UNIVERSITEIT GENT

FACULTEIT ECONOMIE
EN BEDRIJFSKUNDE

HOVENIERSBERG 24
B-9000 GENT
Tel. : 32 - (0)9 – 264.34.61
Fax. : 32 - (0)9 – 264.35.92

WORKING PAPER

Determinants of consumption smoothing

Gert Peersman [1]

Lorenzo Pozzi [2]

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[1] Department of Financial Economics, Ghent University, Gert.Peersman@ugent.be, http://fetew.rug.ac.be/Fineco/gert.htm

[2] Department of Social Economics and SHERPPA, Ghent University, Lorenzo.Pozzi@ugent.be , http://www.sherppa.be

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Determinants of consumption smoothing*

Gert Peersman† Lorenzo Pozzi‡

Abstract

As is generally acknowledged, the failure of the perfect credit markets assumption underlying the permanent income hypothesis may be responsible for low consumption smoothing and observed excess sensitivity of consumption to current income. The economic literature puts forward a number of potential determinants of liquidity constraints. In this paper we investigate the relevance of these determinants by looking at their impact on household consumption smoothing. Applying a Kalman filter to a state-space version of our model, we find that excess sensitivity in the US is higher in recessions and depends positively on government debt and negatively on financial liberalization.

JEL classification: E21

Keywords: Private consumption, liquidity constraints, consumption smoothing, excess sensitivity, Kalman filter, state-space models.

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†Department of Financial Economics, Ghent University, Gert.Peersman@ugent.be, http://fetew.rug.ac.be/Fineco/gert.html

‡Department of Social Economics and SHERPPA, Ghent University, Lorenzo.Pozzi@ugent.be, http://www.sherppa.be

1 Introduction

Under very strict assumptions, the permanent income hypothesis implies that aggregate private consumption follows a random walk (see Hall 1978). Maximizing forward-looking consumers lend and borrow freely on perfect capital markets to smooth consumption over time. In reality, however, private consumption is found to be excessively sensitive to current income. This observed excess sensitivity can be explained theoretically by dropping some of Hall's assumptions. Liquidity constraints are generally acknowledged to be the most probable reason (Flavin 1985, Jappelli and Pagano 1989, Campbell and Mankiw 1990). In the majority of the empirical studies on excess sensitivity the degree of smoothing and thus the degree of liquidity constraints is considered fixed over time and across countries. The papers that allow for endogenous excess sensitivity usually attribute differences in excess sensitivity across countries and over time to the development of credit markets and financial liberalization (see Campbell and Mankiw 1991, Bacchetta and Gerlach 1997). However, a number of papers have considered other determinants of excess sensitivity. For instance, McKiernan (1996) considers whether the change in the unemployment rate and a credit crunch dummy can explain the time variation in excess sensitivity in the US. Evans and Karras (1998) investigate whether differences in the savings rate explain cross-country differences in excess sensitivity. Pozzi, Heylen and Dossche (2004) look whether the government debt ratio affects excess sensitivity in a panel of OECD countries.

In this paper, we simultaneously consider a number of determinants of excess sensitivity or consumption smoothing in the US. Some of these have not been considered before in the literature on excess sensitivity, but have been put forward in the theoretical literature on liquidity constraints (see Stiglitz and Greenwald 2003 among others). We thus link the literature on liquidity constraints to the literature on excess sensitivity. This allows for a more systematic investigation of the determinants of consumption smoothing than is usually encountered in the excess sensitivity literature. Besides financial liberalization, the traditional determinant in the excess sensitivity literature, we also consider unemploy-

ment, the wealth ratio, the government debt ratio and, most importantly, the state of the business cycle. To the best of our knowledge, no other paper has yet analyzed the influence of the wealth ratio[1] and the business cycle on the ability of consumers to smooth consumption. With respect to the business cycle, there is a literature that claims that liquidity constraints are more severe in recessions than in booms (see Stiglitz and Weiss 1981, Bernanke and Gertler 1989). This raises the possibility of a so-called financial accelerator effect, i.e. the possibility that an initial shock in the economy is further propagated and amplified because the shock increases liquidity constraints by households and/or firms. Kashyap, Lamont and Stein (1994), Gertler and Gilchrist (1994) and Vermeulen (2000) find that small firms are more liquidity constrained during downturns. This negatively affects investment. Peersman and Smets (2002) find macroeconomic evidence of a role of a financial accelerator effect in the economy during recessions at the industry level. All these papers, however, investigate the role of the business cycle on liquidity constraints of firms. In this paper the focus is completely different. We look at the effects of the business cycle on liquidity constraints of households.

Methodologically, we estimate a system (one-step approach) consisting of a consumption growth equation and a multivariate process for the excess sensitivity or consumption smoothing parameter. This approach is different from the methods applied until now in the existing excess sensitivity literature. In that literature, if a multivariate process for the excess sensitivity parameter is considered, either a two-step approach is used (Mc Kiernan 1996) or the excess sensitivity parameter is, rather restrictively, assumed to be a deterministic function of the variables considered (see e.g. Evans and Karras 1998 ; Pozzi, Heylen and Dossche 2004).

