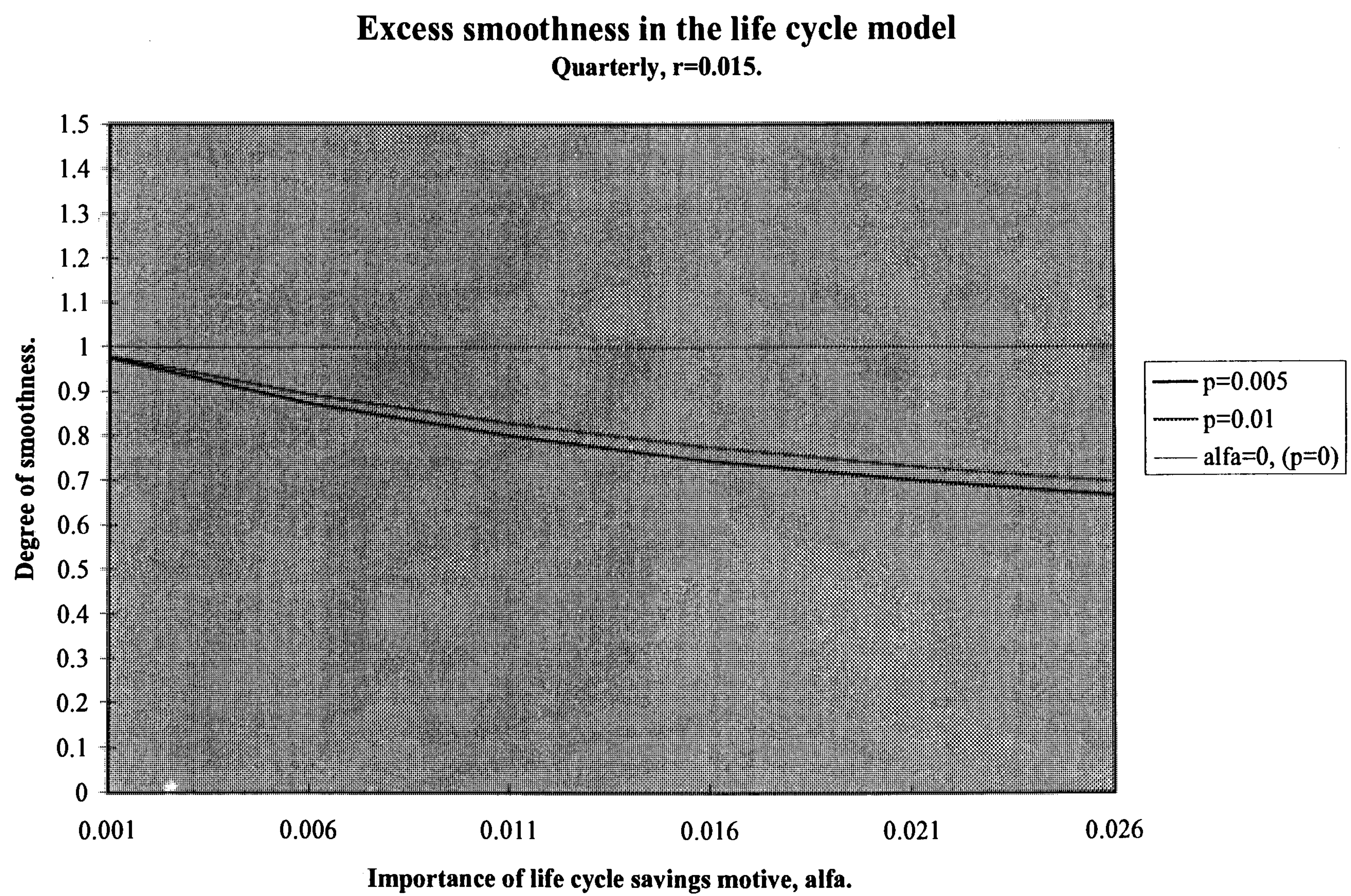
Table 3. Different ARMA-represtations of first differences of Norwegian Disposable Labour Income.

		AR(1)	AR(2)	MA(1)	MA(2)	stand. dev. current income	stand. dev. perm. income	R^2
	ARMA(1,0)	-.227 (.099)				1.45	1.18	.052
Quar Data.	ARMA(0,1)			-.236 (.099)		1.36	1.04	.053
	ARMA(2,1)	-.868 (.383)	-.144 (.132)	.653 (.380)		1.45	1.20	.059
	ARMA(1,2)	-.613 (1.78)		.372 (1.79)	-.114 (.475)	1.46	1.14	.055
	ARMA(1,0)	.596 (.176)				1.54	3.51	.354
Ann Data	ARMA(0,1)			.462 (.194)		1.59	2.28	.276
	ARMA(2,1)	probl. of conver-gence						
	ARMA(1,2)	probl. of conver-gence						

Table 4. Different ARMA-represtations of first differences of Norwegian Disposable Labour Income.

		AR(1)	AR(2)	MA(1)	MA(2)	stand. dev. current income	stand. dev. perm. income	R^2
Quar Data .	ARMA(1,0)	-.227 (.099)				1.45	1.18	.052
	ARMA(0,1)			-.236 (.099)		1.36	1.04	.053
	ARMA(2,1)	-.868 (.383)	-.144 (.132)	.653 (.380)		1.45	1.20	.059
	ARMA(1,2)	-.613 (1.78)		.372 (1.79)	-.114 (.475)	1.46	1.14	.055
Ann Data	ARMA(1,0)	.596 (.176)				1.54	3.51	.354
	ARMA(0,1)			.462 (.194)		1.59	2.28	.276
	ARMA(2,1)	probl. of conver-gence						
	ARMA(1,2)	probl. of conver-gence						

Figure 1.



Excess smoothness in the life cycle model
Quarterly, r=0.015.
Degree of smoothness.
1.5
1.4
1.3
1.2
1.1
1
0.9
0.8
0.7
0.6
0.5
0.4
0.3
0.2
0.1
0
0.001
0.006
0.011
0.016
0.021
0.026
Importance of life cycle savings motive, alfa.
p=0.005
p=0.01
alfa=0, (p=0)

Figure 2.

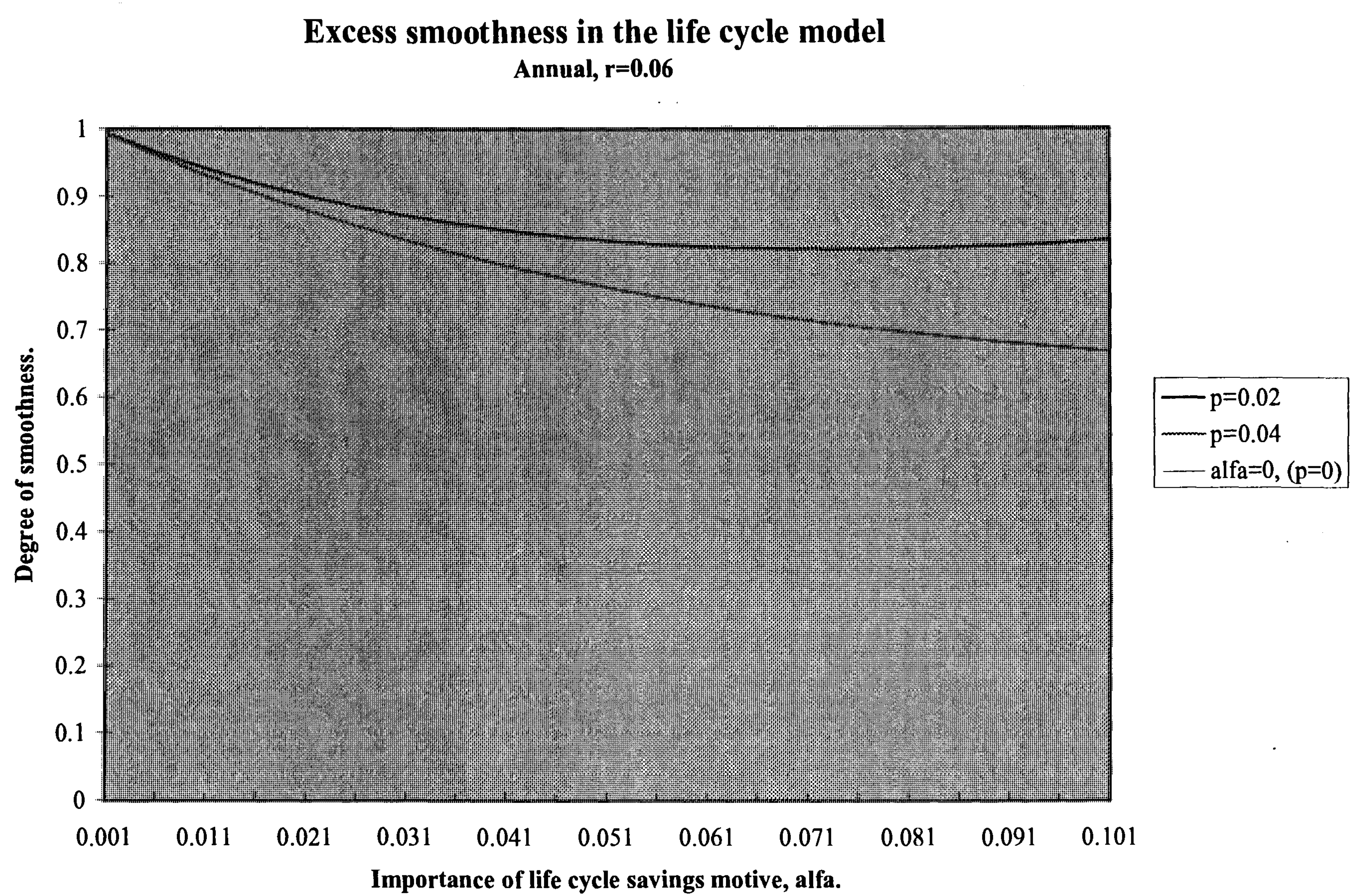
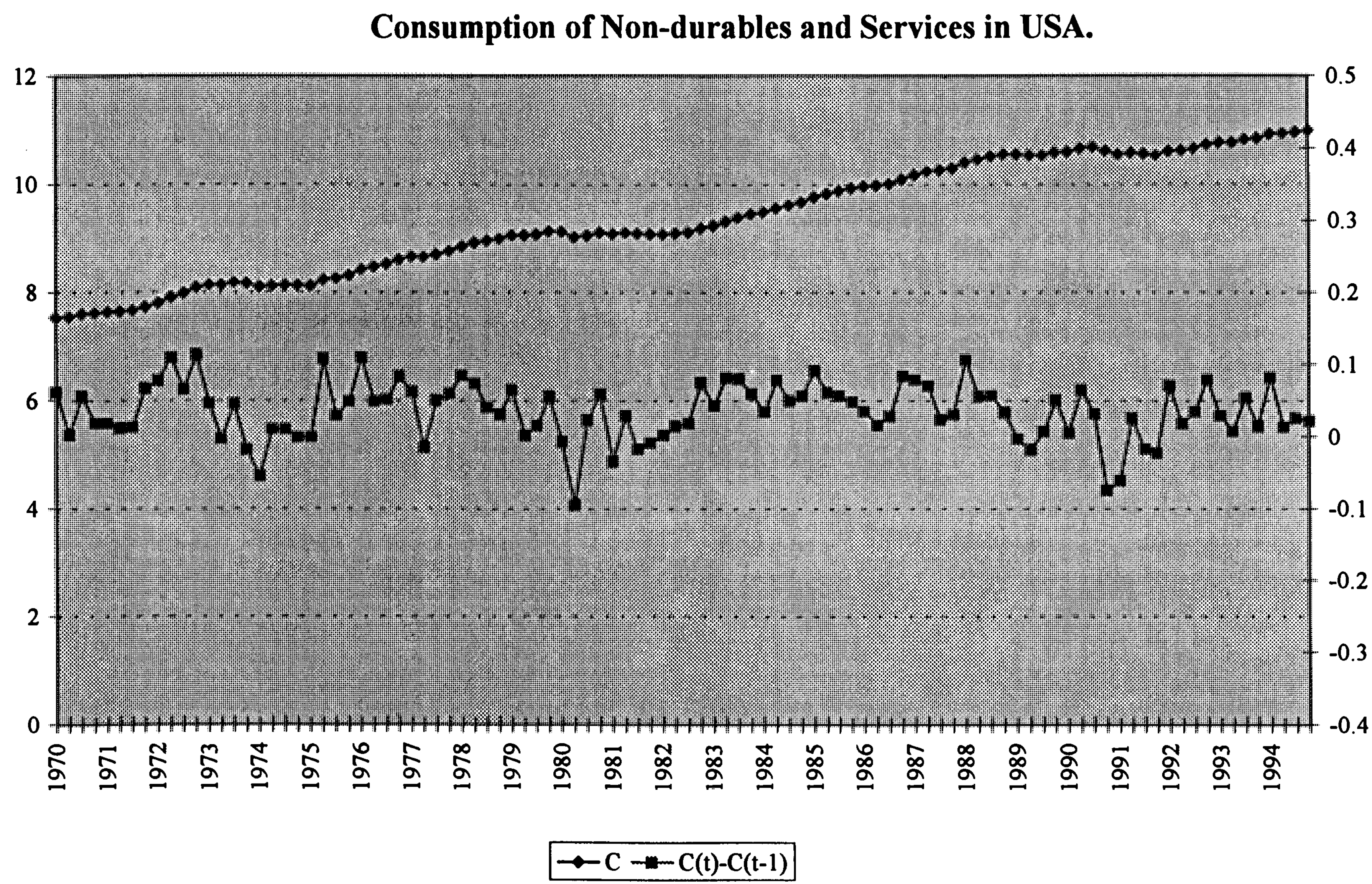

Excess smoothness in the life cycle model
Annual, r=0.06
Degree of smoothness.
1
0.9
0.8
0.7
0.6
0.5
0.4
0.3
0.2
0.1
0
0.001
0.011
0.021
0.031
0.041
0.051
0.061
0.071
0.081
0.091
0.101
Importance of life cycle savings motive, alfa.
p=0.02
p=0.04
alfa=0, (p=0)

Figure 3.



Consumption of Non-durables and Services in USA.
12
10
8
6
4
2
0
0.5
0.4
0.3
0.2
0.1
0
-0.1
-0.2
-0.3
-0.4
1970
1971
1972
1973
1974
1975
1976
1977
1978
1979
1980
1981
1982
1983
1984
1985
1986
1987
1988
1989
1990
1991
1992
1993
1994
C
C(t)-C(t-1)

Figure 4.

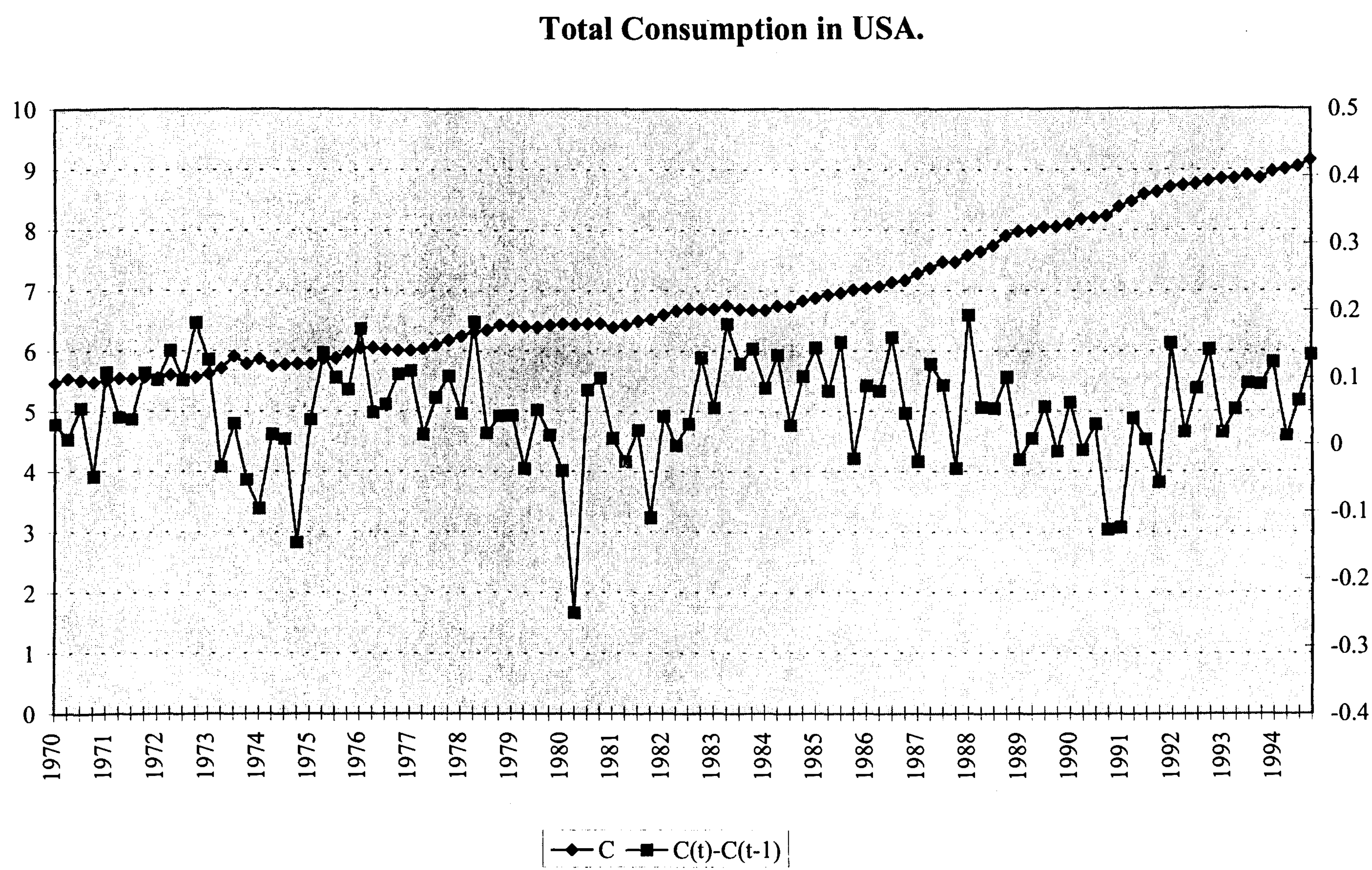

Total Consumption in USA.
10
9
8
7
6
5
4
3
2
1
0
0.5
0.4
0.3
0.2
0.1
0
-0.1
-0.2
-0.3
-0.4
1970
1971
1972
1973
1974
1975
1976
1977
1978
1979
1980
1981
1982
1983
1984
1985
1986
1987
1988
1989
1990
1991
1992
1993
1994
C
C(t)-C(t-1)

Figure 5.

Disposable Labour Income in USA.

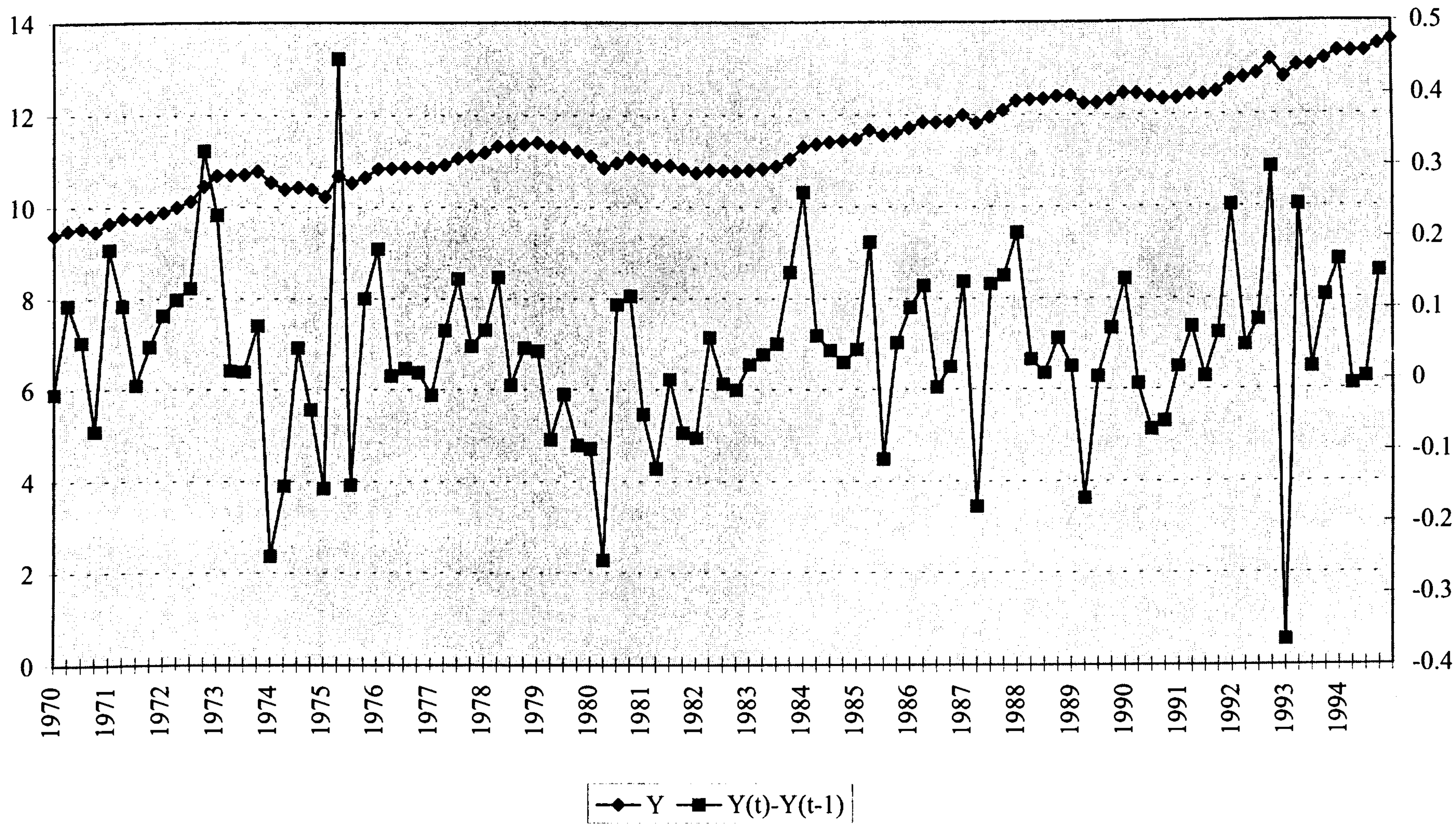


Figure 6.

