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Liquidity Constraints and the Permanent Income Hypothesis *

Pseudo Panel Estimation with German Consumption Survey Data

Martin Beznoska[†], Richard Ochmann[‡],

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Abstract

This paper empirically investigates the relevance of liquidity constraints and excess sensitivity in intertemporal household consumption. Using a pseudo panel that has been constructed on rich German consumption survey data, we estimate the consumption responses to permanent and transitory income shocks, as well as the presence of excess sensitivity to anticipated income changes. A switching regression approach with unknown sample separation is applied to identify the two regimes whether to be liquidity constrained or not. The results are used to test whether liquidity constraints affect the validity of the permanent income hypothesis. For households in the constrained regime, reactions to changes in transitory income are found to be significantly greater than for households in the unconstrained regime. Furthermore, we provide evidence for excess sensitivity to anticipated income changes for households in the constrained regime if total consumption, durable as well as non-durable, is considered.

Keywords: Liquidity constraints, excess sensitivity, household consumption, switching regression, permanent income hypothesis.
JEL Classification: C34, D91, E21

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1. Introduction

There is extensive empirical literature on the question whether households smooth consumption over the life-cycle or rely mainly on their current income. The validity of either the former or the latter hypothesis has great importance for the implications of fiscal policy, in a structural sense as well as in the context of business cycles. The *Life Cycle-Permanent Income Hypothesis* refers to the work of Modigliani and Brumberg (1954) and Friedman (1957) and assumes that agents have rational expectations on their lifetime income and wealth. Lifetime resources are allocated among all periods of the lifecycle. Thereby, *permanent income* results as a long-term resource flow, which subsequently determines consumption in each period following the theory. If this hypothesis holds in reality, it would have many policy related implications, e.g. current income could not hold as a proxy for welfare or measuring inequality any more. Short-time fiscal spending in recessions should have no effect on consumption demand of private households whereas tax cuts in the income taxation should have effects, and so on.

The permanent income hypothesis (PIH) specifically predicts the effect of a non-anticipated shock to permanent income on consumption to be near one. Lifetime income is allocated to current and future consumption over the remaining periods, with respect to this new information. Consequently, in theory, a transitory shock has an influence near zero on consumption. The reason is that a positive transitory deviation from permanent income is saved nearly completely while a negative one is compensated by dissavings or taking-up a credit. The agent perceives the deviation and knows that its expected value will be zero over lifetime, and thus a transitory shock should have no relevant effect on consumption.

However, households are often found to behave differently in response to a shock than theory predicts (Jappelli and Pistaferri, 2010). There are numerous empirical findings suggesting that the PIH does not hold in the data.[1] Consumption has been found to be excessively smooth with respect to permanent income shocks (e.g. Campbell and Deaton, 1989; Attanasio and Pavoni, 2011), as well as excessively sensitive towards transitory income shocks (e.g. Hall and Mishkin, 1982; Souleles, 2002). In particular, the elasticity of permanent income is frequently estimated below one while the elasticity of transitory income is found to lie below the permanent one, but significantly different from zero (e.g. Blundell, Pistaferri, and Preston, 2008b). Excessive response to transitory shocks has been put in the context of transitory income uncertainty. It implies that current income is relatively more important for intertemporal consumption allocation than permanent income. This has been motivated in the literature with either precautionary motives (e.g. Carroll and Samwick, 1998), or deviations from rational behavior,

[1] See Jappelli and Pistaferri (2010), or Attanasio and Weber (2010) for literature reviews on the consumption response to anticipated income changes and non-anticipated income shocks.

such as myopia, inertia, loss aversion, or habit formation (e.g. Shea, 1995a), or with the presence of liquidity constraints (e.g. Zeldes, 1989b).[2]

Another aspect of the PIH, which has gained much attention in the literature, is the excess sensitivity of consumption to anticipated income changes (Flavin, 1981). Anticipated changes in income should have no effect at all on consumption because they are assumed to be already internalized. The literature on this topic finds significant evidence for excess sensitivity rejecting the PIH, which is also explained by the appearance of liquidity constraints (e.g. Zeldes, 1989a; Garcia, Lusardi, and Ng, 1997; Jappelli, Pischke, and Souleles, 1998).

In this paper, we empirically investigate to which extent deviations from the PIH can be traced back to the presence of liquidity constraints in household consumption, utilizing a pseudo panel constructed on rich German consumption survey data. In a switching regression framework, we estimate marginal propensities to consume out of permanent and transitory income shocks, and we test for excess sensitivity of consumption to anticipated income changes.