Our results using quarterly data over the period 1970-2000 suggest that consumption smoothing by US consumers is lower in recessions and when the government debt is high. The first result is new, the second result seems to confirm earlier evidence by Pozzi, Heylen and Dossche (2004) who find a strong positive effect of government debt on the

[1] Even though the idea of wealth as collateral to make access to credit easier is the same as in Evans and Karras (1998). Because of data availability they use the saving rate instead of the wealth ratio.

excess sensitivity of private consumption in a panel of OECD countries. Furthermore, we find that the level of consumption smoothing is significantly higher from the 1980s onward, which probably reflects financial liberalization. We find no significant effect of the wealth ratio or the unemployment rate on the degree of smoothing. Our results thus seem to be in accordance with some of the predictions of the literature on liquidity constraints.

The rest of the paper is structured as follows. In section 2 we discuss the theoretical foundations of the framework that we use. In section 3.1 we discuss our empirical specification and we put it in state-space form. In section 3.2 the methodology is discussed (estimation of a state-space model, application of the Kalman filter and use of a bootstrap to obtain standard errors) and in section 3.3 data problems are tackled. In section 4.1 we present the basic results and in section 4.2 we discuss the robustness of these results. Finally, in section 5 we conclude and we discuss some policy implications of our results.

2 Theoretical framework

We start from the standard random walk result for private consumption (see Hall 1978). If consumers can borrow and lend at the same risk-free interest rate (which is assumed to be constant and equal to the subjective rate of time preference), if instantaneous utility is of the constant relative risk aversion type and if the level of (real per capita) private consumption is log-normally distributed, we can write the first-order condition for an optimizing representative consumer as

$$\Delta c_t = \alpha + \varepsilon_t \tag{1}$$

where Δc_t is the growth rate of real per capita consumption and ε_t is an innovation that is uncorrelated with lagged variables. Private consumption growth cannot be forecasted.

A large literature has demonstrated, however, that private consumption growth is not unpredictable. Typically it is found that there is excess sensitivity of private consumption to anticipated disposable income. Thus reality may be better approximated by

$$\Delta c_t = \alpha + \beta E_{t-1} \Delta y_t + \varepsilon_t \tag{2}$$

where Δy_t is the growth rate of real per capita disposable income, E_{t-1} is the expectations operator conditional on information available at time $t-1$ and β is the excess sensitivity parameter ($0 \leq \beta \leq 1$). Thus the lower β, the higher is consumption smoothing by households. The most common interpretation for $\beta > 0$ is that a proportion of the population consumes its current income because it is liquidity constrained (see Campbell and Mankiw 1990). A less common interpretation for $\beta > 0$ refers to uncertain consumers with a precautionary savings motive. Carroll (1994) shows that US consumption depends on current income rather than lifetime income. He also shows that uncertainty about future income is an important determinant of private consumption. He reconciles these findings with the existence of "buffer stock" consumers. These consumers keep a small stock of assets as an insurance against future income fluctuations while their consumption tracks disposable income quite closely. In this paper our main emphasis is on the first interpretation of excess sensitivity, though we interpret our results along the lines of the second possibility as well.

Whatever its interpretation, it is unlikely that β will be constant. A number of studies consider the degree of excess sensitivity (or the degree of smoothing) to be endogenous by assuming that β varies over time and/or across countries (see Campbell and Mankiw 1991, Bacchetta and Gerlach 1997, McKiernan 1996, Evans and Karras 1998, Pozzi, Heylen and Dossche 2004). In the present paper we also consider a time-varying β. Specifically, we introduce a number of determinants of β that have not been considered yet in the literature on excess sensitivity and that are consistent with the interpretation of β as stemming from the occurrence of liquidity constraints. In the next section we discuss our specification for consumption growth with time-varying excess sensitivity.

3 Empirical specification and estimation issues

3.1 The model and its state-space representation

We consider the following specification,

$$\Delta c_t = \alpha_t + \beta_t E_{t-1}\Delta y_t + \varepsilon_t + \theta\varepsilon_{t-1} \tag{3}$$

$$\alpha_t = \alpha_0 + \alpha_{t-1} + \zeta_t \tag{4}$$

$$\beta_t = \beta_0 + \beta_1 fl_t + \beta_2 bc_{t-2} + \beta_3 u_{t-2} + \beta_4 w_{t-2} + \beta_5 b_{t-2} + \mu_t \tag{5}$$

where eq.(3) is the consumption growth equation extended to allow for time-varying coefficients (see McKiernan 1996, Bachetta and Gerlach 1997). We allow for an MA(1) error in consumption growth ($-1 \leq \theta \leq 1$) which may be due to time aggregation (Working 1960), the presence of durable components in our consumption measure (Mankiw 1982) or a transitory component in the log of consumption. Following Bachetta and Gerlach (1997) we specify α_t as a random walk with drift α_0 in eq.(4).

Eq.(5) is our basic specification for the time-varying excess sensitivity parameter β_t. We model β_t as a straightforward linear function[2] of financial liberalization (fl_t), a variable reflecting the state of the business cycle (bc_{t-2}), the unemployment rate (u_{t-2}), the wealth ratio (w_{t-2}) and the government debt ratio (b_{t-2}). We expect $\beta_1 < 0$ since financial liberalization should make access to credit and thus consumption smoothing possibilities easier for consumers (see Campbell and Mankiw 1991, Bacchetta and Gerlach 1997). Further, if bc_t is defined so that it takes on higher values in recessions, we expect $\beta_2 > 0$. Recessions may lead to more liquidity constraints.[3] First, the deterioration of households' balance sheets in a recession decreases "internal" finance possibilities (i.e. through income

[2] One drawback of the specification in eq.(5) is that it does not restrict β_t to lie between 0 and 1. We have experimented with logistic transformations of β_t as excess sensitivity parameters. Using an approximate non-linear Kalman filter we obtained plausible parameter estimates and state projections, though the bootstrapped standard errors on the parameters were extraordinary large to draw any conclusions.