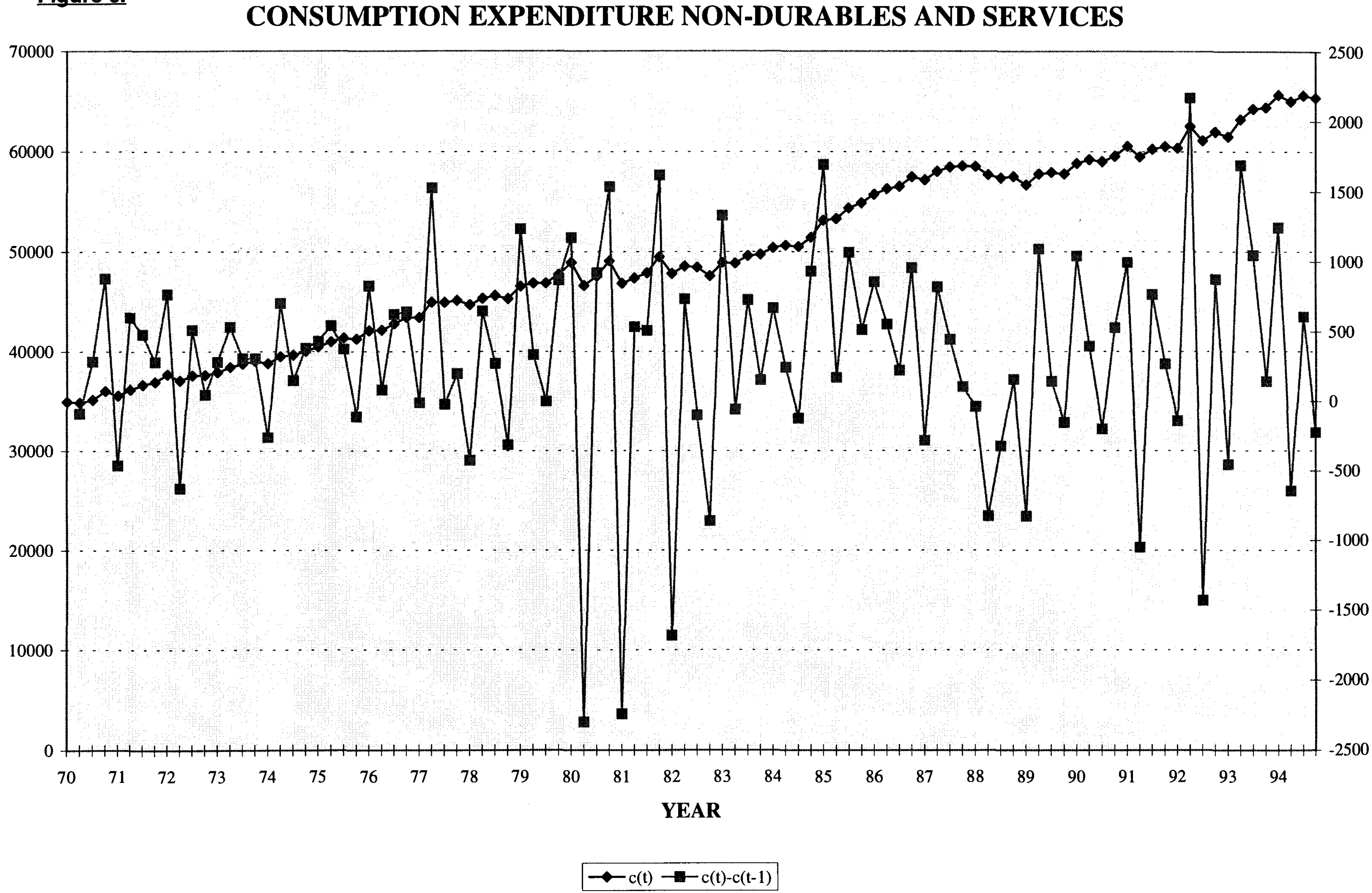
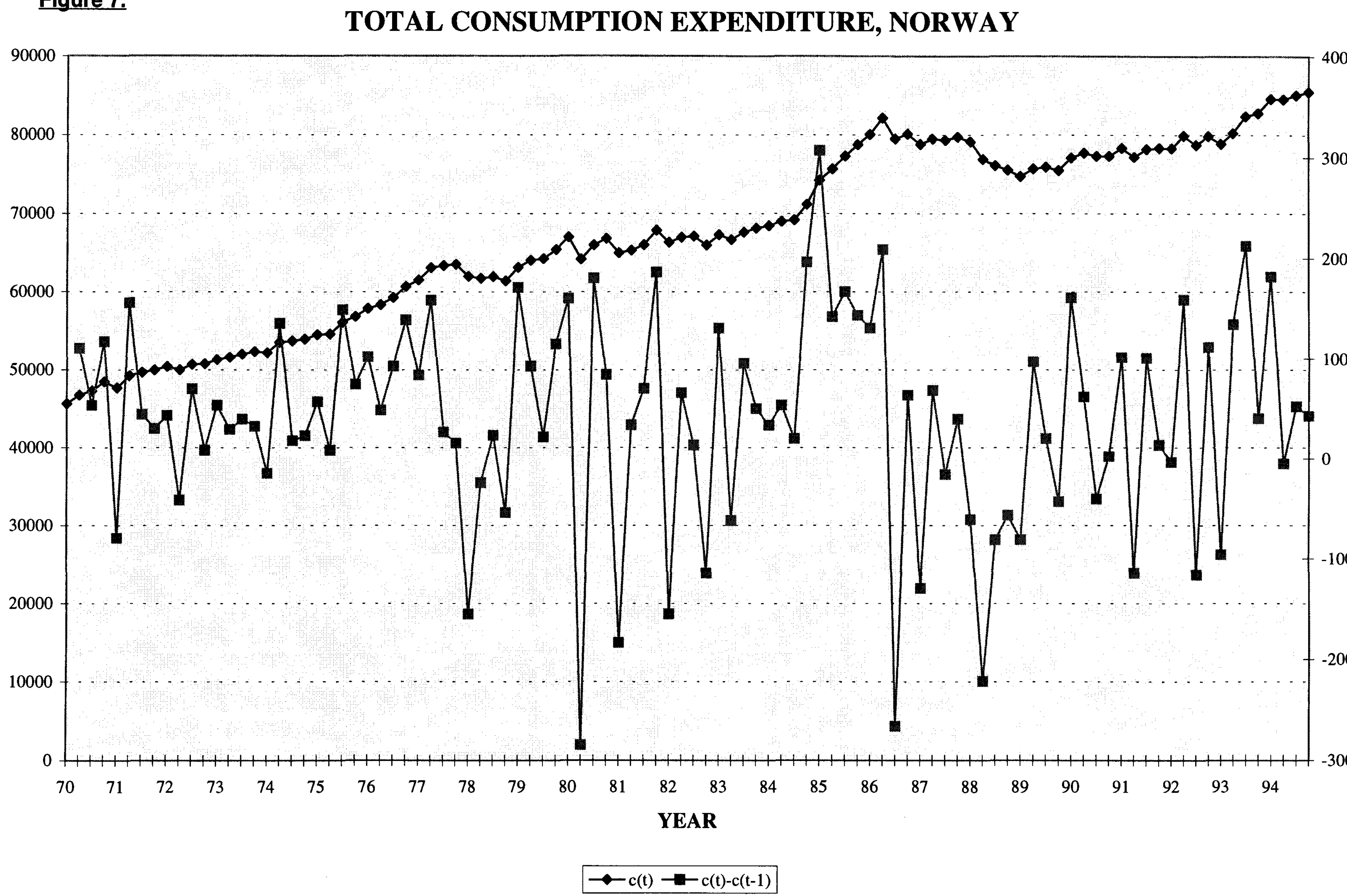

CONSUMPTION EXPENDITURE NON-DURABLES AND SERVICES
70000
60000
50000
40000
30000
20000
10000
0
2500
2000
1500
1000
500
0
-500
-1000
-1500
-2000
-2500
70
71
72
73
74
75
76
77
78
79
80
81
82
83
84
85
86
87
88
89
90
91
92
93
94
YEAR
c(t)
c(t)-c(t-1)

Figure 7.



TOTAL CONSUMPTION EXPENDITURE, NORWAY
90000
80000
70000
60000
50000
40000
30000
20000
10000
0
4000
3000
2000
1000
0
-1000
-2000
-3000
70
71
72
73
74
75
76
77
78
79
80
81
82
83
84
85
86
87
88
89
90
91
92
93
94
YEAR
c(t)
c(t)-c(t-1)

Figure 8.

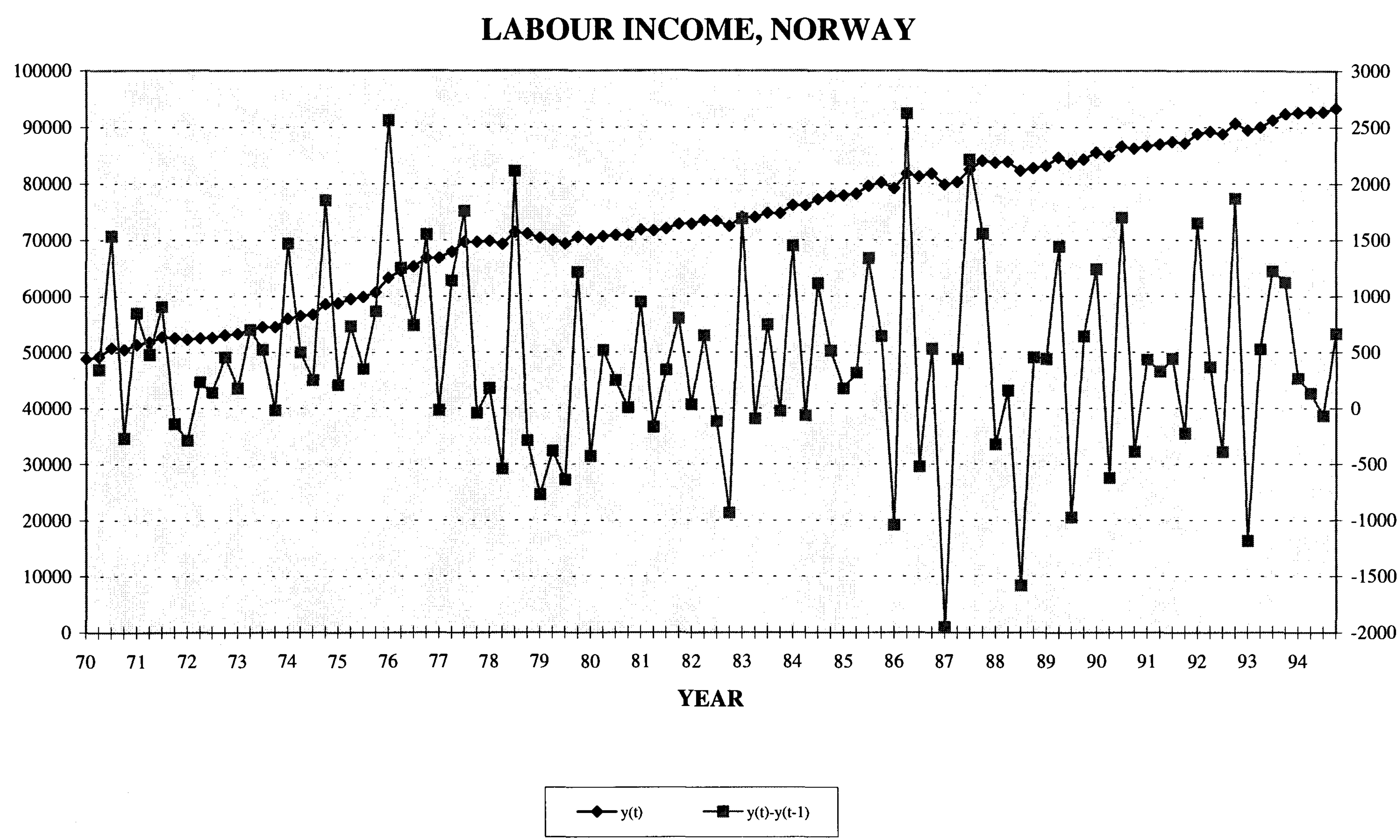

LABOUR INCOME, NORWAY
100000
90000
80000
70000
60000
50000
40000
30000
20000
10000
0
3000
2500
2000
1500
1000
500
0
-500
-1000
-1500
-2000
70 71 72 73 74 75 76 77 78 79 80 81 82 83 84 85 86 87 88 89 90 91 92 93 94
YEAR
y(t)
y(t)-y(t-1)

Figure 9.

AR(1) estimator for US Quarterly Disposable Labor Income.

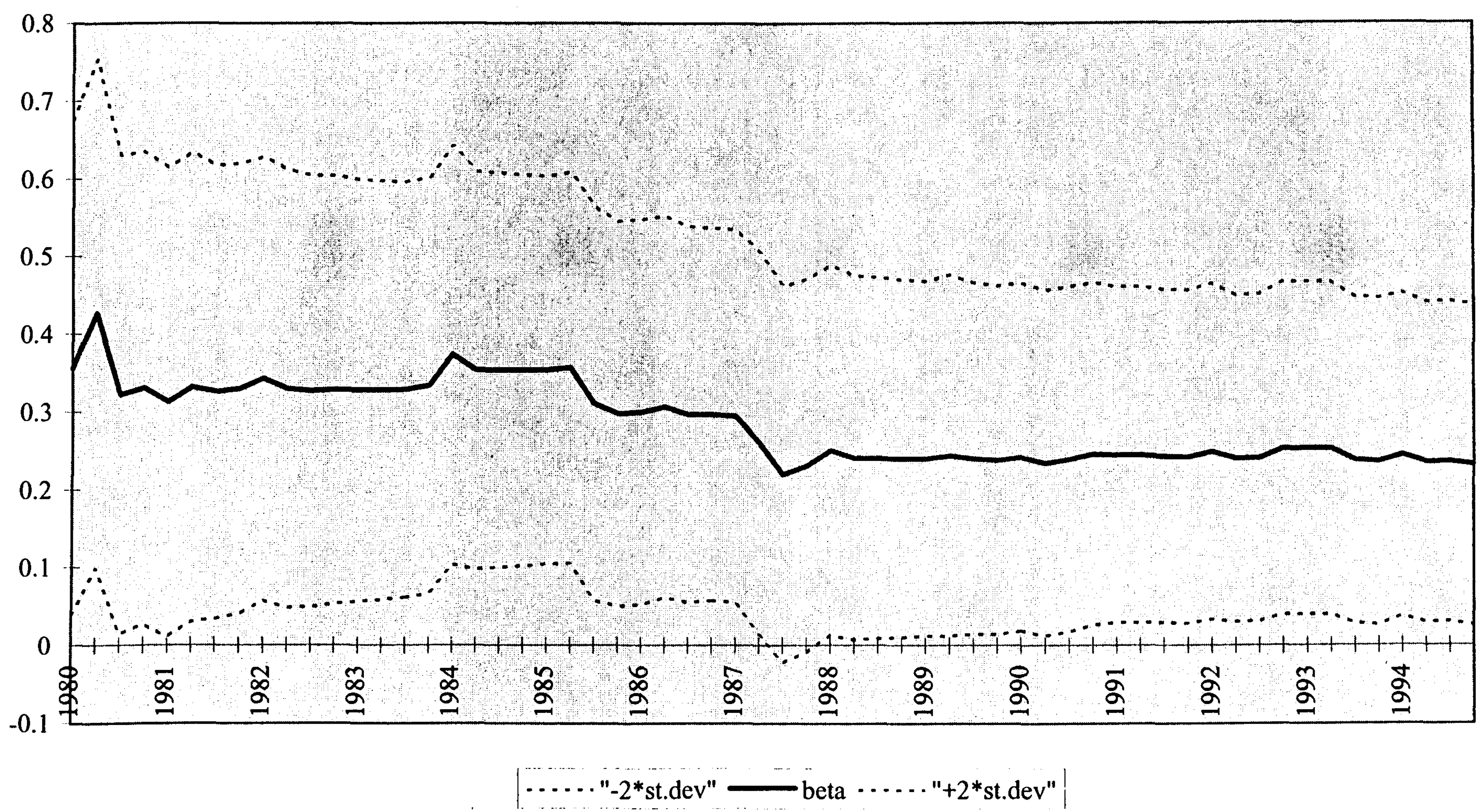


Figure 10.

AR(1) estimator for US Annual Disposable Labor Income

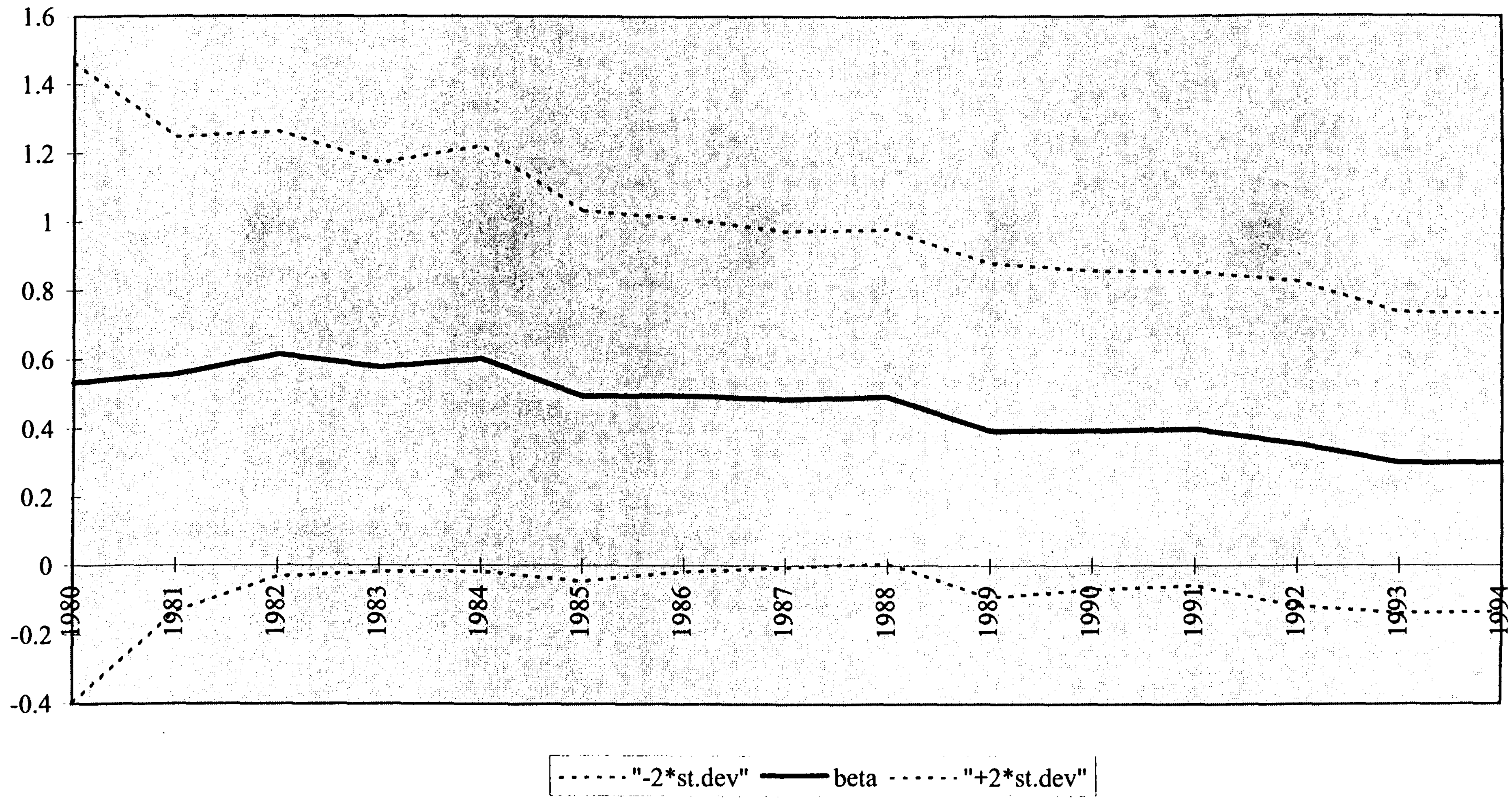


Figure 11.

AR(1) estimat for Norwegian Quartely Disposable Labor Income.

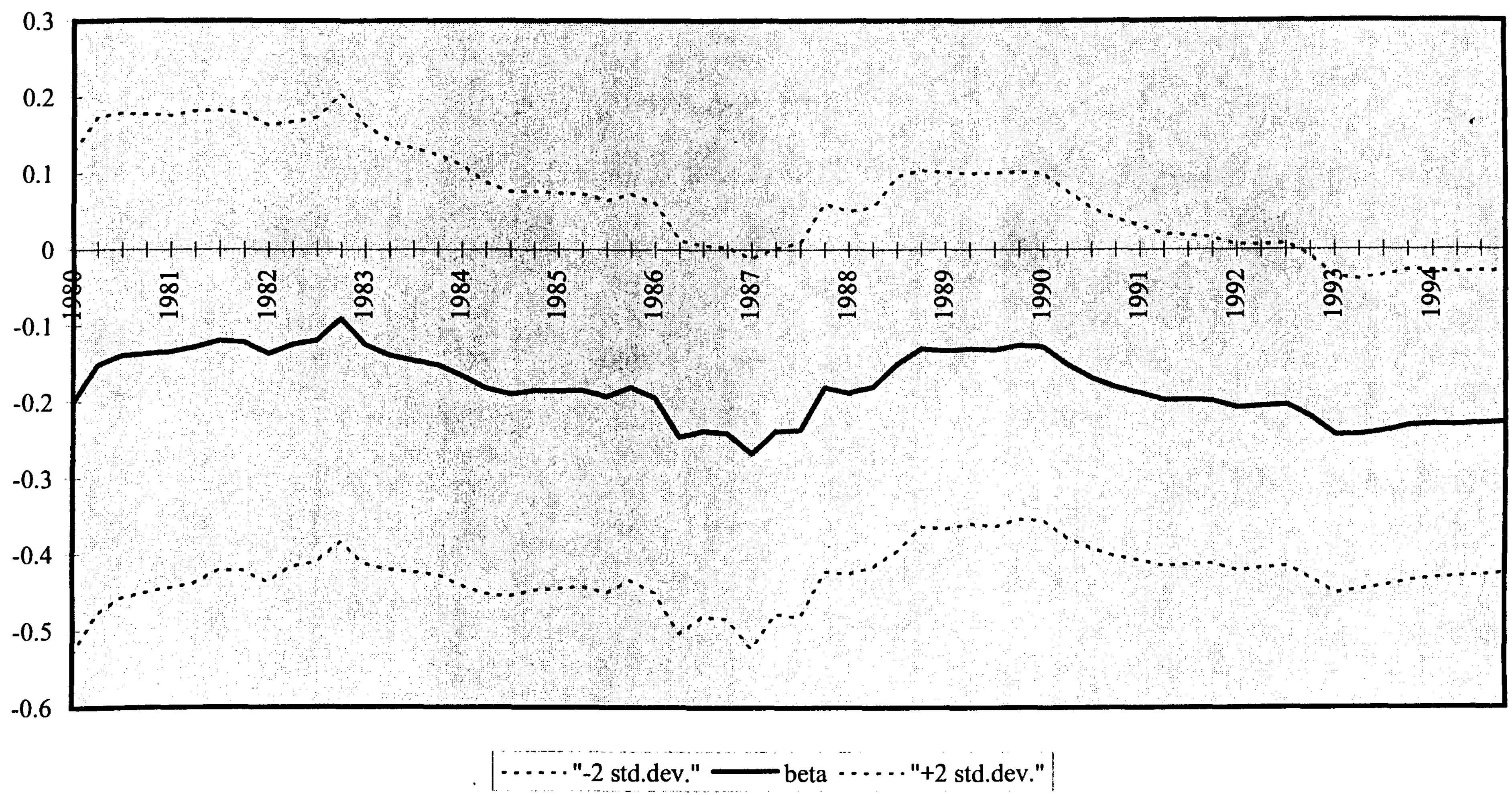


Figure 12.