The concept of liquidity constraints in intertemporal consumption is usually placed in the environment of incomplete markets, where agents' possibilities to insure consumption levels are limited to self insurance (Kaplan and Violante, 2010). The necessary condition for liquidity constraints is that agents do not hold enough liquid assets to keep up *permanent* consumption, which is determined by lifetime income. The sufficient condition is usually assumed fulfilled if agents are not able to borrow as desired at an interest rate that is in an acceptable range around the market lending rate or can not allocate the amortization payments in the preferred periods.[3] Instead of offering a higher interest rate, the credit institution would typically turn down the agent's request because of adverse selection issues.

Evidence for liquidity constraints in the literature is mixed. There are numerous studies that find evidence for liquidity constraints (e.g. Zeldes, 1989b; Kaplan and Violante, 2010), while other studies do not find any support for their relevance (e.g. Shea, 1995b). Evidence is also found to be mixed with respect to the size of the income change (e.g. Hsieh, 2003). Many studies that find evidence for liquidity constraints affecting consumer behavior focus on clear identification of exogenous income changes. In this context, several papers have looked at spending of tax rebates (e.g. Souleles, 2002; Johnson, Parker, and Souleles, 2006), others at repayment of car loans (e.g. Stephens, 2008); again others have utilized external information, such as credit card data (e.g. Gross and Souleles, 2002).

Often, availability of adequate data in this context is an issue. Household panel data on (total) consumption is only rarely available over a longer period of time. Thus, often repeated

[2]See Browning and Lusardi (1996) for a survey on savings motives and empirical evidence on household savings behavior.

[3]This is in line with the definition of liquidity constraints in the literature (see e.g. Garcia et al. (1997)).

cross-sections on micro consumption data are applied in the literature to investigate intertemporal consumption decisions (e.g. Blundell, Low, and Preston, 2008a). In addition, typically, information on income and detailed consumption (non-durable as well as durable consumption) is not available jointly in micro data, so that information must be imputed (Blundell et al., 2008b), or consumption (pseudo) panel data be constructed (Alessie, Devereux, and Weber, 1997).

We make use of a pseudo panel constructed on repeated cross-sections of rich consumption survey data for Germany to investigate the consumption effects of income shocks in the context of liquidity constraints. In this data set, we observe income and consumption jointly, both durable and non-durable consumption, very detailed by purchases of durable goods and spending on non-durable commodities. By applying specific treatment for purchases of durables we account for relevant effects of liquidity constraints among durable consumption (Alessie et al., 1997; Attanasio, Goldberg, and Kyriazidou, 2008). Thus, we have relatively precise measures of the individual income and consumption processes, and can utilize their joint evolution over time to disentangle the consumption effects of income shocks into transitory and permanent elements.

We apply two models, one is referring to the marginal propensities to consume out of permanent and transitory shocks and the other to test for excess sensitivity. For the former, a two-stage approach is applied to pseudo panel data, where current disposable income has to be split up into a permanent and a transitory part to identify its shocks at the first stage. At the second stage, we estimate a consumption growth equation including the permanent and transitory income shocks from first stage estimation as explanatory variables. For the latter, we re-specify the consumption growth equation to perform the excess sensitivity test.

For the consumption growth equations, we use a switching regression approach with unknown sample separation by applying an iterated two-step procedure with the EM algorithm (see Dempster, Laird, and Rubin (1977)) to identify the two regimes whether to be liquidity constrained or not. Our contribution is twofold. While these kinds of models are used in the literature to test for excess sensitivity (Garcia et al., 1997), identification of two regimes based on the two income components has not been applied yet. Also, consumption processes based on a pseudo panel of rich consumption data have not been analyzed for Germany so far in this context to the best of our knowledge; in particular not allowing for explicit treatment of durable consumption. We find that for households in the constrained regime, reactions to changes in transitory income are significantly greater than for households in the unconstrained regime. In addition, there is evidence for excess sensitivity to anticipated income changes for households in the constrained regime if total consumption, durable as well as non-durable, is considered.

The remainder of the paper is organized as follows. In the next section, the model and the empirical strategy are introduced. Section 3 presents the data and some descriptive evidence. Results are provided in Section 4, and Section 5 concludes.

2. The Model

The model is presented in three steps. Firstly, the underlying income process and the consumption growth equation are derived. Then, estimation of the model in a switching regression approach with unknown sample separation is explained and income shocks are integrated into the model. Finally, the model is adjusted to allow for a test for excess sensitivity.