[3] Implying that fiscal policy or monetary policy measures also affect liquidity constraints in so far as they affect the state of the business cycle.

or accumulated wealth) raising demand for "external" finance. Second, higher monitoring and contract enforcement costs and information asymmetries may increase the risk for banks of giving loans in bad times and diminish the supply of credit. These factors may lead to a higher "external finance premium", i.e. the difference between the cost of "external" and "internal" finance. Only potential borrowers willing to pay the higher price will receive a loan. Thus, as also noted by Jappelli and Pagano (1989), a high "external finance premium" may be the source of liquidity constraints and excess sensitivity.[4] Also, non-price credit rationing is possible. A high default risk of borrowers in bad times could lead to a reduction of lending by banks to some potential borrowers even if these are willing to pay a higher price to obtain the loan, i.e. equilibrium credit rationing (Stiglitz and Weiss, 1981). Higher liquidity constraints that are the result of these factors may propagate and amplify the initial downturn, i.e. the so-called financial accelerator effect. Further, we expect $\beta_3 > 0$ since the unemployed usually have no access to credit. More unemployment can thus be expected to raise liquidity constraints and excess sensitivity.[5] Given that the state of the business cycle is included in eq.(5), the variable u_t captures the more structural part of unemployment. Then, we expect $\beta_4 < 0$ since wealthier households have more collateral and should thus have easier access to credit. Finally, we expect $\beta_5 > 0$. If the government's demand for the available savings in the economy rises, banks may prefer to give loans to the government rather than to the private sector because the government is a more reliable debtor (crowding out effects, see also Stiglitz and Greenwald 2003). Another possibility is that an increasing government debt implies higher future tax liabilities for households so that, due to their increased risk of default, banks may give them less credit (see Yotsuzuka 1987).[6]

[4] Note that the possibility of a positive external finance premium (e.g. a wedge between lending and deposit rates) is a deviation from the standard permanent income hypothesis. The latter theorem is based on the assumption that the same interest rate applies to both lenders and borrowers.

[5] Note that in recessions some people still employed may have a high risk of becoming unemployed in the near future. This is captured by the business cycle variable included in eq.(5).

[6] Of course if government debt is used to finance tax reductions, constrained consumers will experience a rise in disposable income and thus a loosening of their constraint (Aiyagari and McGrattan 1998). Thus the net effect of government debt on liquidity constraints may be unclear a priori.

It is possible that the excess sensitivity of consumption is caused, not only by liquidity constraints, but also by consumer uncertainty about the future and precautionary savings (see Carroll 1994). The expected signs of the coefficients on cycle, unemployment, wealth and government debt are identical under this less common interpretation of excess sensitivity. Uncertain "buffer stock" consumers give a large weight to current income in their consumption decisions. Recessions, high government debt and unemployment may raise precaution and excess sensitivity. This alternative interpretation of excess sensitivity will be taken into account when interpreting our results (see section 4).

Note that all variables in eq.(5) are lagged twice to make sure they are predetermined in the system considered (given the MA(1) error in eq.(3)). The exception is fl_t for which we use a variable which reflects a structural regulatory change (see below) and which can, as a consequence, be considered exogenous. Note that lagging the determinants of β_t can also be given an economic motivation. It is likely that macro-economic aggregates enter the information sets of financial institutions and consumers with some delay (see Campbell and Mankiw 1990). Finally, note that the errors ε_t, ζ_t and μ_t are assumed to be independent Gaussian white noise terms.[7]

A state-space representation of eqs.(3)-(5) with state vector $\mathbf{S}_t$ is given by

$$\Delta c_t = \mathbf{H}_t' \mathbf{S}_t \tag{6}$$

$$\mathbf{S}_t = \mathbf{F}\mathbf{S}_{t-1} + \mathbf{D}\mathbf{Z}_t + \mathbf{v}_t \tag{7}$$

$$\text{where } \mathbf{H}_t = \begin{bmatrix} 1 \\ E_{t-1}\Delta y_t \\ 1 \\ \theta \end{bmatrix}, \mathbf{D} = \begin{bmatrix} \alpha_0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\ \beta_0 & \beta_1 & \beta_2 & \beta_3 & \beta_4 & \beta_5 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \end{bmatrix}, \mathbf{S}_t = \begin{bmatrix} \alpha_t \\ \beta_t \\ \varepsilon_t \\ \varepsilon_{t-1} \end{bmatrix}$$

[7] Estimations that allow for a non-zero correlation between ζ_t and μ_t give very similar results to the ones reported.

$$\mathbf{F} = \begin{bmatrix} 1 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 \\ 0 & 0 & 1 & 0 \end{bmatrix}, \mathbf{Z}_t' = \begin{bmatrix} 1 & fl_t & bc_{t-2} & u_{t-2} & w_{t-2} & b_{t-2} & 0 & 0 \end{bmatrix} \text{ and } \mathbf{v}_t = \begin{bmatrix} \zeta_t \\ \mu_t \\ \varepsilon_t \\ 0 \end{bmatrix}$$

and where $\mathbf{v}_t \sim N(\mathbf{O}, \mathbf{Q})$ with $\mathbf{Q} = E(\mathbf{v}_t \mathbf{v}_t') = \begin{bmatrix} \sigma_\zeta^2 & 0 & 0 & 0 \\ 0 & \sigma_\mu^2 & 0 & 0 \\ 0 & 0 & \sigma_\varepsilon^2 & 0 \\ 0 & 0 & 0 & 0 \end{bmatrix}$