AR(1) estimat Norwegian annual disposable Labour Income.

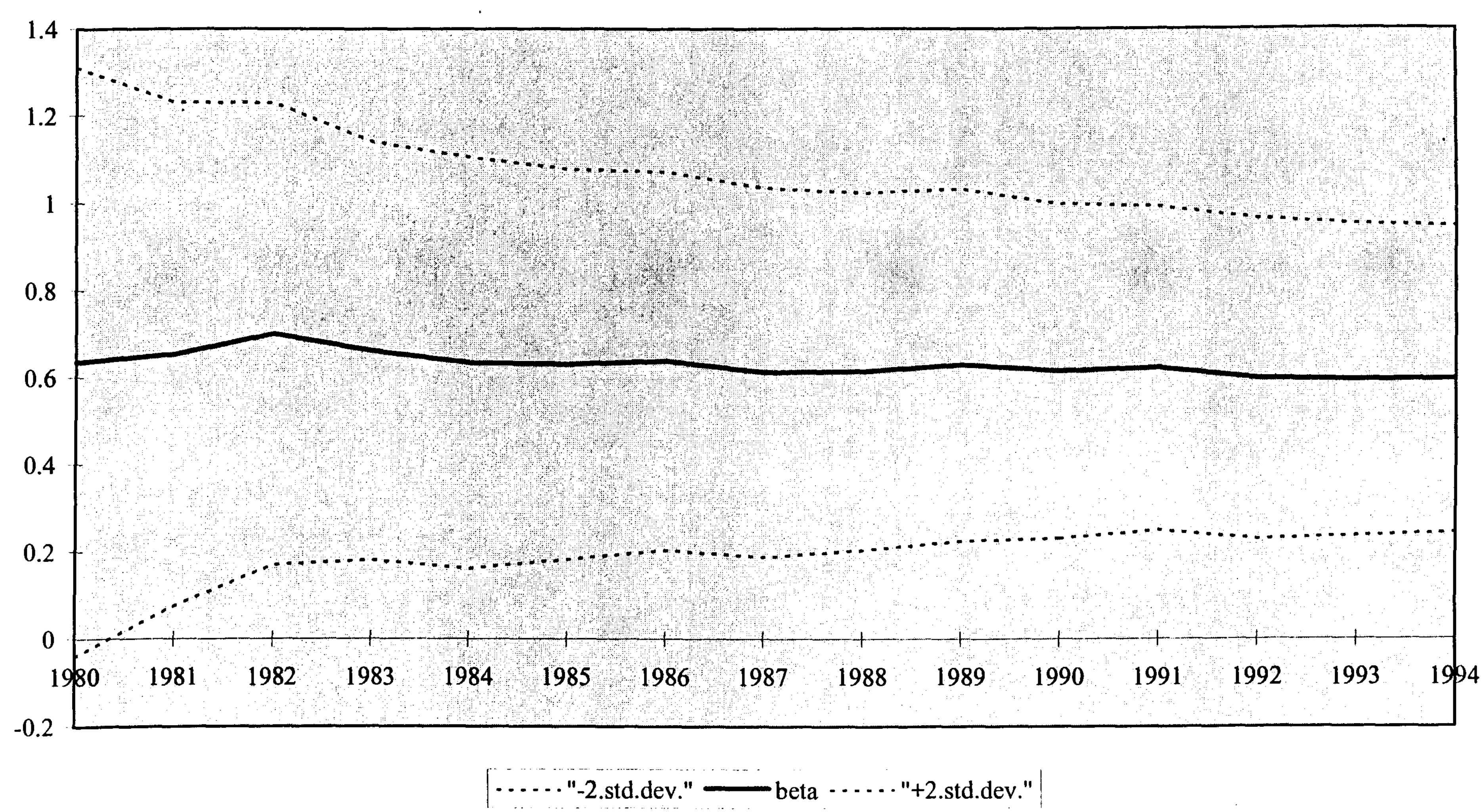


Figure 13.

AR(1)-estimator for quarterly US consumer expendityre, (non-durables and services).

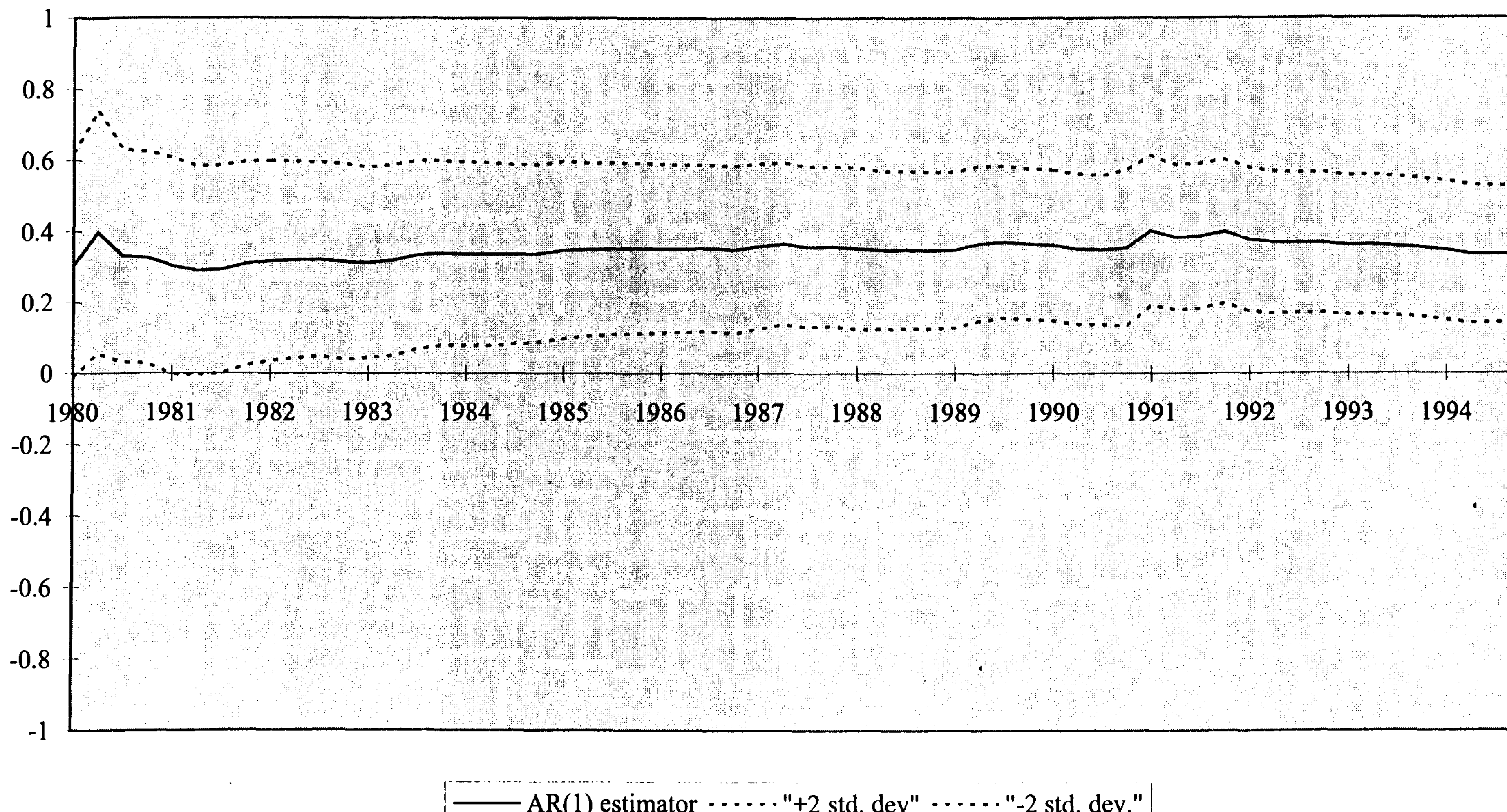
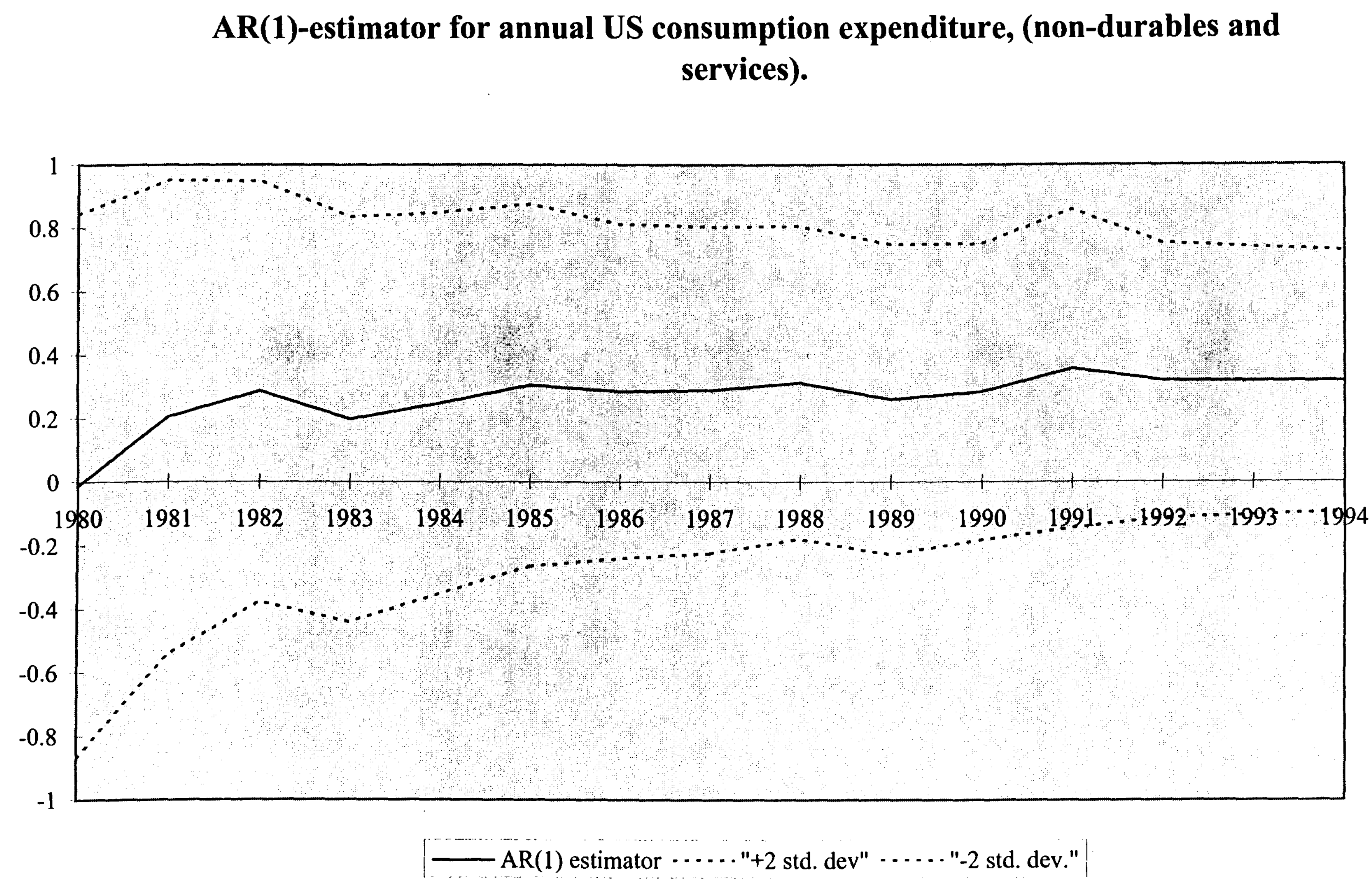


Figure 14.



AR(1)-estimator for annual US consumption expenditure, (non-durables and services).
1
0.8
0.6
0.4
0.2
0
-0.2
-0.4
-0.6
-0.8
-1
1980
1981
1982
1983
1984
1985
1986
1987
1988
1989
1990
1991
1992
1993
1994
AR(1) estimator
"+2 std. dev"
"-2 std. dev."

Figure 15.

AR(1)-estimator for quartely Norwegian consumer expendture, (non-durables and services)

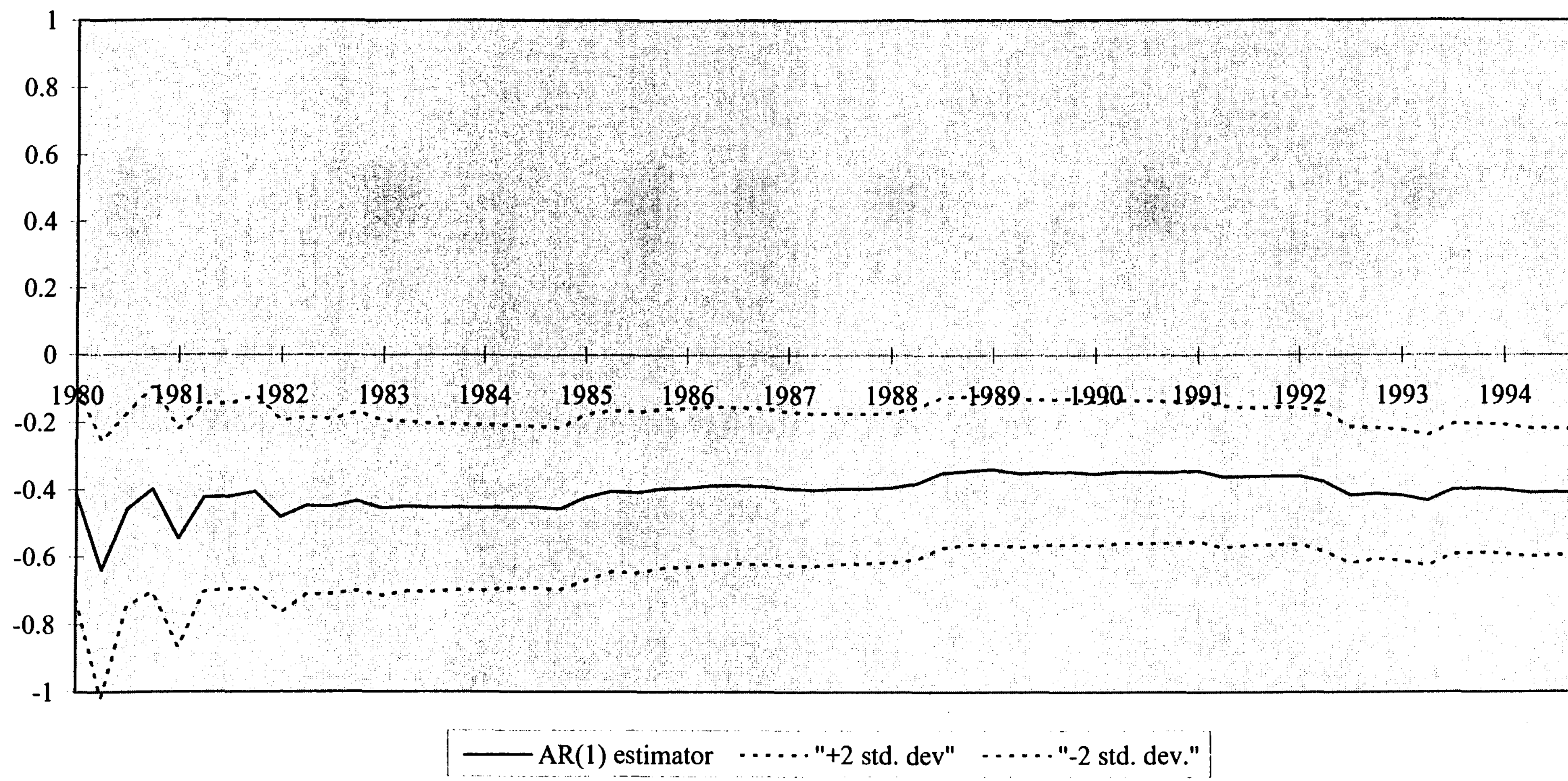


Figure 16.

AR(1)-estimator for annual Norwegian consumer expenditure,(non-durables and services)

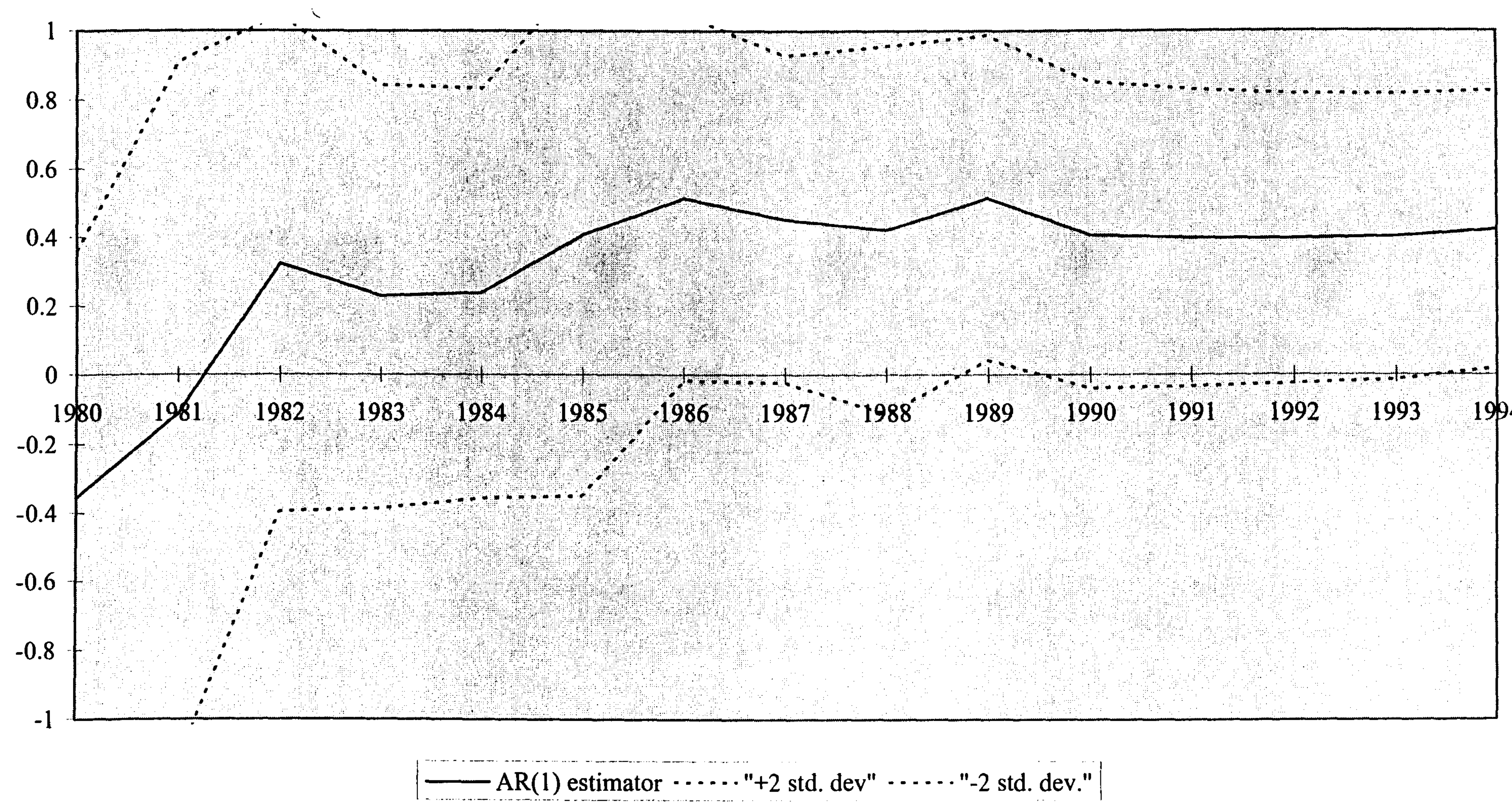


Figure17.

VOLATILITY IN PERMANENT INCOME MEASURES, US QUARTERLY DATA

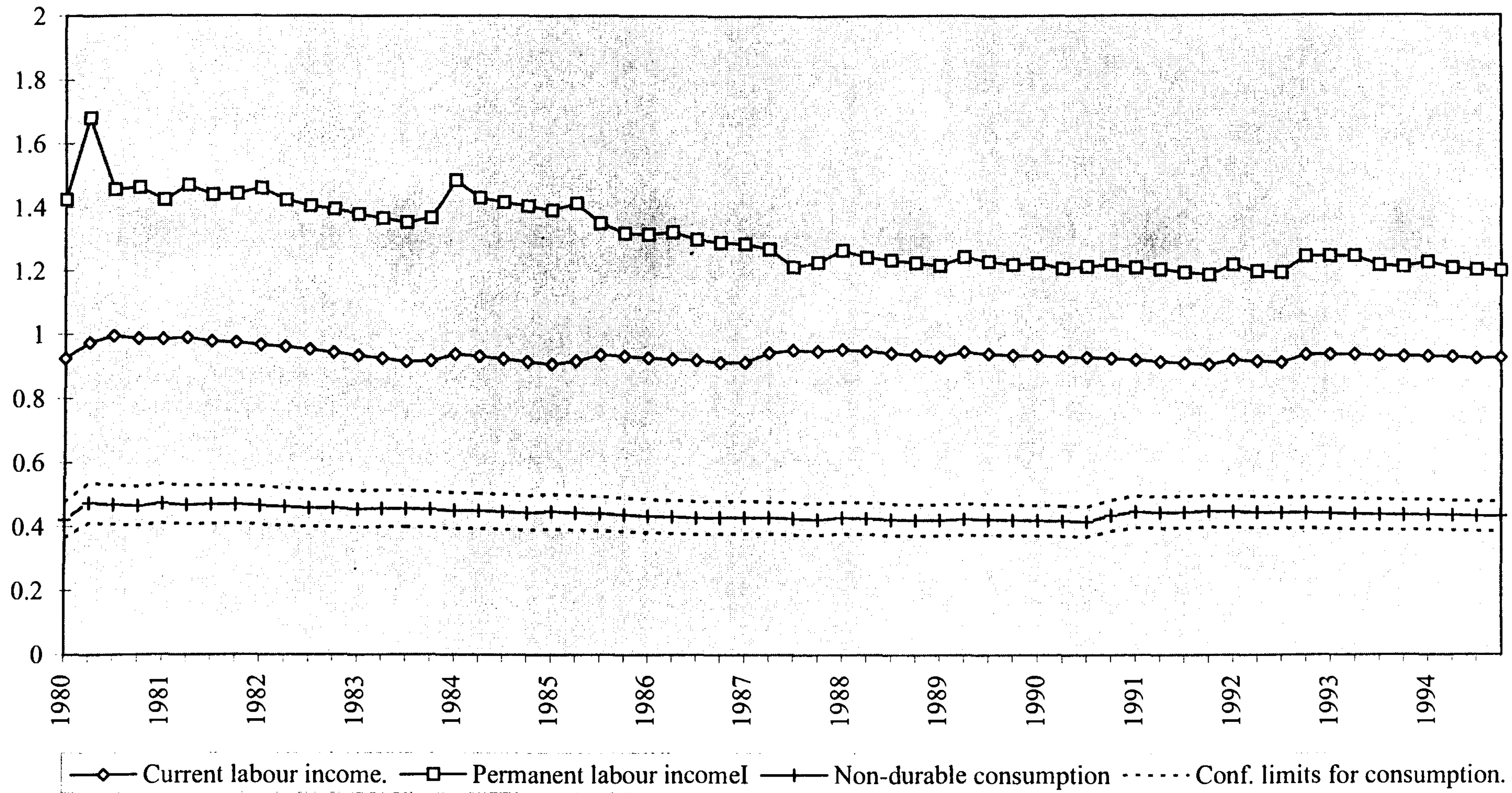


Figure 18.