The Income Process and the Consumption Growth Equation

Firstly, we focus on the estimation of the marginal propensities to consume out of permanent and transitory shocks. Therefore, the first stage concerns the income process. Current disposable income is the observed variable and thus has to be split up into a permanent and a transitory part. For this issue, we run a fixed-effects regression of current income on covariates at the pseudo-panel level (see Section 3 for details on constructing the pseudo-panel). The associated equation looks like this:

$$\ln(y_{it}) \;=\; \delta_1 \text{age}_{it} + \delta_2 \text{age}^2_{it} + X'_{it}\beta + \alpha_i + \omega_{it} \tag{1}$$

where we allow the error-term ω_{it} to be autocorrelated

$$\omega_{it} \;=\; \rho\omega_{it-1} + \epsilon_{it}. \tag{2}$$

$\ln(y_{it})$ denotes the natural logarithm of current income. As covariates, we include age, age squared and a vector X_{it} that contains interactions of the age polynomials with the social status of the household head and household composition. Additionally, we include other household characteristics in X_{it}, such as skill-level of the household head and information on the partner. α_i is a cluster-specific fixed effect, ω_{it} is an autocorrelated error-term and ϵ_{it} is white noise. At this point, we can define a prediction for the permanent and transitory parts of current income, which are

$$\pi^P_{it} \;\equiv\; \widehat{\ln(y_{it})} \;=\; \hat{\delta}_1 \text{age}_{it} + \hat{\delta}_2 \text{age}^2_{it} + X'_{it}\hat{\beta} + \hat{\alpha}_i \tag{3}$$

for permanent income and

$$\pi^T_{it} \;\equiv\; \hat{\omega}_{it}\,, \;\; E[\omega_{it}] = 0 \tag{4}$$

for transitory income.

We are interested in the *non*-anticipated shock to permanent income, in this first analysis. Therefore, we investigate the dynamic processes that underlie π_{it}^P. Non-anticipated permanent income changes should be free of any autocorrelation structures; they should have an expected value of zero. This should also hold for π_{it}^T so that the autoregressive parameter in Eq. (2) should be close to zero. We provide more details on this procedure in Section 4, where we further investigate and discuss the dynamics of the income process. We assume for the further analysis that permanent income (π_{it}^P) and transitory income (π_{it}^T) have an expected value of zero so that changes to them can be interpreted as non-anticipated income shocks.

At the second stage of the approach, a consumption equation is applied, in which the validity of the permanent income hypothesis can be tested. Jappelli and Pistaferri (2010) show how starting from an Euler equation and making some assumptions about the consumption and income processes leads to a consumption growth equation, in which the parameters can be interpreted as structural, and thus they allow for testing the theory. Hall and Mishkin (1982) identify these parameters via contemporaneous and serial correlation between income growth and consumption growth. For our switching regression approach, a parameter identification via variance-covariance matrix is not feasible. Instead, we start off with a reduced form level equation in our empirical specification:

$$\ln(c_{it}) \;=\; \phi\pi_{it}^P + \psi\pi_{it}^T + M_i^{'}\gamma_1 + Z_t^{'}\gamma_2 + \gamma_3\text{age}_{it} + \gamma_4\text{age}_{it}^2 + \xi_{it} \tag{5}$$

where $\ln(c_{it})$ denotes the natural logarithm of consumption, M_i is a vector of time-invariant household characteristics, Z_t is a vector of time dummies and ξ_{it} is an independent error-term. In our main specification, consumption contains all non-durable consumption. In a robustness check, we analyse total consumption, non-durable as well as durable, by adding generated user costs for durable goods to non-durable consumption. See Appendix A for details on how user costs have been generated.

First-differencing of Eq. (5) leads to:

$$\Delta\ln(c_{it}) \;=\; \phi\Delta\pi_{it}^P + \psi\Delta\pi_{it}^T + Z_t^{'}\gamma_2 + \gamma_5\text{age}_{it} + \Delta\xi_{it} \tag{6}$$

where $\Delta\xi_{it}$ is an independent error-term, which is assumed to be uncorrelated with the shocks $\Delta\pi_{it}^P$ and $\Delta\pi_{it}^T$.[4] This assumption is crucial to gain consistent coefficients. Eq. (6) should in general be consistent with the one proposed by Jappelli and Pistaferri (2010). The non-differenced time dummies Z_t appear here, because we allow this equation to have an intercept for estimation purposes. This does, however, not prevent us from capturing all variation over

[4] Note that first-differencing of the age polynomial yields: $\Delta(\text{age} + \text{age}^2) \;=\; 1 + 2\cdot\text{age}$.

time in Eq. (6) that would be captured with differenced time dummies.