3.2 Methodology

Before estimating the system (6)-(7), we construct the anticipated disposable income growth series $E_{t-1}\Delta y_t$. We construct it as the fitted values of a regression of disposable income growth on a number of "instruments" suggested in the literature.[8] As "instruments" to forecast disposable income, we use a constant, lagged consumption growth (see Campbell and Mankiw 1990), lagged disposable income growth, lagged personal wealth growth, lagged changes in the nominal interest rate (see Campbell and Mankiw 1990 and McKiernan 1996), the current and lagged growth rate of government expenditures (consumption and investment; see Blanchard and Perotti 2002) and a lagged error correction term, i.e. log consumption minus log disposable income (see Campbell 1987). We use lags 2 to 5 for all instruments[9] except for the error correction term (only lag 2) and for the lagged growth rate of government expenditures (lags 0 to 5). The reason for starting with lag 2 in the construction of $E_{t-1}\Delta y_t$ is the presence of an MA(1) term in eq.(3). For $E_{t-1}\Delta y_t$ to be predetermined (see below) in the system (6)-(7), variables must be lagged

[8] A drawback of this procedure is that the uncertainty about the estimated parameters in this forecasting equation is not taken into account when estimating the system (6)-(7). An alternative approach would be to add an income growth equation to the system as a second observation equation. This, however, would imply the use of a non-linear Kalman filter which is approximate and thus sub-optimal. From a practical point of view it would imply an important increase in the number of parameters to estimate which could considerably complicate the maximum likelihood procedure (i.e. starting values, ...).

[9] A sixth lag is added as a robustness test, see section 4.2

at least twice. Since government expenditures will most likely not be affected by economic activity in the same quarter, we do include lags 0 and 1 for government expenditures.

Given that the variables in $\mathbf{H}_t$ and $\mathbf{Z}_t$ are either exogenous or predetermined, the Kalman filter equations (see for instance Hamilton 1994) can be applied to the system (6)-(7). To initialize the filter we use a diffuse prior, i.e. we assume that the initial state vector $\mathbf{S}_0$ is random with covariance matrix $\kappa\mathbf{I}$ where $\kappa \to \infty$ and where $\mathbf{I}$ is an identity matrix. We use the Kalman filter to construct the sample log likelihood function which is then maximized numerically with respect to the unknown parameters in $\mathbf{H}_t$, $\mathbf{D}$ and $\mathbf{Q}$. This procedure provides the state projections $\{\mathbf{S}_{t|t}\}_{t=1}^{T}$ and the maximum likelihood estimates of the parameters in $\mathbf{H}_t$, $\mathbf{D}$ and $\mathbf{Q}$.[10]

Note that to take into account a potential small sample bias we bootstrap the standard errors of the state projections (thus replacing the usual standard errors obtained through the standard Kalman formulas) and the standard errors of the maximum likelihood estimates (thus replacing the usual standard errors obtained from the information matrix). The bootstrap procedure is explained in appendix A.

3.3 Data issues

We use quarterly data for the US over the period 1969:03-1999:02 (this actual sample period is determined by the availability of data and does not include the observations lost due to the use of lags). Data sources are reported in appendix B. Data are seasonally adjusted where necessary. When considered in real terms, appropriate deflators for the data are used. For Δc_t we use the growth rate of real per capita expenditures on non-durables and services. This is standard since Hall (1978). To construct $E_{t-1}\Delta y_t$ we use the fitted values of a regression of the growth rate of real per capita household disposable income Δy_t on the variables mentioned in section 3.2: lagged real per capita non-durable consumption growth, lagged real per capita disposable income growth, lagged real per capita personal

[10] Note the estimation procedure restricts the variances in $\mathbf{Q}$ to be strictly positive and the parameter θ to lie in the interval $[-1, 1]$.

wealth growth, lagged nominal change in the 3 month T-bill rate, current and lagged real per capita growth rate of government expenditures and log real per capita non-durable consumption minus log real per capita disposable income.[11] Quarterly population figures are obtained through interpolation of annual figures, assuming that quarterly population growth equals annual population growth divided by four. For u_t, w_t and b_t we respectively use the unemployment rate (in percent), total personal wealth to GDP ratio (in percent; GDP on annual basis) and the gross government debt to GDP ratio (in percent; GDP on annual basis). For bc_t we use both the readily available NBER recession dummy and a constructed measure. The NBER dummy equals 1 during recessions and 0 otherwise. As an alternative, we use a constructed variable for cycle, which is obtained through a two-state Markov Switching model applied to real GDP growth that endogenously determines the probability of being in a state of recession. This variable takes on values between 0 and 1.[12] The filtered probabilities of being in a recession, together with the NBER defined turning points of the business cycle are shown in figure 1. These probabilities closely resemble the NBER dating of the cycles. For fl_t existing work typically uses the stock of outstanding consumer credit (e.g. Bachetta and Gerlach, 1997). For our purpose this variable is inadequate because, besides the fact that it is driven by both credit supply and demand, more importantly, it may also be highly endogenous to other variables included in eq.(5). For instance, as we have argued before, one of the reasons why the government debt ratio affects the excess sensitivity of private consumption to disposable income may be that it makes banks less willing to lend. We therefore prefer a variable that has a more exogenous nature.[13] We have chosen a simple measure for fl_t which can be considered exogenous: we capture financial liberalization through a level shift in β_t that occurs in the first quarter of 1982 (i.e. a dummy that takes on the value 0 before 1982:1 and 1 from 1982:01 onward). This date is qualified by Kaminsky and Schmukler (2003) as the point

[11] This regression has an R^2 of 0.36 and an R^2adj. of 0.21.