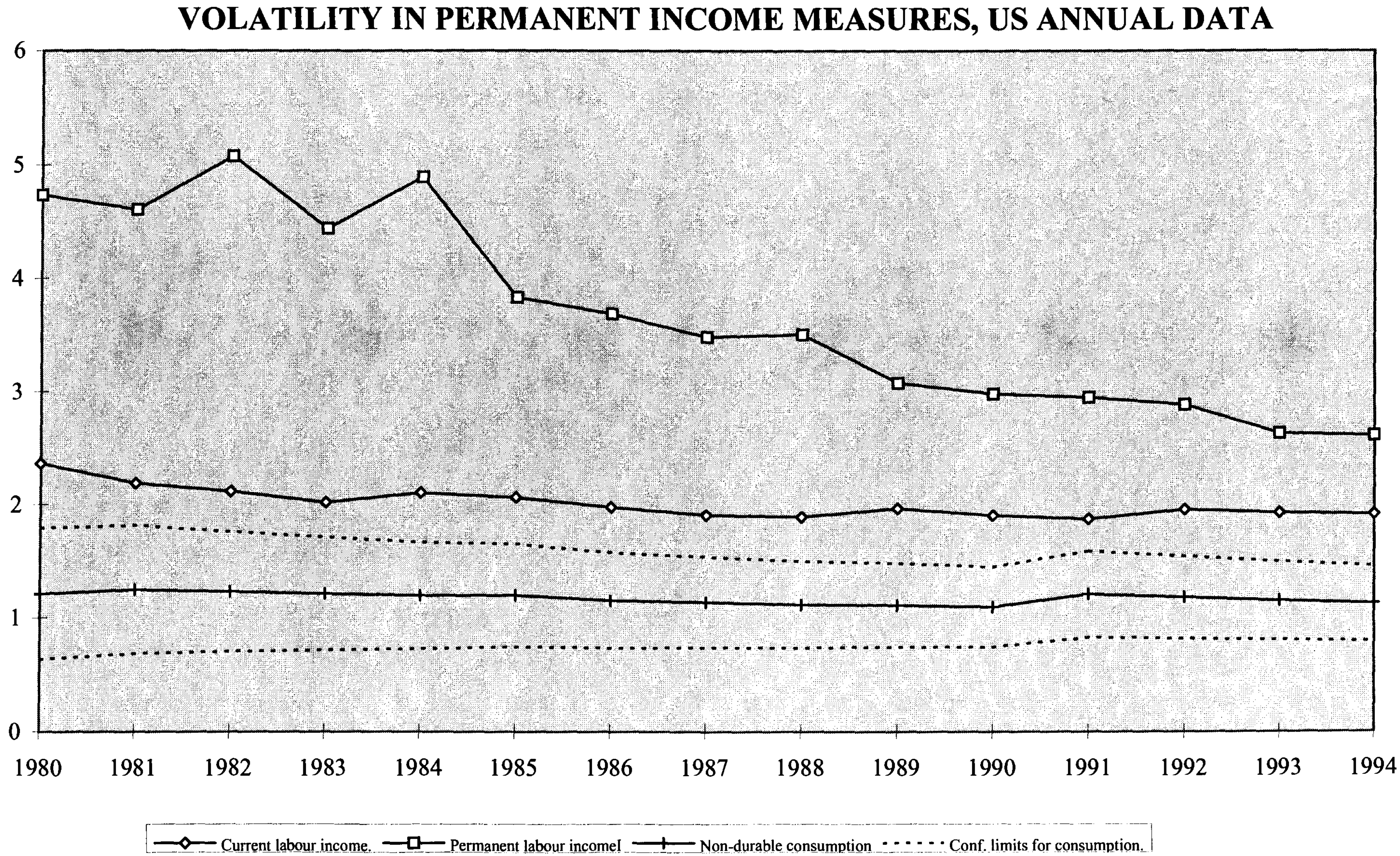
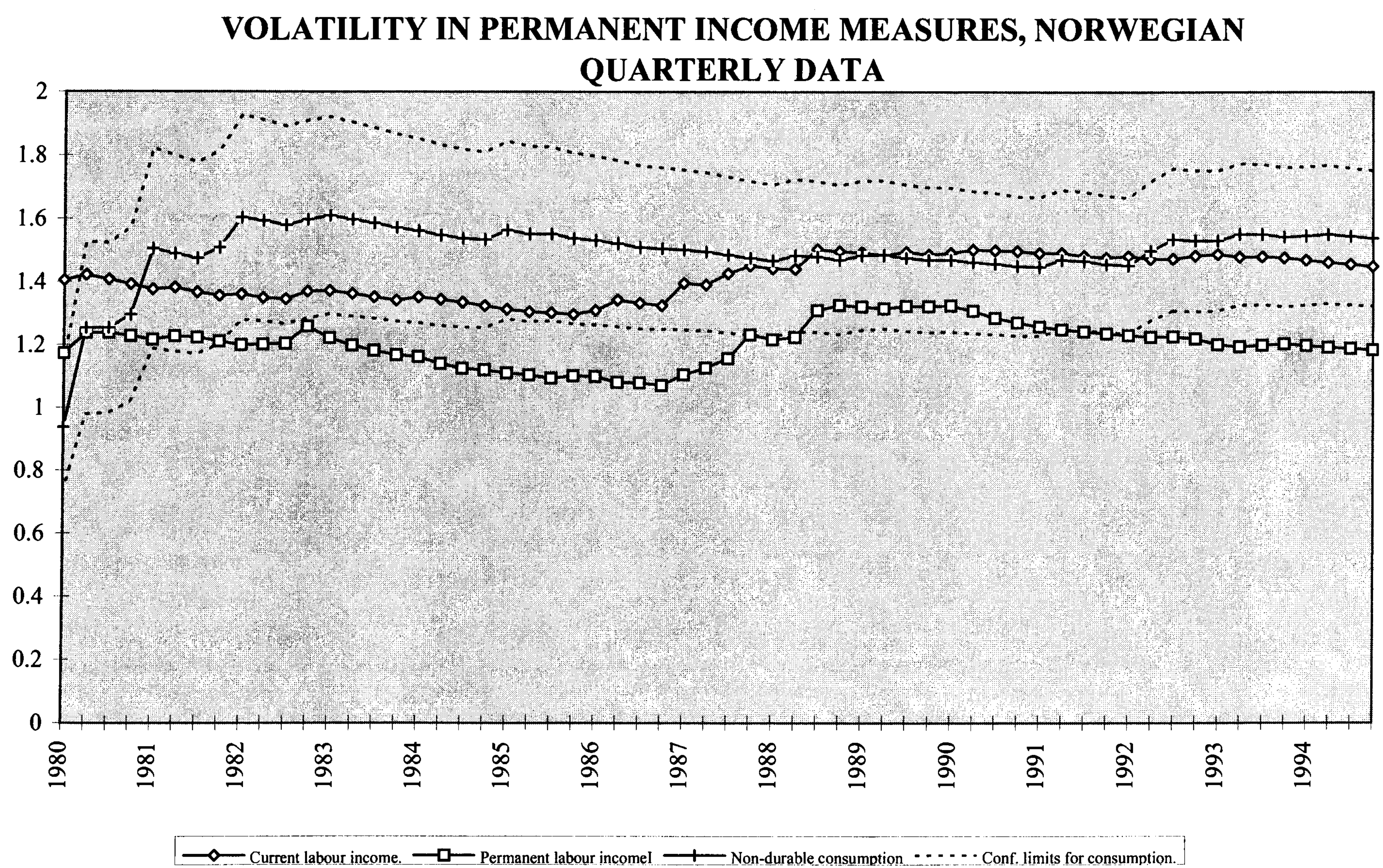

VOLATILITY IN PERMANENT INCOME MEASURES, US ANNUAL DATA
6
5
4
3
2
1
0
1980
1981
1982
1983
1984
1985
1986
1987
1988
1989
1990
1991
1992
1993
1994
Current labour income.
Permanent labour incomel
Non-durable consumption
Conf. limits for consumption.

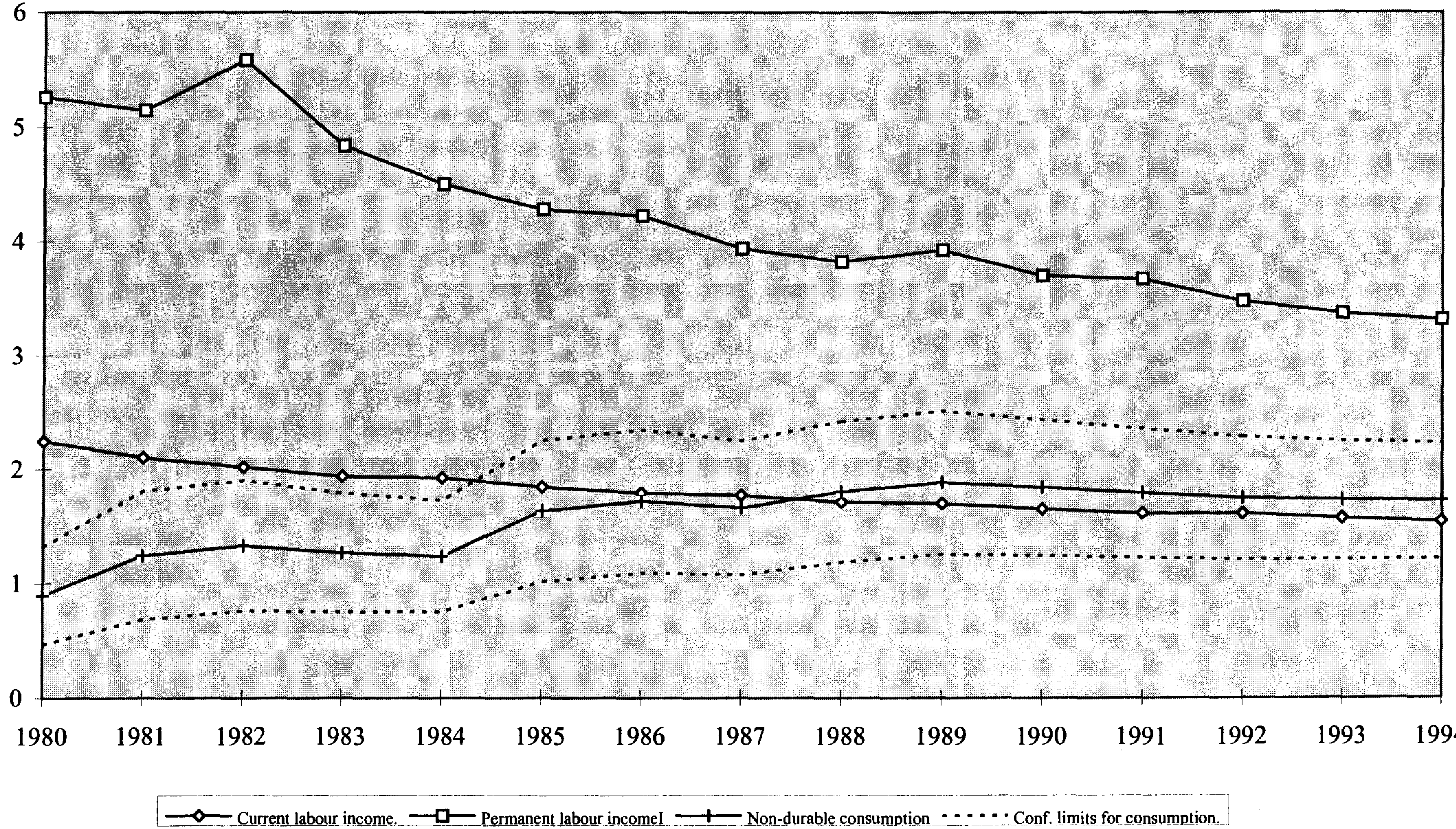
Figure 19.



VOLATILITY IN PERMANENT INCOME MEASURES, NORWEGIAN QUARTERLY DATA
2
1.8
1.6
1.4
1.2
1
0.8
0.6
0.4
0.2
0
1980
1981
1982
1983
1984
1985
1986
1987
1988
1989
1990
1991
1992
1993
1994
Current labour income.
Permanent labour incomel
Non-durable consumption
Conf. limits for consumption.

Figure 20.

VOLATILITY IN PERMANENT INCOME MEASURES, NORWEGIAN ANNUAL DATA



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ESSAY IV.

SEASONAL IMPLICATIONS OF THE PERMANENT-INCOME HYPOTHESIS.

A comparison of Norwegian and Swedish data for consumption and retail sales.

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1. Introduction

The conventional way of testing the permanent-income or life-cycle hypothesis is to derive an Euler equation or other dynamic relation, estimate it on aggregate or household data, and test the overidentifying restrictions on this equation. Another approach is to single out a more specific implication that can be tested in reduced form, perhaps on more easily available data. Shefrin and Thaler (1988) present some examples of this approach, such as the findings of Cagan (1965) that people with pension plans save more rather than less. More recent examples are the studies by Wilcox (1987, 1989) and Shapiro and Slemrod (1994) of the effects of income tax refunds, announcements of changes in social-security benefits, and changes in the timing of tax withholding on the timing of consumption expenditure in the United States. According to the permanent-income hypothesis, the level and timing of consumption spending is affected only by the expected level and variance of income, not by the timing with which it arrives. Yet, all of these studies find that the timing of consumer expenditure in fact is affected by the timing of disposable income.

This paper reports the results of a similar reduced-form examination, namely of the effects of seasonal factors in disposable income on the seasonal pattern of consumption expenditure and retail sales. Paxon (1993) is similar in spirit to our study. She considers how seasonality in agrarian income influence on the seasonality in household consumption expenditure in Thailand. She finds that rural and farm households have a similar seasonal consumption pattern to other households. We exploit the Norwegian tax withholding system. In Norway, income taxes, including social insurance premia, are withheld at an even rate for ten of the twelve months of the year. However, tax withholding is suspended during one of the summer months (June or July, most often June) and done at half the normal rate in December. It may be noted that monthly salaries in Norway conventionally are paid around the middle of the month, so that the extra disposable income in December is received in time before Christmas. For the same reason, those receiving paycheques without tax withholding in June receive the extra money early enough for spending by the end of the second quarter. However, the official reason for suspending tax withholding during one of the summer months is to enable people to go on summer vacations. With July and August being the traditional summer vacation months, it thus should not be surprising if the extra money isn't spent until then.

Of course, the extra disposable income that Norwegian this way occasionally receive is neither permanent nor surprising, and it is offset by higher withholding rates during the rest of the year. Thus, for rational consumers following the permanent-income hypothesis, there should be no reason for the seasonal pattern of spending to be affected by this institutional fact, Christmas and summer vacation notwithstanding. Thus, if the seasonal pattern nevertheless is affected, that would be a contradiction of the permanent-income hypothesis.

Designing a test for any such effect is not trivial, however. Although seasonally unadjusted quarterly data for consumption and household disposable income are available for Norway, the data for disposable income rely on an imputation of tax liability rather than actual withholding. Data for the government's tax revenue are available; but these data tell only when the withheld taxes were paid by the employers, not when the withholding took place. Moreover, even if we had had ideal data for disposable income, we would have had serious difficulties identifying the seasonal movements that are due to the variation in tax withholding as opposed to the «deep structure» seasonal factors in preferences and technology.

If the Norwegian withholding system had been carried out as a proper experiment, data would have been available before and after the change, and there would have been a control group that did not receive the treatment. Of course, no such controlled experiment has taken place. Furthermore, neither the quarterly consumption data nor the monthly retail-sales data go back to the time before this rule was implemented. However, a control group of sorts is available: the withholding of income taxes in Sweden is carried out at an even rate throughout the entire year. A control group needs to be similar to the treated group along the relevant dimensions. In our case, it needs to be true that the seasonal pattern of consumer spending would have been equal for the two countries in the absence of the differences in tax withholding. Under this maintained hypothesis, the effects of the seasonal pattern in tax withholding can be estimated by comparing the seasonal patterns of consumption spending in Norway and Sweden.

While this assumption may not be exactly true, there seem good reasons why it should hold as a close approximation. The two countries lie next to each other at approximately the same latitude and thus essentially share the same climate. They also basically share the same history and culture; even the languages are almost identical. It thus seems quite reasonable to assume that their seasonal factors in technology and preferences are similar. At least, we would be hard put identifying any pair of countries for which this similarity should hold more closely. On the basis of these considerations, we treat the differences in tax withholding between Norway and Sweden as an approximate natural experiment and derive the results we present from a comparison of the seasonal patterns in consumer spending and retail sales of Norway and Sweden. Although we would have liked to go further by adding neighbouring Denmark and Finland as additional control groups, we were prevented from doing so by data limitations.

Our results suggest significant deviations from the permanent-income hypothesis. The deviation is clearest in the quarterly consumption data, where the differences in seasonal patterns correspond quite closely to the differences in tax withholding. The implied marginal propensity to consume (MPC) out of extraordinary tax withholdings is close to unity. The pattern is somewhat less pronounced for retail sales, but it remains true for these data as well that Norwegian spending is concentrated more heavily during the third and fourth quarters compared to Sweden.

Whenever the permanent-income hypothesis fails it is important to consider the reasons for the failure, in other words the alternative hypothesis. We can imagine three reasons why the seasonal pattern of Norwegian spending should be correlated with the tax withholding pattern. The first two coincide with the two alternative hypotheses advanced by Flavin (1986), namely liquidity constraints and myopic behaviour. Liquidity constraints have been analysed rather extensively in the literature, for example by Zeldes (1989) and more recently by Deaton (1991). It is not quite convincing as an alternative in the present case, however, because it is not necessary for a consumer to borrow in order to distribute consumption in a desirable fashion over the year. A Norwegian household wishing to consume more in the first quarter could easily do so by saving from the suspended tax withholdings in the preceding summer and Christmas season[1].

[1] The exception to this rule would be young households entering adulthood in August or January with no financial assets. However, these households receive a very small portion of aggregate income and should not give significant results on aggregate consumption.

The other alternative is that people act myopically in the sense that they constrain themselves to spending extraordinary income when it arrives. This type of behaviour need not be irrational in a wider perspective, but may be consistent with the behavioural life-cycle hypothesis advanced by Shefrin and Thaler (op. cit.). Rather than devoting time and energy to optimisation (an activity they might not be good at anyway), people may find it convenient and even efficient to organise their decisions around «mental accounts» as a guide to which money is to be spent on what and when. With this kind of behaviour it should not be surprising to see Norwegian consumers organising their vacation and holiday spending around the «allowances» allotted to them by government policy. The seasonal spending pattern then would not be solely the product of innate preferences, but influenced by the manipulation by a (perhaps) benevolent government.

The third alternative is that the permanent-income hypothesis actually holds, but that the maintained hypothesis of identical seasonal factors in preferences and technology fails. Indeed, it may be argued that Norwegian consumers have decided to facilitate their peculiar seasonal preferences via a saving program organised by the government, which saves households the transactions cost of organising their own private saving. Although we find this view intriguing, we are reluctant to accept it, at least as a complete explanation. For one thing, the transactions cost of private saving appears to be quite small unless one includes the psychic cost of imposing the discipline of saving on oneself; and in that case we would be back with the behavioural life-cycle hypothesis. For another, if there is a transactions cost of private saving, it should not be significantly different for Norwegians than for Swedes, so we would need another explanation for why Norwegian consumers chose this particular scheme. Third, the seasonal peak is not the summer, but the Christmas season, and for both countries. Thus, if Norwegian consumers wanted to save through the tax withholding system, it would make much more sense to suspend withholding completely in December than in the summer.

The organisation of the rest of the paper is as follows. Because the data collection methods for consumption and retail sales are not identical in Norway and Sweden, the next section starts by taking a careful look at these methods. We also look at the methods for converting retail sales and other source data into quarterly data for consumer spending. We are particularly interested in ways in which the data might be influenced by measurement errors that might potentially differ systematically between the two countries. Section 3 presents a model of seasonality and formulates the hypotheses tested in this paper. Sections 4 and 5 present the estimation and testing results for consumer spending and retail sales, respectively, whereas Section 6 seeks to reconcile the seasonal factors in consumer spending and retail sales for the two countries. Section 7 closes the paper by offering some additional discussion and conclusions.

2. Data Construction for Retail Sales and Consumption

For Norway as well as Sweden, the quarterly consumption data are constructed on the basis of monthly data for retail sales as well as other primary sources. Retail sales are classified according to types of stores, such as food stores, hardware stores, department stores, etc. The data are constructed first as sub-indices for each category of stores and then aggregated to a total.

The construction of the retail indices of both countries start out with a base year, for which monthly level data are constructed from direct observations of a sample of stores. For subsequent years, however, the levels are not recorded directly. In the case of Norway, the stores in the sample are asked to report the change in sales from the previous month, except for January, for which they report the change from January of the preceding year.