The implications that are given by theory suggest the null hypothesis of $\phi = 1$ and $\psi = \kappa$, where κ depends on the interest rate and the marginal propensity to consume out of assets and should be a small value (see Hall and Mishkin (1982)). κ declines if remaining life time increases. As a rule of thumb, κ can be approximated by $\frac{1}{T}$ where T is the expected number of remaining periods of lifetime.[5] This approximation is used in the following to test the permanent income hypothesis. It suggests that κ is a function of age. Thus, in a robustness check, we interact age with the transitory income shock to test whether reactions to transitory shocks differ by age of the household head.

The Switching Regression and the Shock Model

For identification of the two regimes, whether to be liquidity constrained or not, a switching regression approach with unknown sample separation is applied to the consumption equation Eq. (6). The exact classification of households facing liquidity constraints is difficult in household survey data. Questions directly relating to this issue, like "Do you have access to as much credit as desired in your credit institution?" appear rather rarely in common surveys, and there is no such information in the data we use.

Variables that are related to the issue of liquidity constraints are current disposable income and the ratio of financial wealth to permanent income, for example, which refer to the aspect of available household liquidity. Then there are proxies for the uncertainty of future income flows, such as the status of current unemployment or being in education, as negative examples, and being a civil servant, as a positive example. Another indication that should reflect the absence of liquidity constraints, and a status of good creditworthiness, is whether households can afford to repay their loans, i.e. if they have a high ratio of amortisation payments related to the level of debt. But despite all these indicators, which are potentially related to liquidity constraints, it remains challenging to separate the sample into two regimes. Besides the problem that all these variables are continuous so that setting a sample-separating threshold can only be arbitrary, other difficulties arise in terms of multidimensions when interactions between the indicators are constructed.[6]

Zeldes (1989a) splits the sample on the basis of various criteria on the wealth to income ratio and tests the permanent income hypothesis against several thresholds. He finds his results to be quite sensitive with respect to different sample splits. However, because of the arbitrary

[5] This approximation assumes that one extra transitory Euro in the current period will be allocated equally to consumption over all remaining periods of lifetime.

[6] Note that in the pseudo-panel the information about being unemployed, in education or in civil service is a continuous variable, too. It only reflects the cluster share of household heads that are in the specific status.

splitting criteria and the resulting need to estimate the model for numerous specifications of sample splits, we prefer to apply a switching regression approach with unknown sample separation, as suggested by Garcia et al. (1997). While on the one hand, this appears to be an elegant way to "let the data speak" about which two regimes can be identified via selection equation, on the other hand, Maddala (1986) points out that one maybe "asks too much from the data" in this kind of switching regression and that maximum likelihood estimation may result in local, rather than global maxima due to unboundedness of the likelihood function. But he also states that the results from empirical applications are "surprisingly" good.

Our model consists of equations for the two regimes mentioned earlier and a selection equation, resulting in the 3-equation-model

$$\begin{aligned} \Delta \ln(c_{it}) &= \phi_1 \Delta\pi^P_{it} + \psi_1 \Delta\pi^T_{it} + \Delta\xi_{1it}\,, \qquad \text{if} \quad K'_{it}\lambda + u_{it} < 0 \\ \Delta \ln(c_{it}) &= \phi_2 \Delta\pi^P_{it} + \psi_2 \Delta\pi^T_{it} + \Delta\xi_{2it}\,, \qquad \text{if} \quad K'_{it}\lambda + u_{it} \geq 0 \end{aligned} \tag{7}$$

where control variables have been left out for clarification, see Eq. (6) for details.[7] The subscripts 1 and 2 denote the belonging to the respective regime and K_{it} is a vector of variables that are assumed to determine the presence of liquidity constraints. K_{it} contains current disposable income, the ratio of financial wealth to permanent income, and the ratio of amortisation payments to the level of debt. Additionally, K_{it} includes household characteristics, such as age of the household head, household composition, and interactions between the characteristics and current income, as well as time dummies. The error-terms $\Delta\xi_{1it}$, $\Delta\xi_{2it}$ and u_{it} are assumed to be independent and normally distributed with variances $\sigma^2_{\xi_1}$, $\sigma^2_{\xi_2}$ and σ^2_u. The latter is set to be 1 for identification purposes.