[12] The estimated probabilities of being in a recession are obtained using a univariate two-state Markov-switching model with four lags, as in Hamilton (1994). Full results are available upon request.

[13] Other possible proxies, like for instance the trend in the deposit minus lending rates of banks, M2 on M1 or M2 on GDP, were also considered but their decreasing trends in the eighties strongly contradict the general presumption that liberalization increased during that decade.

in time where the domestic financial sector in the US can be considered "fully liberalized" (to be interpreted as the date on which regulations like credit allocation control were fully lifted). Note that for fl_t and w_t we consider other possibilities in the robustness tests presented in section 4.2.

4 Results

4.1 Basic results

In table 1 we present the basic results from the estimation of the system eqs. (6) and (7) over the period 1969:03-1999:02. Note that we do not report $\widehat{\sigma}^2_\zeta, \widehat{\sigma}^2_\mu$ and $\widehat{\sigma}^2_\varepsilon$ and $\widehat{\theta}$. First, in column 1, we report the results of estimating the state-space model (6) and (7) under the restrictions that α_t and β_t are constant.[14] We find a value for the excess sensitivity parameter over the sample period of about 0.27 (significant at the 5 percent level). This value is close to the values of about 0.3 found by Bacchetta and Gerlach (1997) for the US over the period 1970-1995. The question is then whether this value hides important time-variation. In column 2 of table 1 we report the results of estimating (6) and (7) unrestricted with the Markov-switching probabilities for bc_t (the higher these probabilities, the stronger the recession). In column 3 we use the NBER dummy for bc_t instead (equals one in recessions). The 1982 liberalization dummy is used for fl_t in both cases. First, we find that the excess sensitivity level is significantly lower from 1982 onward (with a difference in the average level of 0.55/0.59 between the 1970s and the 1980/1990s). This result seems to indicate that liberalization has a very significant impact on consumer smoothing possibilities. This result contradicts Campbell and Mankiw (1991) who do not find a significant level shift in excess sensitivity in 1980 in the US. They attribute their result to potential offsetting factors (like unemployment) which they do not take into account and to the fact that their 2SLS estimation procedure may not

[14] Following Bacchetta and Gerlach (1997) we have also estimated the system under the restriction that both α_t and β_t follow independent random walks (results available upon request). We find that β_t tends to decline over time.

be powerful enough to detect movements over time in excess sensitivity. To allow for a less drastic shift in excess sensitivity, in section 4.2, we use a linear spline for fl_t: deterministically increasing over time until 1982 and constant from 1982 onward. Next, in column 2, we find a significant positive impact of bc_t on excess sensitivity. This seems to suggest that consumption smoothing is lower in recessions. We do not find a statistically significant confirmation of this in column 3, however, with the NBER dummy (t-value for the coefficient on bc_t is only 1.47). Possibly, the correlation of the business cycle with other variables included in (5), like the wealth rate or the unemployment rate, is responsible for this.[15] We return to this issue in section 4.2. Further, in both columns 2 and 3, there is no evidence of a significant effect of the unemployment rate nor the wealth rate, even though the signs of the coefficients on these variables are as expected and the magnitudes of the coefficients on these variables seem plausible. This result implies that there is no structural effect on liquidity constraints of collateral in hands of consumers and no structural effect of unemployment either. Of course, both wealth/collateral and unemployment may play a role in the cyclical behavior of β_t since the role of the business cycle in liquidity constraints cannot be dismissed. Finally, in both columns 2 and 3, we find a positive effect of government debt on excess sensitivity, confirming the results by Pozzi, Heylen and Dossche (2004). This result suggests that government debt makes consumers more liquidity constrained. An explanation is that there are crowding out effects and that banks simply prefer to lend to the government (see Stiglitz and Greenwald 2003). An alternative explanation may be that the high future tax liabilities this debt imposes on consumers makes banks less willing to lend. A problem with the latter interpretation is that the effect of wealth on excess sensitivity is insignificant. This result might as well be due to measurement error however since the measurement of wealth is not without problems. Note that the results found for bc_t and for b_t are also compatible with an

[15] The sample correlations are 0.64 for $b_t - w_t$, -0.12 for $b_t - u_t$, -0.32 for $b_t - bc_t$ when using NBER dummy, -0.31 for $b_t - bc_t$ when using Markov-Switching variable, -0.53 for $w_t - u_t$, -0.32 for $w_t - bc_t$ when using NBER dummy, -0.34 for $w_t - bc_t$ when using Markov-Switching variable, 0.15 for $u_t - bc_t$ when using NBER dummy,0.23 for $u_t - bc_t$ when using Markov-Switching variable and 0.78 between NBER dummy and Markov-Switching variable.

alternative interpretation for excess sensitivity. Specifically, it may be that both recessions and government debt raise uncertainty and precaution of consumers in the economy so that they will attribute more weight to current income in their consumption decisions. In this reasoning our results suggest that neither wealth nor unemployment affect general consumer uncertainty but government debt and bad times do.