For a model of this procedure, let x_{kt} denote the true value of the level of retail sales for month k in year t, and y_{kt} the recorded level from the method just described. Let the base year be 0 and suppose that the level in month k of the base year is measured with an error ε_k, so that $y_{k0} = x_{k0} + \varepsilon_k$. Suppose further that the change from January of year t-1 to year t is observed with an error denoted E_t and that the monthly change between the months k - 1 and k in year t is observed with an error denoted $e_{kt}, t > 0, k > 1$. Then, for any year following the base year, the recorded levels for the respective months may be written as

$$
(1) \qquad \begin{aligned} y_{1t} &= x_{1t} + \varepsilon_1 + \sum_{s=1}^{t} E_s \\ y_{kt} &= x_{ky} + \varepsilon_1 + \sum_{s=1}^{t} E_s + \sum_{j=2}^{k} e_{jt}, \qquad k = 2,\ldots,12. \end{aligned}
$$

Thus, measurement errors accumulate over time within each year[2]. However, the measurement errors that accumulate over more than a year are not tied to any particular month, so that the observation errors in the seasonal factors of one year tend to disappear the following year. There is no tendency for errors in the seasonal factors to persist or accumulate over time. Note that the measurement errors for the base-year levels also do not persist in a way that distorts the seasonal patterns.

Measurement errors with this pattern should not contribute a seasonal unit root to the observed series. Thus, rejection of a seasonal unit root is consistent with the description of the measurement errors in (1). On the other hand, a seasonal unit root may very well be present in the true series. Thus, failure to reject a seasonal unit root carries no information about the measurement errors per se.

The measurement method for Sweden is somewhat different. Here, the level of the retail-sales index for any month is recorded as the observed level of the same month the year before, plus the observed 12-month change. Suppose again that the true base-year levels x_{k0} are measured with observation errors ε_k. Suppose further that, for month k

[2] There might be an automatic correction of the error accumulation if the same stores were surveyed each month and the store computed the change it reports from the level it observes for the current and the preceding month. However, in practice, the sample changes over time so that some accumulation of errors must be expected.

of year t, the changes from the same month the year before are measured with errors E_{kt}. Then, the observed level for month k of year t may be written as

$$(2) \qquad y_{kt} = x_{kt} + \varepsilon_k + \sum_{s=1}^{t} E_{ks} .$$

This formula reveals two potential differences from the Norwegian case. First, if the seasonal factors in the base year are measured incorrectly, these errors persist indefinitely, as indicated by the term ε_k. Second, if the 12-month growth rates are measured incorrectly, these errors accumulate as distortions of the seasonal pattern[3].

Under the assumption that the measurement errors for the 12-month growth rates are stationary, the measurement errors in the levels in formula (2) have a seasonal unit root. Thus, if the measurement errors in growth rates are nonnegligible the observed series should have a seasonal unit root as well, except in the unlikely case that the measurement errors and the true series are cointegrated on the annual frequency. Rejection of a seasonal unit root then may be taken as an indication that the accumulating measurement errors E_{kt} have negligible variance. As in the Norwegian case, nonrejection of a seasonal unit root has no particular consequence because the true series may very well contain such a root.

As seen in the next section, we will also study seasonality in the form of varying seasonal means, in other words, as constant coefficients of seasonal dummies. If the accumulating measurement errors in (2) are important, we would expect to observe different seasonal means for different subsamples. Small changes over time then may serve as another indication that the variance of these measurement errors are negligible[4].

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[3] Again, one might hope that the accumulated errors could be corrected by the reporting stores themselves if the same stores were surveyed each month. However, the sample changes over time for Sweden as well.
[4] Note that the measurement errors for the seasonal base-year levels also persist in the Swedish case, even though they do not give rise to a seasonal unit root. For the years following 1984 we were able to check

Most of the quarterly data for goods consumption are derived from the retail-sales indices. The main exceptions are household energy consumption including electricity and fuel for home heating and expenditure on motor vehicles. The retail-sales index contains no sub-index corresponding to household energy use; it does, however, contain a sub-index for the sales of automotive stores which resembles expenditures on automobiles. Data for the consumption of services are derived from other sources, in some cases as interpolations of annual data. From what we have been able to find out, no interpolation has been used for the components of goods consumption.

The conversion of retail-sales data into data for consumption expenditure on goods other than automobiles and household energy use goes essentially as follows. Because retail sales are classified by type of store and consumption by end use, conversion keys are needed to convert the sales of various kinds of stores to the respective enduses of consumption. For example, a certain percentage of the sales of food stores is considered food consumption. This procedure works fairly well for goods that make up most of the sales of a particular kind of store, such as food.

The procedure becomes more complicated when the correspondence is less one-to-one. One such example is hardware stores, where a fair share of the sales is considered intermediate inputs to production rather than consumption. Another example is department stores, which carry a wide selection of goods. For such stores, the composition of the selection may vary over the seasons; but such variations do not show up in the statistics because the conversion keys are kept constant across seasons. Fortunately, department stores were responsible for only a minor portion of total retail sales during our sampling period.

For Norway, the level data for retail sales are converted into level data for consumer spending for each quarter. For Sweden, the construction of the quarterly consumption

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for this possibility by comparing the old series, reported to use 1970 as its base year, with a newly constructed one, based on 1985. Although the sample for this comparison is short, we find it reassuring that the seasonal means of the two series appear to be quite similar.

data starts with a base year, for which quarterly consumption levels are constructed on the basis of the level data for the corresponding retail-sales series. From then on, level data for the respective quarters of subsequent years are constructed as the levels of the corresponding quarter the previous year plus the four-quarter change implied by the changes in the retail-sales data. For both countries, at the end of each calendar year, the quarterly levels are reconciled with the annual consumption data, which are based on larger samples. The method used in this reconciliation is a least-squares method which does not disturb the overall seasonal pattern.

We see no particular reason for this procedure to introduce important new errors in the seasonal patterns[5]. Nevertheless, the conversion process is complicated enough that the seasonal patterns of consumption and retail sales could differ, even when we compare variables that should correspond fairly closely to each other. Moreover, measurement errors in the seasonal patterns of retail sales are transplanted into those of consumption. This possibility should be of particular concern in the case of Sweden, where the seasonal measurement errors may accumulate over time.

The organisation of the retail-sales series is somewhat different for Norway and Sweden. For Norway, monthly data are collected for total retail sales as well as a number of subcategories, and all series are available in nominal as well as inflationadjusted terms. For Sweden, correspondingly comprehensive data were not available until the construction of a new retail-sales series, starting with 1984. Before that date, all we have available is a series referred to as «Actual Retail Sales,» a term used in Swedish statistics («egentlig detaljhandel») for retail sales excluding the sales of automotive dealers, wine and liquor stores (a government monopoly), and pharmacies. Moreover, this series is only available in nominal form. However, this older series has been updated to the present. Because we feel the period after 1984 is too short to allow meaningful statistical inference, we base our comparison of retail sales on this older series.

[5] Of course, an observation for a certain quarter in a given year may be measured with error as discussed in essay III. There seems to be no reason, however, for this error to contain a seasonal struture over time. Thus, errors in the overall seasonal pattern should cancel out and become small.

3. Modelling Seasonality

Consider a quarterly or monthly consumption series c_{it} where t now numbers the observations consecutively, and the first subscript i identifies the particular series. If the seasonality is deterministic, at least in part, and the series also has an exponential trend, it can be described by an equation of the form

$$(3) \qquad y_{it} \equiv \ln(c_{it}) = \lambda_i \cdot t + \sum_{k=1}^{m} a_{ik} \cdot Q_{kt} + e_{it}$$

Here, m is the number of observations per year, and Q_{kt} are seasonal dummies whose coefficients indicate deterministic seasonal factors (cf. e.g. Barsky and Miron, 1989). If the seasonality varies over time this variation will manifest itself as serial correlation in the residual e_{it}, of the order m, possibly with a seasonal unit root.

The series in which we are interested in this paper have non-seasonal parts that are likely candidates for having unit roots on the zero frequency, so that first-differencing is natural before estimation. Because the first differences of the seasonal dummies sum to zero, the first-difference form of (3) can be written as

$$(4) \qquad \Delta y_{it} = b_{i1} + \sum_{k=2}^{m} b_{ik} \Delta Q_{kt} + \Delta e_{it}$$

where $\quad b_{i1} = \lambda_i ,$

$$b_{ik} = a_{ik} - a_{i1}, \qquad k = 2,\ldots,m,$$

and the seasonal factors of the last m -1 seasons are measured relative to that of the first. Clearly, if the series has no deterministic seasonality the coefficients $b_{i2},\ldots b_{im}$ all equal zero.

A comparison of the seasonality of two series y_{it} and y_{jt} can be done by subtracting the equations corresponding to (4) for each of the series from each other and estimating

$$(5) \qquad \Delta y_{it} - \Delta y_{jt} = \beta_1 + \sum_{k=2}^{m} \beta_k \cdot \Delta Q_{kt} + u_t$$

where $\beta_k = b_{ik} - b_{jk}, \quad k = 1,\ldots.,m,$

$u_t = \Delta e_{it} - \Delta e_{jt}.$

The hypothesis that both series share the same seasonal pattern implies

$$(6) \qquad H_0\colon \quad \beta_2 = \ldots = \beta_m = 0.$$

In addition, the residual u_t also should have no seasonal unit root. Stationary fourth-order serial correlation is a further indication of seasonality, but an unreliable one because fourth-order serial correlation may be caused by a variety of other factors.

When the normalisation in (4) is used for quarterly data, the deterministic seasonal factors for the second through fourth quarters are identified as deviations from the first quarter. For monthly data, we would be looking at deviations from January. For better comparison between monthly and quarterly data it is useful to identify the seasonal factors for the monthly data as deviations from the first quarter as well, that is, as deviations from the mean of the first three months. To obtain this effect, we constrain the coefficients of the first three months to sum to zero, as opposed to requiring that the first one equal zero, which is implicit in (4). Then, (4) is replaced by

$$\Delta y_{it} = b_{t1} + b_{t2} \cdot (\Delta M_{2t} - \Delta M_{1t}) + b_{i3} \cdot (\Delta M_{3t} - \Delta M_{1t}) + \sum_{k=4}^{m} b_{ik} \cdot \Delta M_{kt} + \Delta e_{it} \tag{7}$$

where the seasonal dummies are denoted M_{kt}, to emphasise that they are monthly. Equation (5) is replaced by a similar expression. The two series y_{it} and y_{jt}, are natural logarithms of aggregate consumption expenditure, subcategories thereof, or of retail sales for the respective countries. Under the maintained hypothesis that consumers in both countries have the same seasonal preferences, the seasonality in consumption should be the same in both countries, so that the constraint H_0 (6) is implied by the permanent-income hypothesis.

Because seasonality may be stochastic as well as deterministic, an additional implication of the seasonal implication of the permanent-income hypothesis is that the difference between the Norwegian and the Swedish consumption (or retail-sales) growth rates should not contain a seasonal unit root. We carry out such tests in addition to the tests of the coefficients $\beta_2,\ldots\beta_m$.

Under the permanent-income hypothesis, the residuals e_{it} and e_{jt} each contain a unit root on the zero frequency, so that the seasonally adjusted consumption series is difference stationary in the sense of Nelson and Plosser (1982). In the case of non-durable consumption it also is uncorrelated with lagged information, except perhaps for its own lagged values as a reflection of «taste shocks». If durables are included the series still will be difference stationary, but its first differences will contain a moving-average term (Mankiw, 1982). Alternatively, if e_{it} does not contain a unit root, the seasonally adjusted series is trend stationary, and the permanent-income hypothesis fails.

Under the alternative hypothesis of myopia, the restriction in (6) fails because the seasonality of consumption follows the seasonality in income. If consumers are strictly myopic in the sense that their consumption follows disposable income completely over the seasons, the difference between the Norwegian and Swedish seasonal patterns should, on quarterly data, be about twice as large for the total of the second and third quarters as for the fourth because the reduction in Norwegian tax withholding is twice as large in June or July as in December. This strict—and perhaps naive version of the myopic alternative thus can be formulated as

(8) $$H_1: \quad b_2 + b_3 = 2 \cdot b_4 > 0 .$$

Under H_1, we may define $\mu = b_2 + b_3$ as the total addition to (the log of) consumption expenditure in the second and third quarters. Furthermore, define α as the share of this added spending that is carried out in the second quarter. An alternative formulation of the myopic alternative then is

(8') $$H_1: \quad \begin{cases} b_2 = \alpha \cdot \mu, \quad & \mu > 0 \\ b_3 = (1-\alpha) \cdot \mu & \\ b_4 = \mu/2 & \end{cases}$$

The estimate of μ may be used to derive a rough measure of the marginal propensity to consume under the myopia alternative. On average, Norwegians pay about 30% of their incomes in taxes. Accordingly, the tax suspension for one month in the summer corresponds roughly to a 10% addition to the disposable income for either the second or the third quarter. Thus, because we multiply all our data by 100 so as to express the growth rates as percentages per quarter, $\mu/10$ may serve as a rough estimate of the marginal propensity to spend out of the extra income from the reduced tax withholding.

The presence of a seasonal unit root in the residual u_t is consistent with the alternative hypothesis. It should be noted that the usual test criteria for the coefficient constraint in (8) are invalid if such a root is present. However, because no seasonal unit root is present under the permanent-income hypothesis, the test of the constraint H_0 in (6) is valid on first-difference form even if the hypothesis of a seasonal unit root cannot be rejected.

4. Comparisons of the Seasonality in Norwegian and Swedish Consumption

We have obtained quarterly, seasonally unadjusted data for personal consumption expenditure from the Central Bureaus of Statistics of Norway and Sweden. We have nominal and inflation-adjusted data for total consumption as well as for components, such as goods consumption, energy consumption, and expenditures on motor vehicles. The data for goods consumption unfortunately leave out some of the typical vacation expenses, such as transportation and hotel and restaurant services. On the other hand, however, the seasonality of these series probably is much less reliable because the seasonal factors of many subcomponents of service consumption are imputed and thus uncertain. From what we have been told by the statistical agencies, the seasonal factors of the components of goods consumption are derived from the primary sources without the use of interpolation methods.

The data series for consumption expenditure are derived from retail sales and other primary sources, as explained above. However, whereas retail sales are defined as transactions having taken place within the country, the data for consumer spending should represent the spending of domestic consumers both at home and abroad. For this reason, the consumption series for both countries contain corrections for the spending of domestic consumers abroad and for the spending by foreigners in the home country. Unfortunately, however, these corrections are known to be unreliable. For better reliability as well as better comparison with the retail-sales data we therefore employ the consumption series that have not been thus corrected. Although we regret that this procedure prevents us from including the spending on vacations abroad, we believe that data quality is more important than coverage.

For Sweden, the consumption data reflect a revision undertaken in 1993. In this process, data going back to 1980 were revised, and the level data before this date were rescaled to preserve comparability over time. The rescaling amounted to level adjustments of 0-0.5% for each of the subgroups of consumption. The Norwegian data from before 1978 have been rescaled in a similar way and by a similar amount. The data for total consumption and goods consumption start in 1966:1 for Norway and in 1970:1 for Sweden, allowing us a sample for comparison of 1970:1-1993:1. First-differencing and correction for fourth-order serial correlation furthermore imply the loss of the observations for 1970:1-1971:1, so that our effective estimation sample is 1971:2-1993:1. Unfortunately, the data for energy consumption and expenditure on automotive equipment start later, so that the comparisons based on data excluding these components do not start until 1981:2.

In the figures in at the end of the paper, we have plotted the main series that we employ. The consumption data exhibits a clear seasonal pattern with a peak in the fourth quarter and a through in the first. These effects appear to be somewhat stronger for Norway than for Sweden. Further, Swedish consumption seems to be a little down in the third quarter. The retail sales index is seasonally similar but the through in the first quarter is more evident. This reflects that energy consumption is left out of these data.

In Table 11, we report on the integrational properties of the data. The tests are based on the procedure of Hylleberg, Engle, Granger and Yoo (1990). We have included seasonal means. A trend term is used in the level data but not in first-differenced data. The consumption data are in real terms and the retail sales index are based on nominal values. We can not exclude a unit root at the zero frequency in any of the series in levels, which is consistent with the implications of the theory. In the consumption data this root disappears when data are first-differenced. In the retail sales index we can not draw the same conclusion. We can not reject a unit root when data are in first differences as well, although the situation improves. This is possibly due to non-stationarity in the price index. However, if logs of both real retail sales and the price data is I(1) first differencing should result in a stationary series. Thus, the non-stationarity of the price data must be more complicated than that.

We find little evidence of roots on the seasonal frequencies. The only exception is actual retail sales in Norway, where we can not disregard the possibillity of a root on the biannual frequency. This carries over to the difference between Norway and Sweden. However, overall there is little evidence of seasonal nonstationarity. The lack of seasonal unit roots increases our confidence in the data and helps us assure that our testing procedure is valid.

Table 1 reports estimates of equation (4) for the inflation-adjusted, Norwegian consumption data. The parameters ρ_1, ρ_2, ρ_3, and ρ_4 are autoregressive serial-correlation parameters of the residual Δe_{it} estimated simultaneously with the b_{ik} by Generalised Least Squares (GLS). GLS in this case avoids the pitfall of Flood and Garber (1980) because the seasonal dummies are exogenous. The last two rows in Table 1 present the results of tests for seasonal unit roots in the error term, based on the test with seasonal means and fourth-order autoregression proposed by Dickey, Hasza, and Fuller (1984), with the refinement by Osborn, Chui, Smith, and Birchenhall (1988)[6]. A t-value less (further away from zero) than the approximate critical value in the last row indicates rejection of the presence of a seasonal unit root.

The results for total consumption and goods consumption for the full sample are shown in the first two columns. The seasonal fluctuations clearly are substantial and reflects the

[6]The two tests seem to give much the same answers here. The HEGY test carries more information of the frequency of seasonal integration.

visual observations we made from the figures. Norwegian consumption peaks every year in the fourth quarter, apparently reflecting the Christmas shopping season. This peak is followed by a trough in the first quarter, whereas the difference between the second and third quarters is negligible for total consumption. Removing services from the data raises the peak in the fourth quarter and introduces a decline from the second to the third quarter. However, the differences between total consumption and goods consumption are not large. In the following, we concentrate mainly on goods consumption for the reasons mentioned above.

The tests for seasonal unit roots still clearly indicate rejection. This finding is consistent with the view of the measurement errors implied by the Norwegian data collection process as expressed in (1). Thus, it should be meaningful to study the seasonal patterns in terms of seasonal means as indicated by the slope coefficients in this table.

The next four columns of the table repeat these estimates for two subsamples, 1971:1-1985:1 and 1985:2-1993:1. The seasonal pattern of consumption could change over time for three reasons. First, if the series contain seasonal unit roots (due to measurement errors or in the true process), the estimates of the coefficients of the seasonal dummies would be likely to change over time as the seasonal innovations accumulate. Second, seasonal preferences could change over time in other ways than those implied by seasonal unit roots. Third, if, contrary to our prior belief, liquidity constraints are important in shaping the seasonal pattern of consumption, this pattern should be affected by the liberalisation of the Norwegian and Swedish credit markets in the mid-1980s. Naturally, to the extent that changes over time are observed, we are unable to determine which one, if any, of these two forces have been at work, or whether consumers simply have changed their seasonal tastes[7].

[7] The liberalisation of the credit markets was a somewhat gradual process in both countries and thus didn't happen exactly at the beginning of the second quarter of 1985. The choice of this particular date actually was determined by a different consideration, namely, the construction of the new series of retail sales for Sweden, as mentioned in Footnote 4. This series goes back to the beginning of 1984 so that, after first differencing and correction for fourth-order serial correlation, it could be studied from 1985:2. Although this series is preferable to the old one by being available for Total as well as Actual Retail Sales, and in constant as well as current prices, we do not report results using this series because we could not detect any significant differences in the seasonal patterns between the old and the new series. Considering the

We find some changes in the seasonal pattern, but not enough to change the overall picture. The Christmas peak becomes a little less pronounced over time, and consumption in the third quarter is reduced relative to the others. These changes could be taken as indications of relaxed liquidity constraints, perhaps because people may start spending their Christmas income before they receive it and because they may borrow against their summer paycheques to take winter vacations half a year earlier. The changes over time are statistically significant for total consumption, but not for goods consumption, as can be read from the results of the stability test in the third line from below in the first and second columns[8]. However, in both cases the changes are modest.

The fact that the changes in the deterministic seasonal parameters are small provide additional evidence that accumulating measurement errors have not been important. For the first subsample, the hypothesis of a seasonal unit root cannot be rejected; however, given the rejection for the full sample, we suspect that this nonrejection can be ascribed to the lower power in shorter samples. In fact, the rejection for the even shorter second sub sample is quite striking for a test known for its low power.

The data series excluding energy and energy and automotive vehicles start at a later date, as noted above. Therefore, we have not made any attempt to split the sample for these series, but present the full-sample results in the last two columns of Table 1. Only four years separate the starting date of this sample from the second sample of the overall data, so we should be able to compare the seasonal patterns with overall goods consumption for that subsample. The removal of energy deepens the trough in the first quarter, which covers most of the heaviest heating season—there is no significant cooling season in the Nordic climate. Removing expenditure on transportation equipment sharpens the Christmas peak, but causes no other significant changes.

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fact that we so far have less than 40 observations of the new series, we therefore report only the results of the old series.

[8] The numbers reported are the p-values of F tests of the hypothesis that β_2, β_3, and β_4 are constant across the two sub-samples, under the maintained hypothesis that $\beta_1, \rho_1, \rho_2, \rho_3$, and ρ_4 are constant. We also carried out the same test letting the latter set of parameters vary. The results were largely similar, although in many cases the rejections of stability were somewhat clearer.

Table 2 presents the corresponding results for Sweden. The Swedish data for total consumption and overall goods consumption also show a peak in the fourth quarter, but not as tall as for Norway. There is again a trough in the first quarter, but shallower than for Norway. Furthermore, there is a trough in the third quarter relative to the first, which we didn't find for Norway except for a much weaker trough for goods consumption in the second subsample. Thus, we do seem to find evidence that Norwegian consumers tend to concentrate more of their spending in the third and fourth quarters than their Swedish counterparts.

The tests for seasonal unit roots lead to rejection for the full sample also in the case of Sweden, although not quite as decisive as for Norway. The previous HEGY-tests confirm that seasonal roots are unlikely. Thus, the accumulating measurement errors for Sweden may not be as serious as what might be feared.

Again, we repeated the exercise for the two subsamples defined above. The changes now are statistically significant for both series. The main changes seem to be a deepening of the trough in the third quarter and a smoothing of the peak in the fourth. Credit-market deregulation is not the main candidate for explaining such changes in the Swedish case because the Swedish tax authorities withhold taxes at a constant rate year round. Thus, in the Swedish case, it seems more reasonable to ascribe the observed changes either to accumulating measurement errors—which are not suggested by the rejection of the unit-root test—or to changes in seasonal tastes.

It is interesting to note that the changes over time tend to go in the same direction for both countries. This similarity seems to provide some credence to our maintained hypothesis of identical seasonal preferences. Taking out energy and motor vehicles has similar effects for Sweden as for Norway, which may be taken as additional evidence for the same hypothesis.

Table 3 compares the results for Norway and Sweden by presenting estimates of equation (5). The third and fourth lines in the bottom panel show the marginal significance levels of Wald F-test of the hypotheses H_0 and H_1, in (6) and (8). As in the two preceding tables, the last two lines present the test results for a seasonal unit root. Note that the tests of H_0 are valid independently of the outcome of the tests for seasonal unit roots because no such root exists under H_0. The tests of H_1, are strictly invalid in the presence of seasonal unit roots. However, because of the low power of the unit-root test, nonrejection of the seasonal unit-root hypothesis is very weak evidence for the existence of a unit root. Thus, we present the results of H_1—as well as the standard errors of the estimated parameters under the assumption that the residual series are stationary.

The results in Table 3 are quite striking. For the full sample as well as the subsamples, all of the estimates of β_3 and β_4 are positive and quite significant individually. For total consumption and overall goods consumption, the estimates for β_2 are positive as well, although insignificant and not as large. For goods consumption excluding energy the estimate of β_2 is negative, but not significant. Overall, these findings clearly appear to contradict the seasonal implications of the permanent-income hypothesis. The restriction (6) that the seasonal patterns in consumption should be equal for the two countries and thus unrelated to the tax withholding system, is rejected decisively. For the full sample, the alternative hypothesis in (8) is rejected marginally for total consumption and clearly for overall goods consumption. Taking out energy and automotive equipment makes the rejection of (8) more decisive. It should be emphasised that the formulation of the myopic alternative in (8) is rather extreme. Even so, however, the data come much closer to corresponding to this alternative than to the null.

The changes over time are smaller in this table than in the two preceding ones because the changes for the two countries tend to go in the same direction. They are not statistically significant, which provides additional evidence in favour of our maintained hypothesis of similar seasonal preferences. However, we do find that the estimate of β_4 increases over time and thus removes the data further from both the null and the strict alternative H_1, in the second subsample. This change is mainly due to the smoothing of the fourth-quarter peak for Sweden and thus does not seem to be a result of the liberalisation of the Norwegian credit markets.

We also estimated constrained versions of these equations under the alternative hypothesis so as to obtain estimates of μ and α. The estimates for α vary between 0.15 and 0.40, which is consistent with our prior belief that Norwegians spend most of their extra vacation pay in July and August. The estimates of μ turn out to be striking, however. Recall that one-tenth of this parameter is a rough estimate of the marginal propensity to consume. The estimates of this parameter hover around 10, indicating a virtually unitary marginal propensity to consume. This result seems very strong. In fact, we find it so strong that a further examination of the underlying retail-sales data seems in order.

Because the Swedish retail-sales data before 1984 only are available in nominal terms, we are forced to undertake this further examination with nominal data. For easier comparison between consumption data and retail-sales data, we repeated the exercises reported in Tables 1-3 on nominal consumption data. The results are reported in Tables 4-6. Under the assumption that the seasonal properties of the Swedish and Norwegian technologies are similar, the analysis of nominal data should be equally valid. Comparison between Tables 1 and 4 shows that the implicit deflators for the respective categories of consumption for Norway display very little seasonality. Formal tests of the type used in Table 3 find this seasonality to be statistically insignificant. Although similar tests do detect significant seasonality for Sweden and significantly different seasonality for deflators between Norway and Sweden, a comparison of Tables 2 and 5 reveals the seasonality in Swedish prices to be quite modest. Thus, our analysis will proceed under the assumption that any interesting difference in the seasonality between Norwegian and Swedish nominal data stems from the real volumes rather than their prices.

5. Comparisons of the Seasonality in Norwegian and Swedish Retail Sales

As noted in Section 2, the retail-sales data for Sweden are not adjusted for inflation, and they cover only «Actual Retail Sales»[9]. The latter restriction means that the sales of automotive dealers, wine and liquor stores (a government monopoly), and pharmacies are left out. Most of the data for the consumption of goods, except energy and vehicles, are derived from the various subcategories of actual retail sales. We have reconstructed a similar series for Norway on the basis of the definition of the Swedish series.

The retail-sales data are not only seasonally unadjusted, but also unadjusted for the number of business days. They are compiled monthly for both countries. However, for more direct comparison with the consumption data, we start by studying the retail sales data on the quarterly frequency.

The seasonal patterns of these series are compared in Table 7. The first three columns show the results for the sample 1973:1993:1, which is the longest period for which both data series are available. The next three columns make the same comparison for the period for which the data for goods consumption excluding energy and vehicles are available. The last three columns repeats the exercise for the period after the credit-market deregulation.

The results for the full sample resemble those for the consumption data in the sense that the seasonal dummies for the third and fourth quarters are significantly higher for Norway than for Sweden. However, the largest difference now occurs in the fourth rather than the third quarter. Thus, although the null hypothesis of no seasonal difference (H_0) is rejected decisively, the simple myopia alternative (H_1) is clearly rejected as well. Furthermore, the seasonal dummy for the second quarter is lower for Sweden, although not statistically significant (t = - 1.87). To check that the rejection of H_0 does not come from this negative difference we carried out a separate exclusion test for only those dummies whose coefficients come out positive (the third and fourth

[9] See footnotes 4 and 5 regarding a new and more complete data series for Sweden, starting in 1984.

quarters). This hypothesis, denoted $\tilde{H}_0$ in the table, clearly is rejected as well. Nevertheless, it seems fair to conclude that the link between tax withholding and the seasonal pattern of spending does not stand out quite as clearly for the retail-sales data as it does for consumption. The test results for seasonal unit roots and parameter stability (with the sample split between the first and the second quarter of 1985) are similar to those for the consumption data.

The middle columns cover the same samples as those of goods consumption excluding energy and vehicles in Tables 1-6. The principles involved in the construction of the consumption data suggests that the seasonal pattern should be similar for actual retail sales and this subcategory of consumption. The correspondence seems to work out quite well for the Norwegian data. For the Swedish case, however, the seasonal variation is stronger for the retail-sales data than for the consumption data. The Christmas peak is higher in particular. The retail-sales data also show a higher level in the third quarter than the first, which we did not find for the consumption data.

Because this sample is shorter, the standard errors are larger than for the full sample. Nevertheless, the test result for $\tilde{H}_0$ shows that the seasonal dummies for the third and fourth quarter are significantly greater for Norway for this subsample as well. The difference for the second quarter remains negative and more so than for the full sample; but the t-statistic is closer to zero.

For the last subsample, the difference in the Christmas peak is a little smaller because of a slight reduction in the peak for Norway. This reduction, combined with the larger standard errors, prevents $\tilde{H}_0$ from being rejected for this subsample.

Table 8 looks at the seasonal patterns in the same data, but monthly, and based on the method in equation (7). The correction for serial correlation is done for the monthly lags of 1, 2, 3, and 12, which appeared to be sufficient.

The monthly data confirm that the significant difference for the first quarter is concentrated in December, as expected. However, substantial differences in the same direction also are found for all of the months for August through November. Furthermore, the negative difference that we found for the second quarter in Table 7 turns out to extend from March through May, with a peak in April. This pattern clearly is more complicated than the one suggested by our naive myopia alternative, where all extra disposable income is spent during the month it is received or perhaps the following one. However, the pattern still is compatible with Norwegian spending being influenced by the timing of tax withholding in that the difference is positive from the month when the first wage earners receive their vacation pay (i.e. June) through the month after the Christmas salary payment (i.e. January). In between those months, Norwegian consumers seem to spread their spending out somewhat by apparently saving some of their extra vacation pay. However, they don't seem to start spending their vacation pay until they receive it. This behaviour seems quite compatible with the behavioural life-cycle hypothesis.

The negative differences for March through May may further have been influenced by a slight difference in holiday patterns in that Norway has one more legal holiday during Holy Week (Maundy Thursday) and one additional national holiday in May (Constitution Day). Moreover, Norwegian consumers have further developed a pattern of taking additional vacation days during the rest of Holy Week, for an extended Easter vacation; and it has become customary to spend this week in the mountains or at the sea rather than shopping in retail stores.

The hypothesis $\bar{H}_0$ is defined in Table 8 as the constraint that all the monthly coefficients with positive signs (except February) be zero. As can be seen from the third column, this hypothesis is clearly rejected for the full sample.

Interestingly, the hypotheses of seasonal unit roots are also rejected decisively for all series for the full sample. However, we do find significant changes over time in the coefficients of the seasonal dummies (the break point now goes between March and

April of 1985). This change now is significant also for the difference between the two countries, but not nearly as significant as for each country viewed separately. The changes over time make the difference dummies change sign for June. Furthermore, H_0 is not rejected for the sample starting in 1981. However, the positive and significant difference for December remains for both subsamples.

6. Reconciliation of the Series for Consumption and Retail Sales

The differences we find between consumption and retail sales suggest that we take a closer look at the relationship between these series. Table 9 provides some insights into this relationship for the Norwegian data. This table has seven columns. The first two columns present the quarterly seasonal patterns in total and actual retail sales, respectively. The third, fourth, and fifth columns show the corresponding patterns for goods consumption, goods consumption excluding energy, and goods consumption excluding energy and automotive vehicles. The sixth column compares the seasonal patterns for total retail sales and goods consumption excluding energy. According to the information we received from Statistics Norway, the coverage of these two series should be close. Similarly, the seventh column makes a similar comparison between actual retail sales and goods consumption excluding energy and automotive vehicles, whose coverage also should be similar[10]. All the estimates are based on nominal data for 1981:2-1993:1, which is the longest period for which the consumption series for energy and automotive equipment is available for both countries.

The energy component clearly and completely explains the difference in seasonal behaviour between goods consumption and total retail sales for Norway. Because home cooling is not an issue in Scandinavia, it is not surprising to find this spending to be

[10] The energy component of consumption includes home heating and electricity, but not automotive fuels. The latter category is, on the other hand, included with consumption spending on vehicles. we chose to treat automotive fuels this way for easier comparison with the retail-sales data. Total retail sales are reported to include the sales of automotive dealers and gasoline stations, but not deliveries of home heating fuels or electricity. Actual retail sales also exclude the sales of automotive dealers and gas stations.

concentrated in the first quarter. Taking out this component therefore raises the coefficients of the seasonal dummies for the remaining three quarters. The largest change takes place in the third quarter, which is also the warmest. After this correction, the seasonal pattern of goods consumption is virtually identical to that of total retail sales. The differences, displayed in the sixth column, clearly are both insignificant and small. The hypothesis of a seasonal unit root cannot quite be rejected for the series of differences, but this non-rejection may be due to low power. The significant first-order serial correlation may be explained by rounding errors in the reported level data, giving rise to first-order moving averages in the measurement errors of the first differences. We conclude that the Norwegian data for goods consumption display no seasonal patterns other than those derived from the corresponding retail-sales data.

When spending on vehicles (including automotive fuels) are taken out as well, the rest is supposed to look like actual retail sales. Again, these adjustments bring the seasonal patterns in the consumption data close to those of the corresponding retail sales data, although the correspondence now fails slightly for the fourth quarter. As a result, the differences in seasonal patterns become borderline significant. Even so, we feel that the correspondence is close enough to support our confidence in the essential correctness of the seasonal patterns of the Norwegian consumption data.

Because the new Swedish data series for total retail sales only goes back to 1984, we concentrate on the investigation of the Swedish data on the relationship between actual retail sales and goods consumption excluding energy and vehicles. The results are displayed in Table 10. As in the Norwegian case, the exclusion of energy from the consumption data increases the magnitudes of all the seasonal dummies, and the most so for the third quarter. As a result, the seasonal patterns of goods consumption excluding energy are roughly half-way reconciled with those of actual retail sales, compared to those of goods consumption when energy is included. However, when vehicles are taken out of the consumption data as well, the reconciliation is partly reversed. Thus, although the seasonal pattern of goods consumption excluding energy and vehicles is closer to that of actual retail sales than the unadjusted goods consumption data, the difference is highly significant, as shown in the last column.

We consider this discrepancy as the main reason for the difference in our results between the analysis of retail sales and consumption data[11]. This discrepancy does not appear to be a result of the Swedish data-collection method discussed in Section 2. Whereas that method potentially could introduce accumulating measurement errors in the seasonal data, we see no reason why these errors should change to such an extent in the conversion from retail sales to consumption data.

Despite repeated attempts, we have not been able to pinpoint the exact reason for the seasonal discrepancy between consumption and retail sales for Sweden. In general, such discrepancies are hardly surprising because the construction of consumption data from retail-sales and other raw data involve conversion keys as the sales of different kinds of retail stores are converted into consumption of the respective categories of consumption. Because the keys vary from category to category and the individual components have different seasonal patterns, discrepancies may arise between the seasonal patterns of the retail-sales series and the consumption series we have examined here. What we find surprising, however, is that we find such large discrepancies for Sweden when no or only small discrepancies are found for Norway.

This conclusion still does not tell us which country has the better data quality. The consistency between the Norwegian retail-sales and consumption data suggests, however, that the Norwegian seasonal pattern perhaps may be more reliable. Even so, we cannot tell for sure whether the retail-sales data or the consumption data are more reliable for Sweden. Our prior goes in the direction of presuming that the retail-sales data are more reliable because they have been subject to less processing. However, after

[11] Using the new Swedish series for total retail sales starting in 1984, we also have analysed the relationship between this series and goods consumption excluding energy only for Sweden. The results then look a little better than those in Table 10, but the discrepancy remains substantial. For example, the coefficient for the fourth quarter in the comparison remains as high as 79, with a standard error of 0.8.

extended contact with the various departments of Statistics Sweden, we have been unable to identify any errors in the conversion process[12].

7. Discussion and Conclusion

The most robust result of our analysis seems to be that the Christmas spending peak is significantly sharper for Norway than for Sweden. This result is found in consumption data as well as retail-sales data, with and without correction for inflation, and on monthly as well as quarterly data for retail sales. This difference correlates clearly with the practice of withholding taxes at half the normal rate for December, which is followed in Norway, but not in Sweden.

For consumption data, we find an even larger difference, in the same direction, for the third quarter, which correlates with the Norwegian practice of no tax withholding for June or July. This difference is considerably weaker for the retail-sales data, however. Although the difference in coverage explains the difference between consumption data and retail-sales data for Norway, that is not the case for Sweden. Without further information about the relative quality of the retail-sales data and consumption data for Sweden, it seems appropriate to conclude that the observation of higher summer consumer spending for Norway than for Sweden should be considered somewhat uncertain.

For both consumption and retail sales, the data clearly reject the hypothesis of identical or even similar seasonal spending patterns for Sweden and Norway. When we correlate this finding with the difference in seasonal tax withholding, we feel that this finding provides significant evidence against the seasonal implications of the permanent-income hypothesis, with a possible caveat for summer spending in the retail-sales data. Furthermore, we find no evidence that this difference can be explained by liquidity

[12] A possible source to the Swedish discrepancies is the use of different base years for the seasonal pattern. Equation 2 implies that differences will tend to last indefinately. The Swedes were unable to explain the sources of their seasonal pattern fully.

constraints, mainly because it does not go away after the creditmarket liberalisation in the mid-1980s. In contrast, we feel that our findings easily can be explained in terms of self-imposed constraints against using money before it has been received.

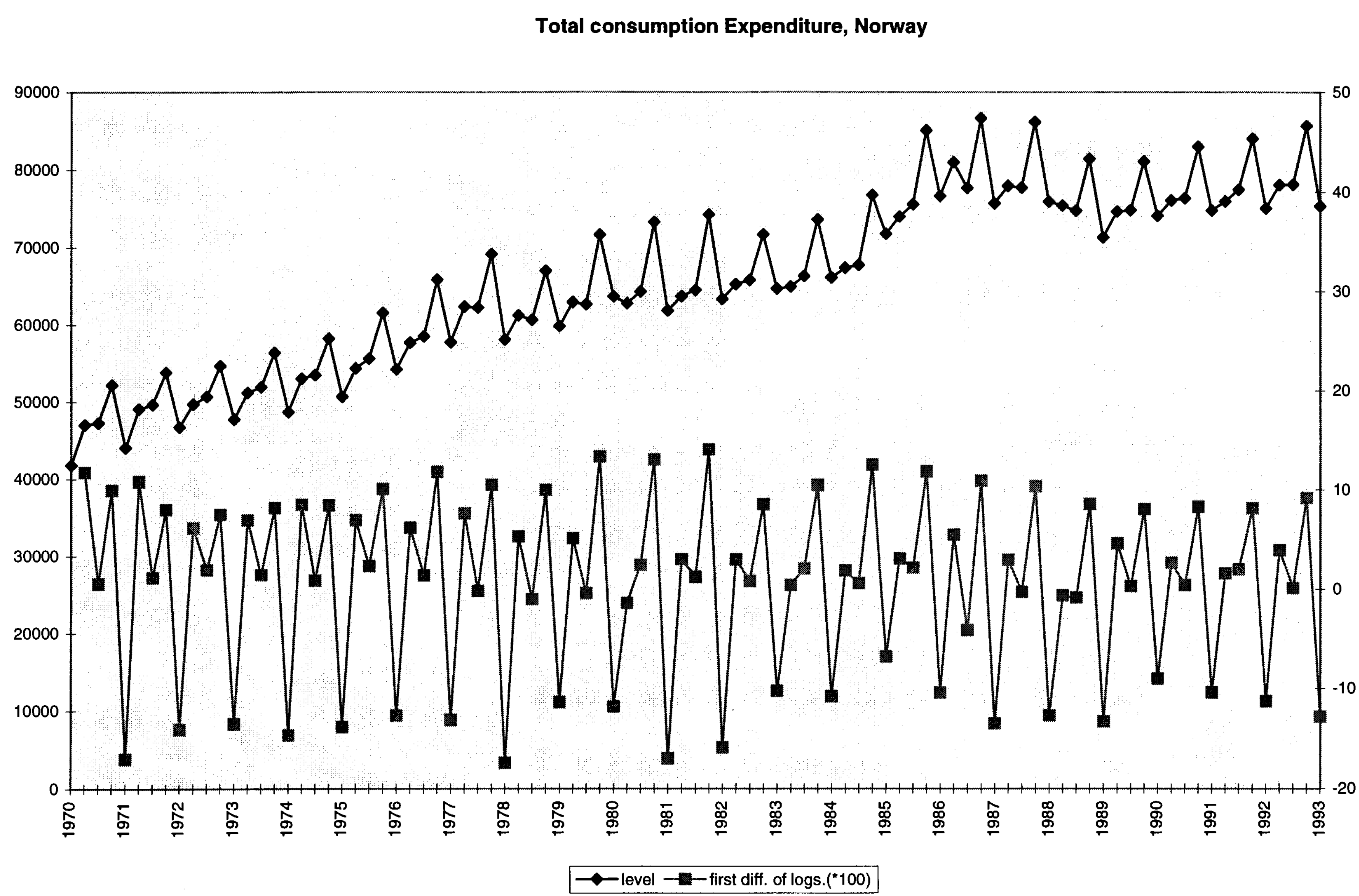
The remaining interesting question is whether our maintained hypothesis of identical seasonal preferences fail. In particular, the question is whether the particular taxwithholding rules for Norway have been caused by a special, national preference—not shared by Swedes—for spending during summer vacation and the Christmas holiday season.

Although this idea of reversed causality may seem appealing to economists believing in rational behaviour, we find some important arguments against it. First, Norwegians do not concentrate their spending during the summer, which is when their tax withholding is the lowest. Rather, Norwegians' spending—like most everybody else's peaks during the Christmas shopping season. If Norwegian consumers wanted to use the government as their saving agent, why then did they pick the summer for a month completely without tax withholding and settle for half the normal withholding rate in December?

Second, the reversed-causality claim does not seem to square well with the history of the introduction of vacation pay without tax withholding. This arrangement was championed as one of the key causes of the Norwegian labour movement, as part of the equal right of the working class to enjoy summer vacations. The purpose was to make it possible for workers to have summer vacation, not to make their saving transactions a little more efficient or convenient.

Third, the transactions costs of private saving are extremely low and have been so for a long time for both Sweden and Norway. Although the credit markets have been regulated, the banking system has been well developed for a much longer period than our data sample. Depositing and withdrawing money from saving accounts surely cannot have been a burden on Norwegian households.

Based on these considerations, we conclude that the timing of tax withholding seems to have affected the timing of spending by Norwegian consumers, despite the predictions to the contrary of the permanent-income hypothesis. However, our conclusion is softened somewhat by the discrepancy between the seasonal patterns in retail sales and consumer spending for Sweden, which suggests some problems of data quality. At this point, we feel more research and development of improved data collection and construction is called for.