Graphs of the evolution of $\beta_{t|t}$, as implied by the estimations in table 1, are presented in figures 2 and 3. In these figures the effects of the level shift in the early 80s and of the recessions (especially in figure 2) are quite clear. Also, the rise in $\beta_{t|t}$ in the 1980s and its subsequent fall in the 1990s reflect the effect of the evolution of the government debt. Note that in some periods $\beta_{t|t} < 0$ though not significantly (see footnote 2).

4.2 Robustness

First, to investigate the robustness of our results with respect to the instrument set, we estimate our model with an additional sixth lag for all instruments (except the error correction term).[16] As can be seen if we compare column 2 in table 1 with column 1 in table 2, this does not affect our conclusions (the only difference is that the coefficient on bc_t is now only significant at the 10 percent level). The second robustness test is to look whether our results are sensitive to the use of an alternative measure for fl_t. In section 4.1 we argue that including a simple dummy for fl_t may be a bit drastic. We therefore use an alternative measure, a linear spline, for fl_t. Liberalization increases deterministically until 1982 and then remains constant on the 1982 level. Obviously, this measure cannot be considered a "better" alternative than our 1982 dummy since there is not a prior reason to assume that liberalization increases deterministically at a constant rate. The results in column 2 of table 2 suggest that excess sensitivity decreases significantly during the seventies with 0.017 percentage points per quarter. Note also that while the coefficient on b_t remains significant, the coefficient on bc_t does not (though its t-value is still above one). Our final robustness check consists of the use of an alternative measure for wealth. Remark that the idea behind inserting wealth in the equation for excess sensitivity is that

[16] The regression of disposable income growth on the instruments has an R^2 of 0.38 and an R^2adj of 0.19.

wealth serves as collateral that could make access to credit easier for consumers. Banks most likely take into account the more "stable" part of wealth (like for instance housing wealth or other assets that can be considered safe) when deciding to attribute credit to individuals. Therefore, it makes sense to use a more structural measure for wealth, i.e. the trend in the personal wealth to GDP ratio. The results in column 3 of table 2 can be compared to the results in column 3 of table 1. We find that wealth and unemployment are still insignificant. The conclusions for financial liberalization and government debt are unchanged. However, bc_t measured with the NBER dummy now enters significantly at the 5 percent level. This result seems to confirm our concern mentioned earlier that the insignificance of the coefficient on bc_t may be due to the correlation of wealth with the business cycle. Note further that we also test whether the insignificance of wealth and unemployment in all our results is due to the inclusion of a business cycle variable. We do this by leaving out the business cycle in our estimations (results unreported but available upon request). This does not change our conclusions even though the coefficient on wealth becomes more significant (with a t-value approaching 1.5). Furthermore, since a significant part of wealth consists of government bonds, we also test whether leaving out government debt from our estimations affects the conclusions for wealth (results also unreported). We find that this is not the case.

Finally, we wish to emphasize that an alternative specification for eq.(5) containing β_{t-1} as an additional regressor has also been estimated. We find that this does never affect our conclusions and that the coefficient on β_{t-1} is never significant.

5 Conclusions

In this paper, we have empirically investigated a number of potential macroeconomic determinants of liquidity constraints mentioned in the literature (the state of the business cycle, the government debt ratio, the personal wealth ratio, the unemployment rate and a dummy capturing financial liberalization). We have analyzed whether these determinants have an impact on households consumption smoothing. This approach makes sense

because the most frequently cited reason for the lack of consumption smoothing by households and the observed excess sensitivity of private consumption to disposable income is the occurrence of liquidity constrained consumers. To the best of our knowledge, no rigorous attempt to investigate the impact of the state of the business cycle and the wealth ratio on the degree of consumption smoothing has yet been undertaken. Our approach thus adds to the existing literature on smoothing, especially because it considers several smoothing determinants simultaneously and because it investigates potential short-run fluctuations in excess sensitivity *in addition to* the long-run liberalization trend usually emphasized in the literature. Also, it adds to the literature on the financial accelerator because it investigates the impact of the business cycle on liquidity constraints of households rather than firms. Methodologically, our approach deviates from the existing literature because we estimate a system containing a consumption growth equation and a multivariate process for the excess sensitivity parameter.

The estimation results using quarterly US data over the period 1970-2000 suggest that consumption smoothing by households decreases in recessions and when the government debt ratio rises. Also, the level of excess sensitivity is significantly lower from the 1980s onward. This probably reflects financial liberalization in the 1980s. Coefficients on the wealth ratio and the unemployment rate have plausible signs and magnitudes but are insignificant. Our results seem to be in accordance with some of the predictions of the literature on liquidity constraints. We note that excess sensitivity of private consumption to disposable income may also be caused by precaution on the part of consumers related to uncertainty about future income and consumption. Our results are in accordance with this alternative interpretation as well.

What are the policy implications of these results ? First, to the extent that monetary or fiscal policy measures affect the state of the business cycle, restrictive policy, for instance, may increase liquidity constraints and negatively affect consumption, thereby reinforcing recessions.