Total consumption Expenditure, Norway
90000
80000
70000
60000
50000
40000
30000
20000
10000
0
50
40
30
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1970
1971
1972
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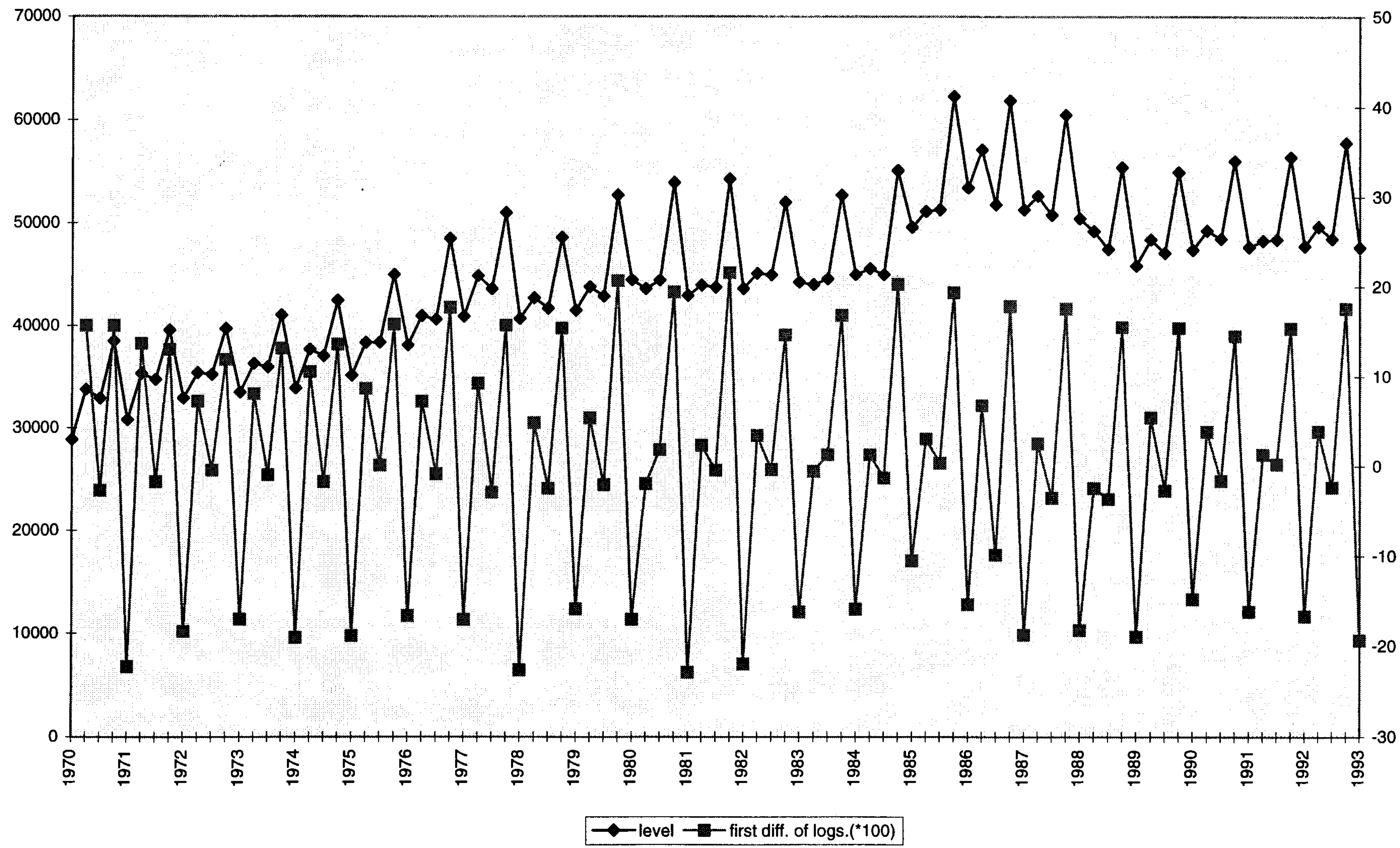
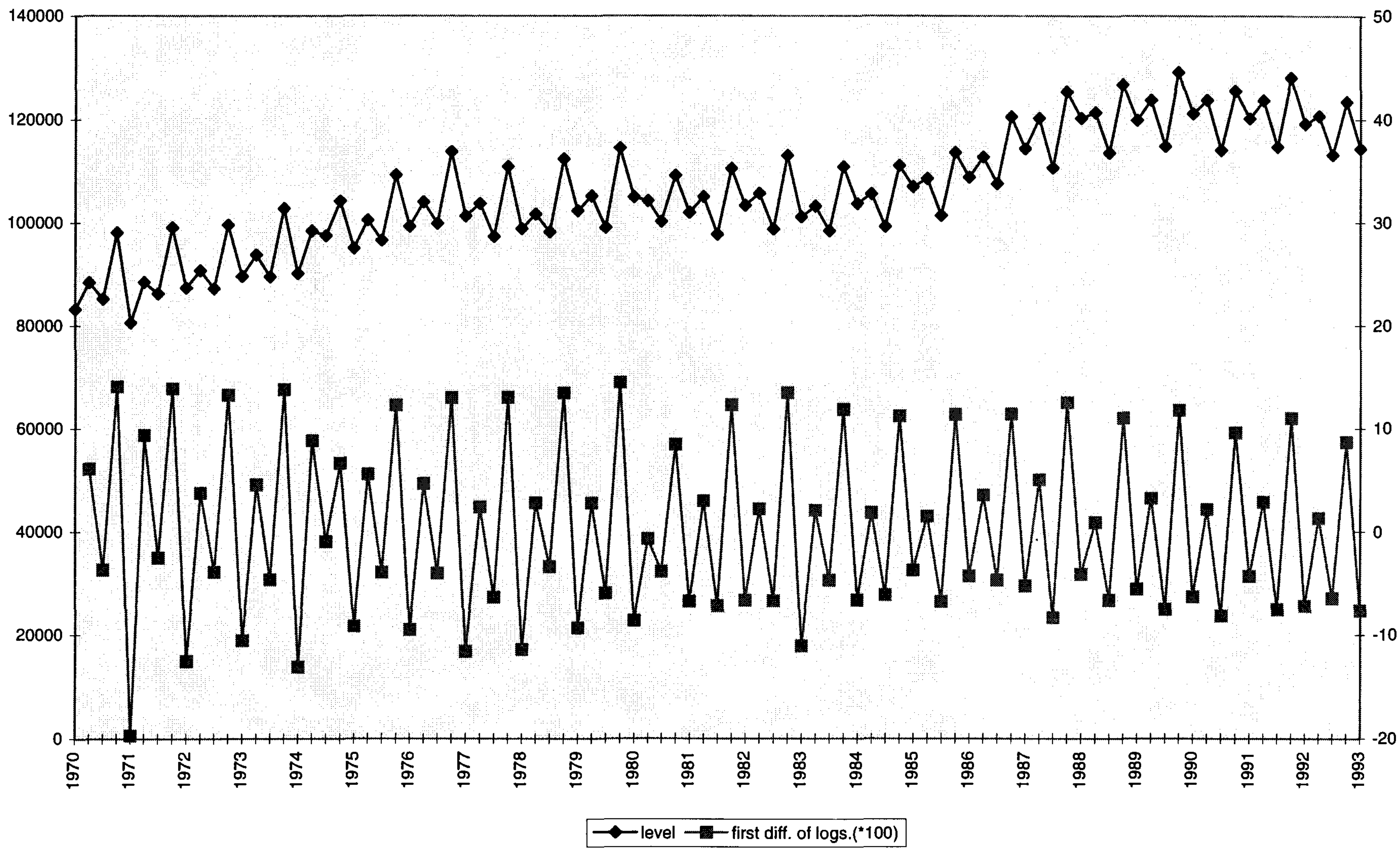
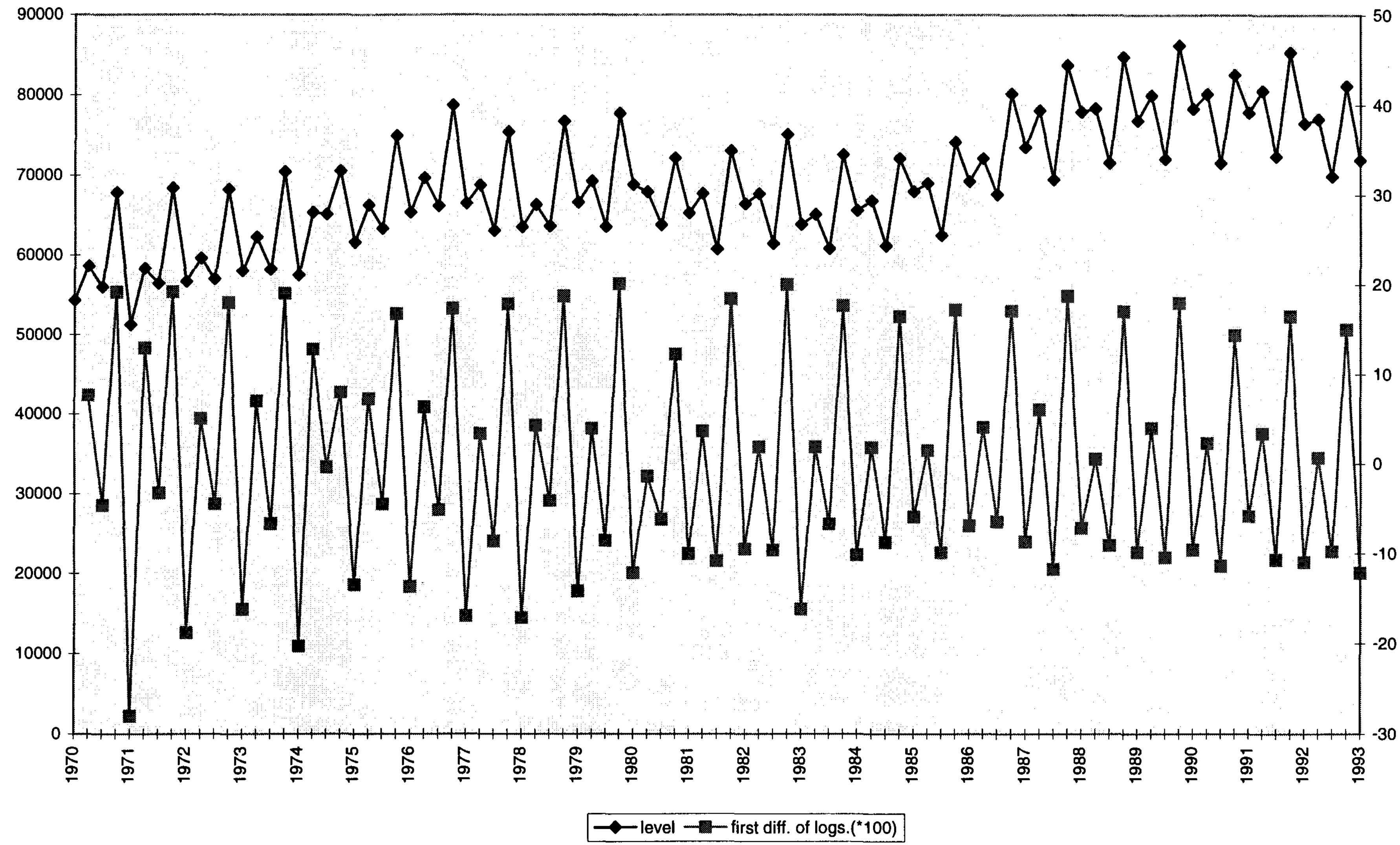
Consumption Expenditure on Nondurables and Services, Norway



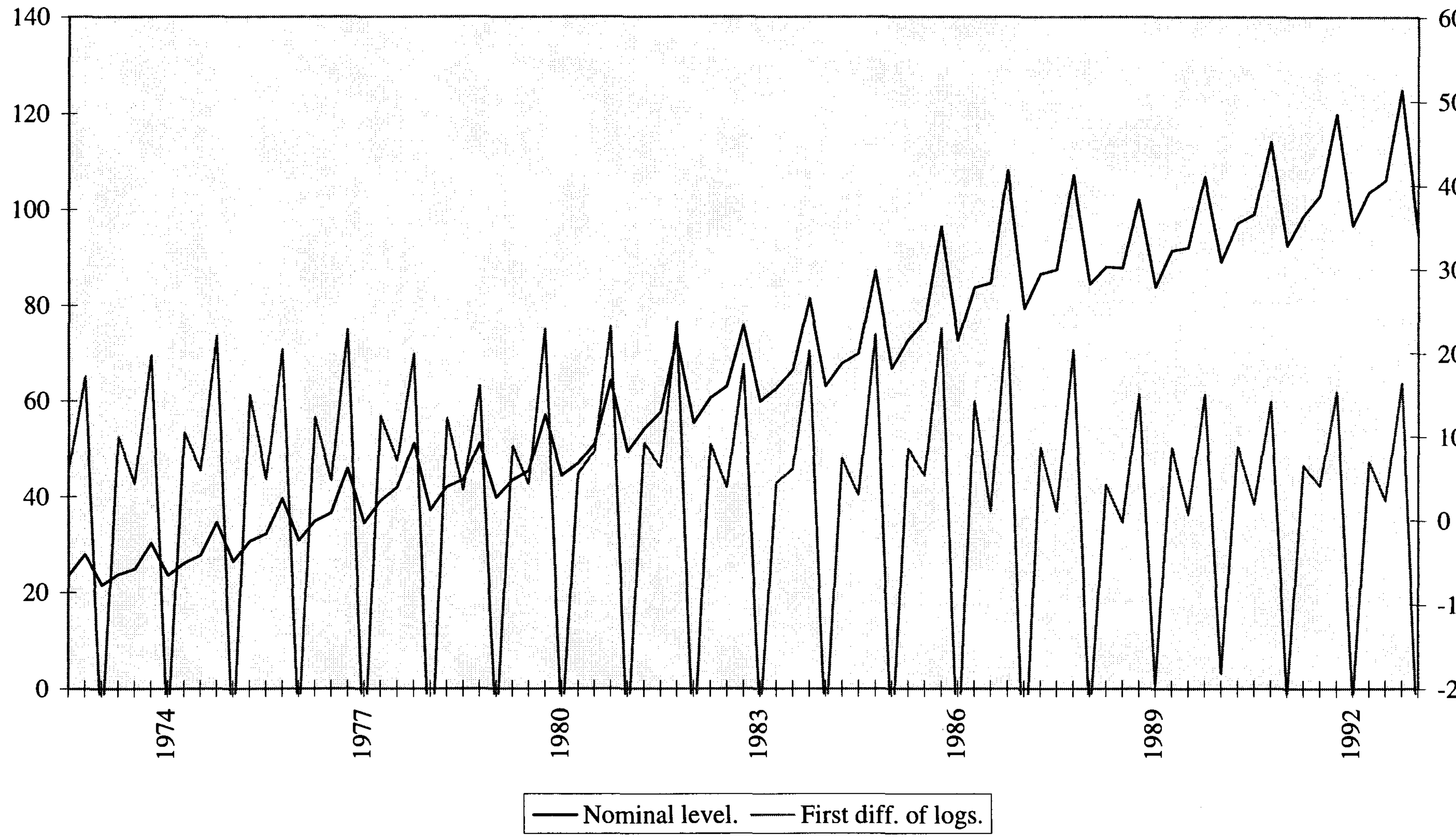
Total consumption Expenditure, Sweden



Consumption Expenditure on Nondurables and Services, Sweden



Actual Retail Sales, Norway



Actual Retail Sales, Sweden

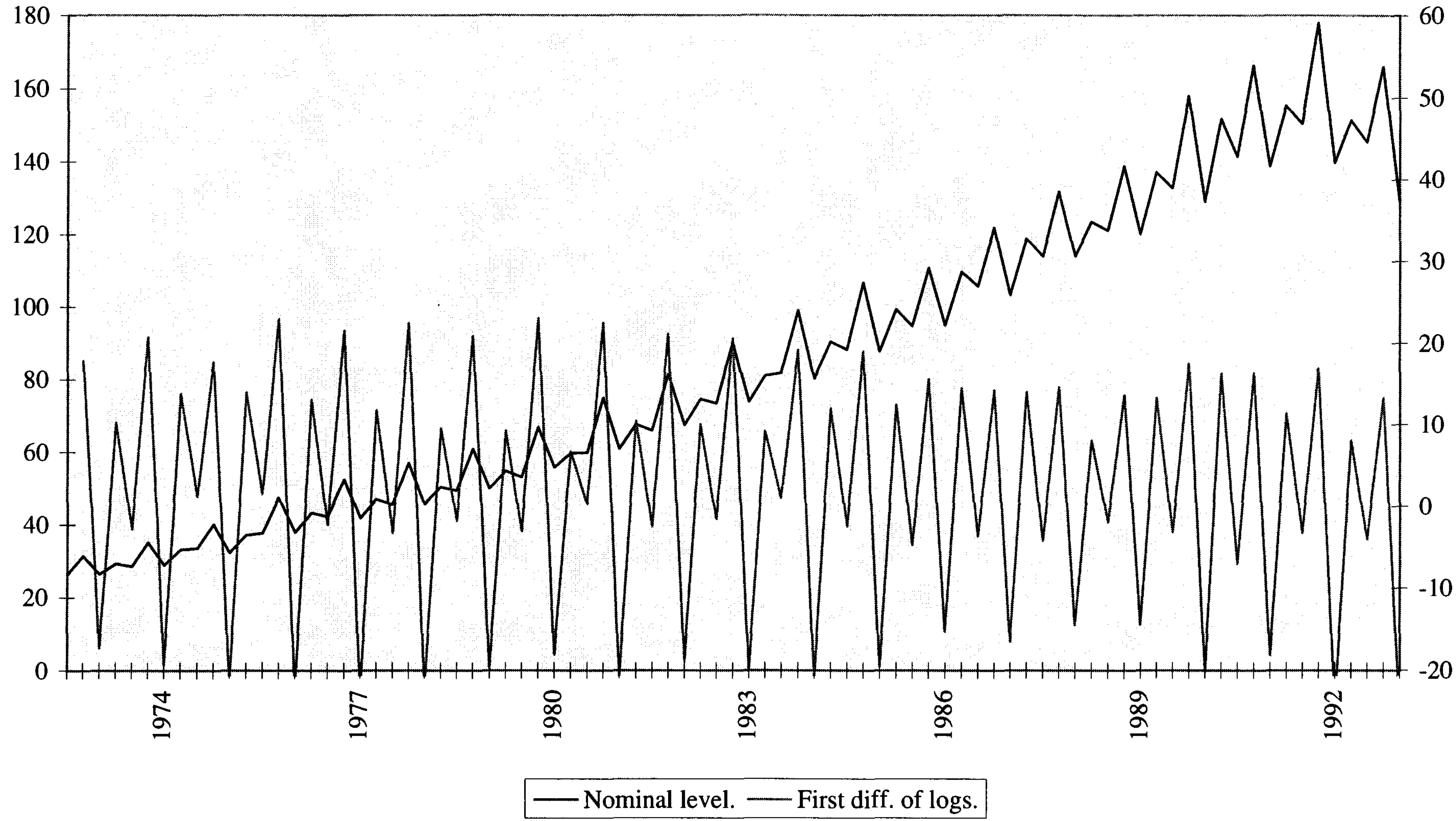


Table 1: **Seasonal Patterns in Norwegian Consumption Expenditure**

Dependent variable: $100 \times \Delta \ln c$
Estimation by GLS
Standard errors in parentheses

	Total Consumption 1971:2–1993:1	Goods Consumption 1971:2–1993:1	Total Consumption 1971:2–1985:1	Goods Consumption 1971:2–1985:1	Total Consumption 1985:2–1993:1	Goods Consumption 1985:2–1993:1	Goods Consumption Excl. Energy 1981:2–1993:1	Goods Consumption Excl. Energy and Vehicles 1981:2–1993:1
Constant	0.568 (0.243)	0.448 (0.324)	0.771 (0.209)	0.727 (0.278)	-0.003 (0.459)	-0.257 (0.594)	0.209 (0.482)	0.283 (0.265)
ΔQ_2	3.363 (0.736)	3.848 (0.951)	2.811 (1.348)	2.957 (2.038)	2.904 (0.581)	3.313 (0.784)	8.009 (0.918)	7.486 (1.110)
ΔQ_3	3.482 (0.854)	1.873 (1.195)	3.219 (1.669)	1.280 (2.625)	2.498 (0.470)	0.314 (0.702)	7.055 (0.830)	7.763 (1.055)
ΔQ_4	13.046 (0.731)	18.086 (0.944)	13.615 (1.320)	18.142 (1.959)	11.787 (0.569)	17.089 (0.773)	20.223 (0.915)	23.825 (1.105)
ρ_1	-0.200 (0.990)	-0.133 (0.101)	-0.327 (0.154)	-0.343 (0.149)	-0.209 (0.179)	-0.170 (0.185)	-0.125 (0.148)	-0.288 (0.139)
ρ_2	-0.125 (0.101)	-0.164 (0.102)	-0.327 (0.154)	-0.343 (0.149)	0.286 (0.170)	0.236 (0.177)	-0.136 (0.149)	-0.105 (0.147)
ρ_3	0.017 (0.101)	0.030 (0.102)	-0.090 (0.155)	-0.096 (0.152)	0.188 (0.171)	0.161 (0.178)	-0.009 (0.146)	-0.157 (0.144)
ρ_4	0.414 (0.093)	0.384 (0.094)	0.391 (0.137)	0.388 (0.134)	0.030 (0.171)	-0.000 (0.176)	0.139 (0.142)	0.228 (0.134)
SER	2.03	2.68	2.09	2.83	2.03	2.68	2.87	2.43
R^2	0.950	0.959	0.955	0.963	0.964	0.970	0.960	0.979
p(stability)	0.022	0.100						
Seas. unit root t-value	-5.07	-5.08	-3.17	-3.04	-5.74	-5.15	-4.95	-4.61
5% crit.value	-4.11	-4.11	-4.14	-4.14	-4.21	-4.21	-4.21	-4.21

Table 2: Seasonal Patterns in Swedish Consumption Expenditure

Dependent variable: $100 \times \Delta \ln c$

	Total Consumption	Goods Consumption	Total Consumption	Goods Consumption	Total Consumption	Goods Consumption	Goods Consumption Excl. Energy	Goods Consumption Excl. Energy and Vehicles
	1971:2–1993:1	1971:2–1993:1	1971:2–1985:1	1971:2–1985:1	1985:2–1993:1	1985:2–1993:1	1981:2–1993:1	1981:2–1993:1
Constant	0.347 (0.130)	0.308 (0.169)	0.376 (0.138)	0.310 (0.188)	-0.127 (0.744)	-0.028 (0.644)	0.100 (0.407)	0.126 (0.354)
ΔQ_2	2.702 (0.749)	3.510 (1.102)	2.740 (0.919)	3.781 (1.475)	2.338 (0.277)	2.648 (0.415)	8.144 (0.579)	5.039 (0.523)
ΔQ_3	-3.407 (0.927)	-4.868 (1.408)	-2.622 (1.090)	-3.568 (1.829)	-4.997 (0.316)	-7.547 (0.478)	-0.433 (0.725)	-3.736 (0.614)
ΔQ_4	7.947 (0.745)	11.917 (1.093)	9.288 (0.896)	13.501 (1.435)	5.682 (0.272)	8.900 (0.417)	12.220 (0.578)	10.831 (0.523)
ρ_1	-0.330 (0.097)	-0.360 (0.099)	-0.462 (0.130)	-0.483 (0.132)	-0.025 (0.206)	-0.023 (0.205)	-0.021 (0.160)	-0.114 (0.158)
ρ_2	-0.328 (0.103)	-0.401 (0.105)	-0.408 (0.141)	-0.488 (0.143)	0.197 (0.201)	0.156 (0.199)	-0.071 (0.155)	-0.010 (0.159)
ρ_3	-0.159 (0.103)	-0.182 (0.105)	-0.223 (0.142)	-0.266 (0.144)	0.513 (0.202)	0.410 (0.200)	0.220 (0.153)	0.197 (0.159)
ρ_4	0.258 (0.097)	0.213 (0.099)	0.112 (0.130)	0.090 (0.133)	-0.065 (0.220)	-0.102 (0.216)	0.099 (0.154)	0.148 (0.159)
SER	1.90	2.75	1.28	2.75	1.90	2.75	2.17	1.91
R^2	0.950	0.950	0.950	0.950	0.979	0.980	0.966	0.973
p(stability)	0.003	0.011						
Seas. unit root t-value	-4.83	-4.98	-4.20	-3.92	-3.03	-3.34	-3.64	-3.73
5% crit.value	-4.11	-4.11	-4.14	-4.14	-4.21	-4.21	-4.21	-4.21

Table 3: Comparison of the Seasonal Patterns in Norwegian and Swedish Consumption Expenditure