Second, the results found imply that in periods of high debt or in recessions there is a stronger link between disposable income and private consumption. Thus fiscal policy

(and, more indirectly, monetary policy) may be more effective in recessions and when the government debt is high. This implication is not necessarily positive because at high debt levels the most likely stance of fiscal policy is a contractionary one. Also, it should be noted that potential higher effectiveness is caused by the impossibility of a large number of consumers to smooth consumption. This inevitably leads to a loss in welfare.

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Appendix A: bootstrap

We bootstrap in the following way:

1) We draw a vector $\mathbf{v}_t^*$ from the multivariate Gaussian distribution $N(\mathbf{0},\widehat{\mathbf{Q}})$ where $\widehat{\mathbf{Q}}$ contains the maximum likelihood estimates $\widehat{\sigma}_\zeta^2, \widehat{\sigma}_\mu^2$ and $\widehat{\sigma}_\varepsilon^2$.

2) Given the initial values $\mathbf{S}_{0|0}$ and given $\widehat{\mathbf{D}}$ and $\mathbf{Z}_t$, we construct pseudo state series $\mathbf{S}_t^*$ through eq.(7).

3) Then, given $\widehat{\theta}$, we construct a pseudo consumption growth series Δc_t^* through eq.(6).

4) We repeat steps 1-3 n times, where n is the number of replications (n=200 for all reported results). We thus obtain n pseudo series $(\Delta c_t^*)_i$ where $i = 1, ..., n$.

5) We use these pseudo series for consumption together with the true series for the other variables in the system. For each generated sample i, we calculate a sequence $\{\mathbf{S}_{t|t}^*\}_{t=1}^T$ and maximum likelihood estimates $\widehat{\mathbf{D}}^*$.

6) From these it is straightforward to calculate the standard error for each projection, $\mathbf{S}_{1|1}$ to $\mathbf{S}_{T|T}$, and for each of the parameters in $\widehat{\mathbf{D}}$.[17]

[17] We do not bootstrap standard errors for the variances in $\widehat{\mathbf{Q}}$ and for $\widehat{\theta}$ since these parameters are

Appendix B: data

Data for consumer expenditure on non-durables and services, disposable income, personal wealth, gross government debt, unemployment rate are from the Bank of England database (2003). Also from this database is the implicit deflator for non-durable consumption (1995=100) which is used to calculate real expenditures on non-durables and services, real disposable income and real personal wealth. Not from this database are:

- GDP and GDP deflator (1995=100). From IMF International Financial Statistics (2003).
- nominal 3 month T-bill rate. From IMF International Financial Statistics (2003).
- government consumption and investment. From IMF International Financial Statistics (2003). Deflated by the GDP deflator.
- population. From OECD Economic Outlook (2003).
- business cycle dummy. From NBER (2003).

restricted in the estimations.

Tables and Figures

Table 1: Maximum likelihood estimates and bootstrapped standard errors for estimation of system (6)-(7) on US data (1969:03-1999:02).

	(1) *time-invariant case*	*(2)* *Markov-Switching probabilities for* bc_t	*(3)* *NBER dummy for* bc_t
α_0	0.0037**	0.0000	0.0000
	(0.0005)	(0.0001)	(0.0001)
β_0	0.2692**	-0.3219	-0.2702
	(0.0607)	(0.8688)	(1.010)
β_1	-	-0.5507**	-0.5867**
		(0.1982)	(0.2051)
β_2	-	0.4241**	0.2165
		(0.2159)	(0.1470)
β_3	-	0.0224	0.0227
		(0.0546)	(0.0556)
β_4	-	-0.0040	-0.0044
		(0.0040)	(0.0045)
β_5	-	0.0258**	0.0270**
		(0.0107)	(0.0108)

Notes: Bootstrapped standard errors between brackets. *denotes significance at the 10% level, **denotes significance at the 5% level. Note that we do not report $\widehat{\sigma}^2_\zeta, \widehat{\sigma}^2_\mu$, $\widehat{\sigma}^2_\varepsilon$ and $\widehat{\theta}$. In column (1) the system (6)-(7) is estimated under the restriction $\beta_1 = \beta_2 = \beta_3 = \beta_4 = \beta_5 = \sigma^2_\zeta = \sigma^2_\mu = 0$ and with an F matrix that contains zeros, except in position (4,3) which contains a 1.

Table 2: Robustness tests for the results reported in table 1.

	(1) *additional lag "instruments" to forecast income*	*(2)* *linear spline for* fl_t	*(3)* *Hodrick-Prescott filtered wealth variable for* w_t
α_0	0.0000	0.0000	0.0000
	(0.0001)	(0.0001)	(0.0001)
β_0	-0.3170	0.9536	0.2772
	(0.8737)	(0.7936)	(1.3275)
β_1	-0.5149**	-0.0175**	-0.5954**
	(0.1888)	(0.0051)	(0.2129)
β_2	0.3777*	0.2654	0.2763**
	(0.2221)	(0.2149)	(0.1377)
β_3	0.0251	0.0400	0.0035
	(0.0484)	(0.0471)	(0.060)
β_4	-0.0032	-0.0060	-0.0079
	(0.0043)	(0.0043)	(0.0073)
β_5	0.0223**	0.0212**	0.0312**
	(0.0105)	(0.0087)	(0.0128)

Notes: Bootstrapped standard errors between brackets, *denotes significance at the 10% level, **denotes significance at the 5% level. Note that we do not report $\widehat{\sigma}^2_\zeta, \widehat{\sigma}^2_\mu$, $\widehat{\sigma}^2_\varepsilon$ and $\widehat{\theta}$. Column (1) is the same case as column (2) in table 1 except that a 6th lag of all instruments is added to construct fitted disposable income. The sample period used here is 1969:04-1999:02. Column (2) is the same case as column (2) in table 1 except that a linear spline (i.e. a variable that increases deterministically over time until 1982 and is constant from 1982 onward) is used for fl_t instead of a dummy. Column (3) is the same case as column (3) in table 1 except that the trend in the wealth to GDP ratio is used for w_t instead of the actual wealth to GDP ratio.