Dependent variable: $100 \times (\Delta \ln c_N - \Delta \ln c_S)$

	Total Consumption	Goods Consumption	Total Consumption	Goods Consumption	Total Consumption	Goods Consumption	Goods Consumption Excl. Energy	Goods Consumption Excl. Energy and Vehicles
	1971:2–1993:1	1971:2–1993:1	1971:2–1985:1	1971:2–1985:1	1985:2–1993:1	1985:2–1993:1	1981:2–1993:1	1981:2–1993:1
Constant	0.225 (0.235)	0.133 (0.325)	0.381 (0.167)	0.386 (0.226)	-0.144 (0.597)	-0.384 (0.748)	0.090 (0.711)	0.147 (0.452)
ΔQ_2	0.810 (0.598)	0.453 (0.952)	0.829 (0.877)	0.128 (1.393	0.438 (0.501)	0.474 (0.687)	-0.243 (0.742)	2.215 (0.827)
ΔQ_3	6.827 (0.716)	6.604 (1.187)	6.306 (1.097)	5.677 (1.799)	7.413 (0.463)	7.779 (0.715)	7.436 (0.765)	11.347 (0.828)
ΔQ_4	4.915 (0.599)	5.927 (0.952)	4.260 (0.879)	4.694 (1.396)	5.959 (0.498)	8.009 (0.684)	8.084 (0.743)	13.138 (0.829)
ρ_1	-0.222 (0.105)	-0.193 (0.106)	-0.398 (0.138)	-0.389 (0.135)	-0.135 (0.179)	-0.023 (0.194)	-0.114 (0.155)	-0.256 (0.149)
ρ_2	-0.146 (0.108)	-0.194 (0.106)	-0.397 (0.152)	-0.444 (0.148)	0.352 (0.179)	0.265 (0.185)	0.186 (0.146)	0.037 (0.149)
ρ_3	0.082 (0.107)	0.040 (0.105)	-0.156 (0.152)	-0.188 (0.148)	0.309 (0.181)	0.234 (0.186)	0.243 (0.149)	0.164 (0.146)
ρ_4	0.217 (0.103)	0.242 (0.102)	0.080 (0.137)	0.107 (0.134)	-0.130 (0.190)	-0.189 (0.193)	0.037 (0.150)	0.237 (0.141)
SER	3.34	3.36	2.31	3.32	3.35	3.36	3.18	2.56
R^2	0.764	0.647	0.750	0.627	0.882	0.824	0.791	0.926
$p(\mathrm{H}_0)$	0.000	0.000	0.000	0.002	0.000	0.000	0.000	0.000
$p(\mathrm{H}_1)$	0.066	0.015	0.443	0.224	0.000	0.000	0.000	0.000
p(stability)	0.099	0.271						
Seas. unit root t-value	-6.52	-5.84	-5.66	-4.98	-3.49	-3.88	-5.76	-5.23
5% crit.value	-4.11	-4.11	-4.14	-4.14	-4.21	-4.21	-4.21	-4.21

Table 4: Seasonal Patterns in Norwegian Nominal Consumption Expenditure

Dependent variable: $100 \times \Delta \ln c$

	Total Consumption	Goods Consumption	Total Consumption	Goods Consumption	Total Consumption	Goods Consumption	Goods Consumption Excl. Energy	Goods Consumption Excl. Energy and Vehicles
	1971:2–1993:1	1971:2–1993:1	1971:2–1985:1	1971:2–1985:1	1985:2–1993:1	1985:2–1993:1	1981:2–1993:1	1981:2–1993:1
Constant	2.252 (0.481)	2.132 (0.515)	2.877 (0.215)	2.824 (0.326)	0.961 (0.136)	0.750 (0.722)	1.494 (0.732)	1.575 (0.628)
ΔQ_2	3.384 (0.833)	3.698 (1.135)	2.632 (1.524)	1.960 (2.685)	3.023 (0.597)	3.383 (0.749)	8.162 (0.942)	7.964 (1.198)
ΔQ_3	3.563 (1.017)	1.998 (1.475)	3.292 (1.929)	0.997 (3.529)	2.425 (0.531)	0.199 (0.699)	7.134 (0.925)	8.094 (1.325)
ΔQ_4	12.837 (0.826)	18.062 (1.126)	13.672 (1.486)	18.399 (2.614)	11.282 (0.574)	16.675 (0.724)	20.208 (0.934)	24.106 (1.181)
ρ_1	-0.072 (0.087)	-0.043 (0.090)	-0.215 (0.126)	-0.186 (0.126)	-0.202 (0.185)	-0.191 (0.185)	-0.068 (0.151)	-0.058 (0.137)
ρ_2	-0.054 (0.088)	-0.123 (0.090)	-0.256 (0.130)	-0.308 (0.129)	0.230 (0.170)	0.201 (0.172)	0.141 (0.148)	0.010 (0.137)
ρ_3	0.112 (0.088)	0.099 (0.091)	-0.007 (0.132)	-0.051 (0.133)	0.306 (0.172)	0.270 (0.173)	0.008 (0.144)	0.012 (0.133)
ρ_4	0.562 (0.083)	0.518 (0.085)	0.504 (0.123)	0.495 (0.120)	0.166 (0.176)	0.081 (0.173)	0.275 (0.143)	0.443 (0.131)
SER	2.01	2.63	1.99	2.54	1.72	2.33	2.87	2.51
R^2	0.950	0.961	0.963	0.968	0.962	0.973	0.960	0.978
p(stability)	0.015	0.043						
Seas. unit root t-value	-4.22	-4.39	-2.43	-2.38	-4.21	-4.61	-4.95	-3.88
5% crit.value	-4.11	-4.11	-4.14	-4.14	-4.21	-4.21	-4.21	-4.21

Table 5: Seasonal Patterns in Swedish Nominal Consumption Expenditure

Dependent variable: $100 \times \Delta \ln c$

	Total Consumption	Goods Consumption	Total Consumption	Goods Consumption	Total Consumption	Goods Consumption	Goods Consumption Excl. Energy	Goods Consumption Excl. Energy and Vehicles
	1971:2–1993:1	1971:2–1993:1	1971:2–1985:1	1971:2–1985:1	1985:2–1993:1	1985:2–1993:1	1981:2–1993:1	1981:2–1993:1
Constant	2.395 (0.217)	2.169 (0.320)	2.731 (0.133)	2.639 (0.178)	1.734 (0.404)	1.043 (0.742)	1.416 (0.777)	1.235 (1.022)
ΔQ_2	2.179 (0.899)	2.919 (1.202)	2.417 (0.961)	3.509 (1.551)	1.744 (0.400)	2.619 (0.405)	7.568 (0.665)	4.440 (0.544)
ΔQ_3	-4.767 (1.095)	-6.160 (1.567)	-3.928 (1.175)	-4.596 (1.996)	-6.319 (0.471)	-8.745 (0.462)	-1.907 (0.828)	-5.037 (0.617)
ΔQ_4	6.127 (0.894)	10.912 (1.189)	7.833 (0.633)	12.394 (1.507)	3.730 (0.398)	8.089 (0.405)	11.292 (0.660)	9.887 (0.541)
ρ_1	-0.175 (0.092)	-0.122 (0.092)	-0.406 (0.123)	-0.400 (0.127)	0.064 (0.200)	0.066 (0.190)	0.042 (0.154)	-0.014 (0.152)
ρ_2	-0.176 (0.097)	-0.209 (0.095)	-0.398 (0.132)	-0.464 (0.134)	0.100 (0.210)	0.228 (0.180)	0.010 (0.155)	0.135 (0.152)
ρ_3	-0.071 (0.097)	-0.000 (0.095)	-0.232 (0.135)	-0.277 (0.137)	0.218 (0.211)	0.376 (0.183)	0.254 (0.148)	0.290 (0.149)
ρ_4	0.427 (0.097)	0.427 (0.096)	0.170 (0.127)	0.156 (0.131)	-0.159 (0.226)	-0.173 (0.206)	0.242 (0.158)	0.268 (0.161)
SER	2.02	2.71	1.85	2.62	1.67	1.86	2.26	1.95
R^2	0.939	0.951	0.957	0.939	0.959	0.979	0.965	0.972
p(stability)	0.015	0.057						
Seas. unit root t-value	-3.54	-3.43	-3.47	-3.24	-3.23	-3.84	-2.71	-3.35
5% crit.value	-4.11	-4.11	-4.14	-4.14	-4.21	-4.21	-4.21	-4.21

Table 6: **Comparison of the Seasonal Patterns in Norwegian and Swedish Nominal Consumption Expenditure**

Dependent variable: $100 \times (\Delta \ln c_N - \Delta \ln c_S)$

	Total Consumption	Goods Consumption	Total Consumption	Goods Consumption	Total Consumption	Goods Consumption	Goods Consumption Excl. Energy	Goods Consumption Excl. Energy and Vehicles
	1971:2–1993:1	1971:2–1993:1	1971:2–1985:1	1971:2–1985:1	1985:2–1993:1	1985:2–1993:1	1981:2–1993:1	1981:2–1993:1
Constant	-0.046 (0.216)	0.033 (0.287)	0.126 (0.155)	0.158 (0.196)	-0.469 (0.436)	-0.353 (0.673)	-0.053 (0.585)	0.065 (0.455)
ΔQ_2	1.390 (0.535)	0.915 (0.792)	1.381 (0.788)	0.512 (1.183)	0.946 (0.378)	0.784 (0.590)	0.393 (0.685)	3.168 (0.845)
ΔQ_3	8.308 (0.650)	7.881 (0.984)	7.862 (0.961)	6.893 (1.485)	8.580 (0.438)	8.723 (0.677)	8.886 (0.742)	12.984 (1.041)
ΔQ_4	6.377 (0.536)	6.744 (0.792)	5.606 (0.784)	5.634 (1.733)	7.412 (0.374)	8.393 (0.581)	8.910 (0.682)	14.399 (0.841)
ρ_1	-0.158 (0.109)	-0.151 (0.107)	-0.338 (0.139)	-0.332 (0.136)	0.085 (0.188)	0.054 (0.199)	-0.102 (0.160)	-0.058 (0.151)
ρ_2	-0.138 (0.110)	-0.170 (0.108)	-0.343 (0.148)	-0.387 (0.145)	0.173 (0.175)	0.165 (0.183)	0.125 (0.149)	-0.084 (0.146)
ρ_3	0.044 (0.110)	0.022 (0.108)	-0.201 (0.148)	-0.241 (0.146)	0.329 (0.177)	0.308 (0.184)	0.226 (0.151)	0.111 (0.144)
ρ_4	0.165 (0.108)	0.220 (0.105)	0.092 (0.136)	0.125 (0.133)	-0.385 (0.198)	-0.251 (0.205)	0.019 (0.162)	0.223 (0.154)
SER	2.20	2.90	2.07	2.96	3.35	2.72	2.96	2.55
R^2	0.841	0.756	0.849	0.758	0.916	0.865	0.841	0.938
$p(\mathrm{H}_0)$	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000
$p(\mathrm{H}_1)$	0.005	0.004	0.219	0.119	0.000	0.000	0.000	0.000
p(stability)	0.097	0.282						
Seas. unit root t-value	-6.13	-5.84	-5.60	-5.15	-4.17	-3.77	-4.90	-4.49
5% crit.value	-4.11	-4.11	-4.14	-4.14	-4.21	-4.21	-4.21	-4.21

Table 7: **Comparison of the Quarterly Seasonal Patterns in Actual Retail Sales, at Current Prices, for Norway and Sweden**

	Norway	Sweden	Norway – Sweden	Norway	Sweden	Norway – Sweden	Norway	Sweden	Norway – Sweden
	1973:4–1993:1	1973:4–1993:1	1973:4–1993:1	1981:2–1993:1	1981:2–1993:1	1981:2–1993:1	1985:2–1993:1	1985:2–1993:1	1985:2–1993:1
Const.	1.794 (0.401)	1.716 (0.706	-0.157 (0.254)	1.316 (0.402)	0.666 (1.976)	-0.208 (0.418)	1.096 (0.453)	0.111 (2.733)	-0.076 (0.610)
ΔQ_2	7.260 (1.213)	9.681 (1.124)	-2.425 (1.297)	7.133 (1.795)	10.358 (1.093)	-3.572 (2.050)	7.139 (2.596)	11.015 (1.081)	-4.147 (3.392)
ΔQ_3	9.344 (1.204)	5.451 (1.427)	3.973 (1.337)	8.559 (1.475)	5.355 (1.278)	2.664 (1.622)	8.602 (1.840)	5.695 (1.050)	2.181 (2.450)
ΔQ_4	27.305 (1.191)	21.941 (1.108)	5.323 (1.275)	25.956 (1.782)	20.317 (1.085)	5.212 (2.047)	24.897 (2.582)	19.645 (1.074)	4.591 (3.393)
ρ_1	-0.177 (0.107)	-0.020 (0.108)	-0.311 (0.114)	-0.206 (0.148)	-0.035 (0.150)	-0.306 (0.149)	-0.202 (0.201)	-0.127 (0.220)	-0.285 (0.193)
ρ_2	-0.039 (0.111)	-0.063 (0.114)	-0.184 (0.120)	-0.048 (0.148)	0.044 (0.153)	-0.076 (0.156)	-0.065 (0.195)	0.193 (0.217)	-0.070 (0.198)
ρ_3	-0.027 (0.111)	0.092 (0.113)	-0.150 (0.119)	-0.161 (0.142)	0.189 (0.147)	-0.106 (0.156)	-0.271 (0.193)	0.289 (0.218)	-0.152 (0.199)
ρ_4	0.477 (0.108)	0.566 (0.114)	0.348 (0.113)	0.408 (0.141)	0.515 (0.147)	0.394 (0.148)	0.353 (0.196)	0.335 (0.225)	0.401 (0.195)
SER	2.69	2.41	2.90	2.73	2.48	3.16	2.98	2.79	3.75
R^2	0.978	0.977	0.751	0.977	0.975	0.757	0.974	0.970	0.729
$p(\mathrm{H}_0)$			0.000			0.000			0.012
$p(\tilde{\mathrm{H}}_0)$			0.000			0.040			0.379
$p(\mathrm{H}_1)$			0.000			0.001			0.031
p(stability)	0.005	0.012	0.146						
Seas. unit root t-value	-4.60	-2.85	-4.57	-3.75	-2.75	-3.84	-4.00	-2.52	-2.95
5% crit.val.	-4.11	-4.11	-4.11	-4.21	-4.21	-4.21	-4.21	-4.21	-4.21

Table 8: **Comparison of the Monthly Seasonal Patterns in Actual Retail Sales, at Current Prices, for Norway and Sweden.**

	Norway 1973:8–1993:2	Sweden 1973:8–1993:2	Norway – Sweden 1973:8–1993:2	Norway 1981:4–1993:2	Sweden 1981:4–1993:2	Norway – Sweden 1981:4–1993:2	Norway 1985:4–1993:2	Sweden 1985:4–1993:2	Norway – Sweden 1985:4–1993:2
Constant	0.604	0.670	-0.066	0.450	0.546	-0.107	0.368	0.439	-0.083
	(0.123)	(0.112)	(0.132)	(0.165)	(0.120)	(0.178)	(0.229)	(0.141)	(0.272)
$\Delta M_2 - \Delta M_1$	-4.931	-5.529	0.636	-4.831	-5.821	1.089	-5.293	-6.084	1.066
	(0.723)	(0.737)	(0.717)	(0.996)	(1.038)	(0.918)	(1.195)	(1.264)	(1.448)
$\Delta M_3 - \Delta M_1$	2.788	5.897	-3.189	2.968	7.824	-4.977	2.942	7.713	-4.989
	(0.672)	(0.684)	(0.716)	(0.924)	(0.902)	(0.913)	(1.197)	(1.137)	(1.382)
ΔM_4	3.266	8.449	-5.102	2.953	7.696	-4.797	3.792	8.479	-4.277
	(0.901)	(0.904)	(1.009)	(1.186)	(1.194)	(1.297)	(1.479)	(1.466)	(1.942)
ΔM_5	8.384	11.177	-2.808	7.405	12.037	-4.540	8.524	13.522	-5.095
	(0.909)	(0.898)	(1.010)	(1.197)	(1.121)	(1.295)	(1.533)	(1.354)	(1.919)
ΔM_6	10.749	8.991	1.216	9.999	10.587	-0.645	9.578	11.115	-1.487
	(0.941)	(0.923)	(1.040)	(1.240)	(1.159)	(1.364)	(1.599)	(1.396)	(2.069)
ΔM_7	6.943	4.054	2.926	5.045	3.772	1.174	4.693	3.745	1.259
	(0.995)	(0.974)	(1.070)	(1.321)	(1.209)	(1.374)	(1.694)	(1.453)	(2.069)
ΔM_8	11.439	6.177	5.318	10.821	7.208	3.744	11.111	7.865	3.164
	(0.959)	(0.937)	(1.077)	(1.243)	(1.185)	(1.419)	(1.658)	(1.433)	(2.143)
ΔM_9	9.945	5.043	4.864	9.292	4.946	4.250	9.003	5.583	3.699
	(0.981)	(0.959)	(1.054)	(1.317)	(1.205)	(1.374)	(1.686)	(1.451)	(2.066)
ΔM_{10}	16.880	12.604	4.302	16.229	10.819	5.401	15.696	10.746	4.946
	(0.927)	(0.910)	(1.022)	(1.243)	(1.161)	(1.365)	(1.599)	(1.401)	(2.071)
ΔM_{11}	14.941	10.429	4.567	14.224	9.792	4.489	13.713	8.517	5.226
	(0.894)	(0.882)	(0.990)	(1.198)	(1.118)	(1.294)	(1.537)	(1.358)	(1.924)
ΔM_{12}	46.854	39.290	7.512	44.546	37.779	6.752	43.036	36.857	6.540
	(0.884)	(0.884)	(0.655)	(1.185)	(1.185)	(1.293)	(1.490)	(1.451)	(1.945)
ρ_1	-0.644	-0.653	-0.531	-0.680	-0.723	-0.484	-0.642	-0.769	-0.491
	(0.065)	(0.065)	(0.066)	(0.086)	(0.080)	(0.086)	(0.112)	(0.108)	(0.108)
ρ_2	-0.313	-0.358	-0.304	-0.319	-0.422	-0.223	-0.311	-0.486	-0.182
	(0.075)	(0.074)	(0.072)	(0.100)	(0.093)	(0.093)	(0.128)	(0.128)	(0.120)
ρ_3	0.056	0.016	-0.107	0.102	-0.083	-0.129	0.149	-0.112	-0.106
	(0.065)	(0.064)	(0.065)	(0.087)	(0.080)	(0.083)	(0.115)	(0.109)	(0.108)
ρ_{12}	0.201	0.240	0.235	0.187	0.288	0.247	0.119	0.249	0.272
	(0.054)	(0.051)	(0.057)	(0.071)	(0.063)	(0.072)	(0.101)	(0.087)	(0.098)
SER	3.21	3.02	3.45	3.38	2.79	3.39	3.76	2.91	3.99
R^2	0.967	0.965	0.524	0.962	0.971	0.519	0.953	0.969	0.476
$p(\mathrm{H}_0)$			0.000			0.000			0.000
$p(\bar{\mathrm{H}}_0)$			0.000			0.113			0.004
p(stability)	0.013	0.000	0.024						
Seas. unit root t-value	-7.94	-8.02	-6.72	-7.34	-6.27	-4.88	-6.11	-4.96	-3.97
5% crit.val.	-5.82	-5.82	-5.82	-5.86	-5.86	-5.86	-5.86	-5.86	-5.86

Table 9: **Comparison of the Seasonal Patterns in Norwegian Retail Sales and Consumption, Current Prices**

All samples 1981:2–1993:1

	Total Retail Sales	Actual Retail Sales	Goods Consumption	Goods Consumption Excl. Energy	Goods Consumption Excl. Energy and Vehicles	Total Retail Sales – Goods Consumption Excl. Energy	Actual Retail Sales – Goods Consumption Excl. Energy and Vehicles
Constant	1.163 (0.797)	1.316 (0.402)	1.486 (0.742)	1.494 (0.732)	1.575 (0.628)	-0.234 (0.134)	-0.371 (0.272)
ΔQ_2	7.660 (1.258)	7.133 (1.795)	2.266 (0.896)	8.162 (0.942)	7.964 (1.198)	-0.336 (0.606)	-0.813 (0.968)
ΔQ_3	6.980 (1.018)	8.559 (1.475)	-0.133 (0.942)	7.134 (0.925)	8.094 (1.325)	-0.081 (0.566)	0.445 (0.914)
ΔQ_4	20.514 (1.254)	25.956 (1.782)	16.666 (0.885)	20.208 (0.934)	24.106 (1.181)	0.226 (0.608)	1.787 (0.968)
ρ_1	-0.013 (0.154)	-0.206 (0.148)	-0.056 (0.151)	-0.068 (0.151)	-0.058 (0.137)	-0.552 (0.153)	-0.251 (0.132)
ρ_2	0.170 (0.148)	-0.048 (0.148)	0.103 (0.146)	0.141 (0.148)	0.010 (0.137)	-0.250 (0.173)	-0.059 (0.138)
ρ_3	-0.056 (0.145)	-0.161 (0.142)	0.116 (0.142)	0.008 (0.144)	0.012 (0.133)	0.009 (0.180)	-0.015 (0.139)
ρ_4	0.364 (0.145)	0.408 (0.141)	0.308 (0.143)	0.275 (0.143)	0.443 (0.131)	0.255 (0.158)	0.475 (0.136)
SER	2.82	2.73	2.63	2.79	2.51	1.43	1.59
R^2	0.962	0.977	0.962	0.962	0.978	0.360	0.558
$p(H_0)$						0.802	0.052
Seas. unit root t-value	-3.28	-3.75	-4.37	-4.24	-3.88	-4.11	-3.53
5% crit.val.	-4.21	-4.21	-4.21	-4.21	-4.21	-4.21	-4.21

Table 10: Comparison of the Seasonal Patterns in Swedish Retail Sales and Consumption, Current Prices

All samples 1981:2–1993:1

	Actual Retail Sales	Goods Consumption	Goods Consumption Excluding Energy	Goods Consumption Excluding Energy and Vehicles	Actual Retail Sales – Goods Consumption Excluding Energy and Vehicles
Constant	0.666 (1.976)	1.520 (0.595)	1.416 (0.777)	1.235 (1.022)	0.004 (0.274)
ΔQ_2	10.358 (1.093)	1.940 (0.544)	7.568 (0.665)	4.440 (0.544)	6.194 (1.127)
ΔQ_3	5.355 (1.278)	-8.558 (0.676)	-1.907 (0.828)	-5.037 (0.617)	11.474 (1.234)
ΔQ_4	20.317 (1.085)	8.855 (0.542)	11.292 (0.660)	9.887 (0.541)	11.392 (1.234)
ρ_1	-0.035 (0.150)	0.026 (0.156)	0.042 (0.154)	-0.014 (0.152)	-0.709 (0.171)
ρ_2	0.044 (0.153)	0.001 (0.153)	0.010 (0.155)	0.135 (0.152)	-0.688 (0.170)
ρ_3	0.189 (0.147)	0.303 (0.148)	0.254 (0.148)	0.290 (0.149)	-0.634 (0.164)
ρ_4	0.515 (0.147)	0.118 (0.155)	0.242 (0.158)	0.268 (0.161)	-0.008 (0.160)
SER	2.48	2.16	2.26	1.95	1.55
R^2	0.975	0.970	0.965	0.972	0.956
$p(H_0)$					0.000
Seas. unit root t-value	-2.75	-3.97	-2.71	-3.35	-3.58
5% crit.val.	-4.21	-4.21	-4.21	-4.21	-4.21

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