Figure 1: Estimated filtered probabilities of being in a recession and NBER turning points (vertical lines)

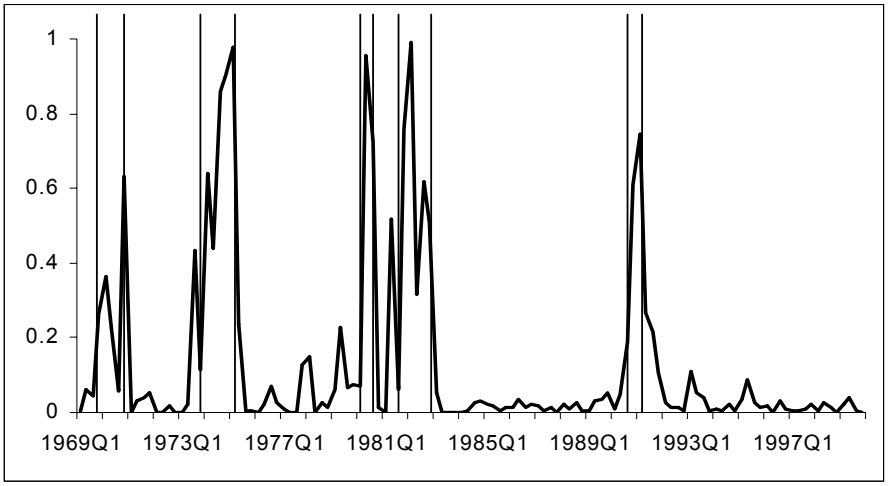


Figure 2: Time-varying excess sensitivity parameter ($\beta_{t|t}$) with 95% confidence bands (US data, 1969:03-1999:02, result for specification (2) in table 1)

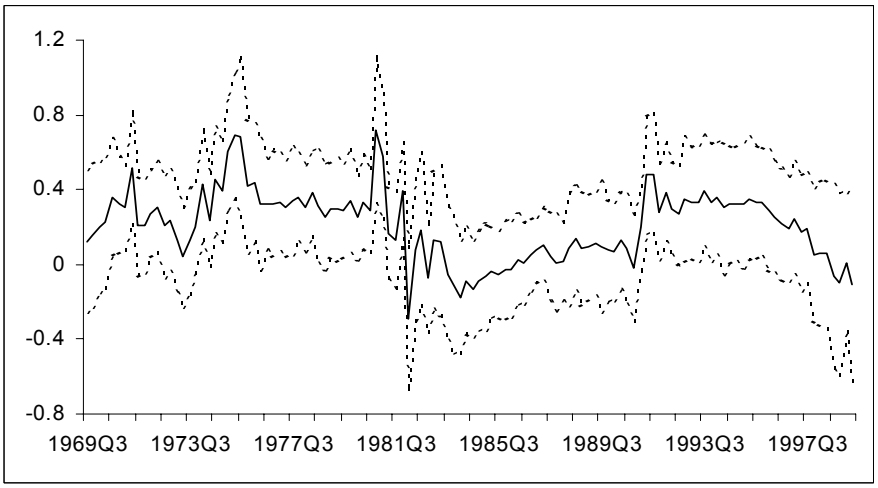
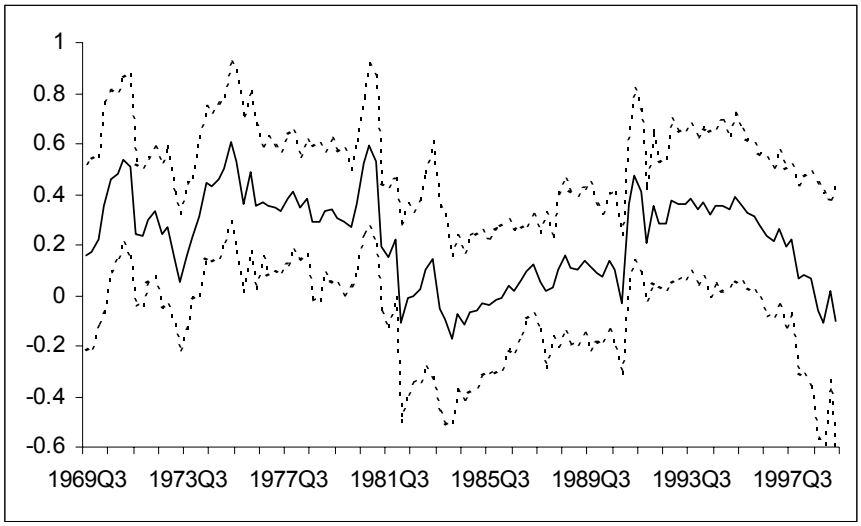


Figure 3: Time-varying excess sensitivity parameter ($\beta_{t|t}$) with 95% confidence bands (US data, 1969:03-1999:02, result for specification (3) in table 1)



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