The proxies for the income shocks (π^P_{it}, π^T_{it}) in Eq. (7) are derived from the income equation (Eqs. (1) and (2)) estimated separately from the switching regression at a first stage. There is potentially heterogeneity in the income processes over the two regimes. However, the switching regression approach that we apply here does not allow for an integretation of the income processes, the selection equation, and the consumption growth equation. We assume that the two regimes face identical income processes.

The model can be estimated by maximizing the likelihood function (Garcia et al., 1997)

$$f(\Delta \ln(c_{it})) = P_{it} f(\phi_1, \psi_1, \Delta\xi_{1it}) + (1 - P_{it}) f(\phi_2, \psi_2, \Delta\xi_{2it}) \tag{8}$$

where the $f(\cdot)$ function denotes the density of the normal distribution. This maximization

[7]The control variables are of course allowed to vary between the two regimes, too.

problem is solved by applying an iterated two-step procedure with the EM algorithm (see Dempster et al. (1977)).[8] In in the second step (the main equation) of the two-step alogorithm, the observations are weighted by their probability of belonging to each of the regimes, which in turn depends on the first step (the selection equation).

Let therefore

$$P_{it} = \frac{f(\Delta\xi_{1it})\Phi(-K_{it}'\lambda)}{f(\Delta\xi_{2it})(1-\Phi(-K_{it}'\lambda))} + f(\Delta\xi_{1it})\Phi(-K_{it}'\lambda) \tag{9}$$

denote the probability of belonging to the first regime, where the $\Phi(\cdot)$ stands for the normal cumulative distribution function.[9]

The Excess Sensitivity Test

In the absence of excess sensitivity, the response to *anticipated* changes to permanent and transitory income should be zero. The PIH predicts that consumption has already been adjusted with respect to anticipated income changes when the intertemporal consumption allocation has been planned for all periods in advance. There should be no additional response to anticipated changes. This is what we what to test. We extend our basic model from the first analysis to allow for a test for excess sensitivity to anticipated income changes.

Flavin (1981) pointed out that, in the absence of excess sensitivity, consumption should follow a martingale, in which information gained in the past does not matter for the present consumption decision.

$$\ln(c_{it+1}) = \ln(c_{it}) + \nu_{it+1} \tag{10}$$

where ν_{it+1} is an error-term that covers all *new* information that is available to the household in $t+1$, including shocks. If we subtract $\ln(c_{it})$ on both sides of the Eq. (10) and add on the right-hand side a rational expectation term of income changes, which is based on information available at time t and should thus be zero, this should not alter the basic functional relation. Thereby, we nest the excess sensitivity test within an Euler equation derived from dynamic utility optimization in the context of intertemporal consumption allocation, following Runkle (1991) and Zeldes (1989a):

$$\Delta\ln(c_{it+1}) = \delta E_{it}\left[\Delta\ln(y_{it+1})\right] + Z_{t+1}'\gamma_6 + \gamma_7\text{age}_{it+1} + \gamma_8\Delta\text{adults}_{it+1} + \gamma_9\Delta\text{kids}_{it+1} + \nu_{it+1} \tag{11}$$

where in our empirical specification, we add the same control variables that are included in the consumption growth equation Eq. (6). To control for taste shifts of the household, we include

[8] This has been done here using the user-written Stata routine "switchr" (see Zimmerman (1999)).

[9] In the first iteration, the densities that stem from the residuals of the main equation have been set to one.

the first differences of the numbers of adults as well as children in the household, where we follow the specifications applied in Garcia et al. (1997) and Jappelli et al. (1998). In these specifications, it is suggested to use as a proxy for the expectational term $E_{it}[\Delta \ln(y_{it+1})]$ the observed income in t, $\ln(y_{it})$. We follow this approach.

The null hypothesis on the validity of the PIH in the context of excess sensitivity that we test is $\delta = 0$ in Eq. (11). In applying the switching regression approach from the first analysis (Eq. (7) - Eq. (9)) on Eq. (11), we again aim at identifying two regimes according to the presence of liquidity constraints and test the PIH for both of them.[10]

3. Data and Descriptive Evidence

Firstly, the data set applied is introduced and the conversion from the household level to the synthetic panel level is described. Then, some descriptive evidence on the composition of the clusters generated for the pseudo panel, as well as on cluster-average income and consumption shares is presented.

Data

The micro data applied in this analysis stems from the Continuous Household Budget Survey for Germany (*Laufende Wirtschaftsrechnungen*, LWR). The LWR data is repeated cross-sectional consumption survey data, which is repeated every year since 2002. It is maintained by the German Federal Statistical Office (*Statistisches Bundesamt*).[11] It contains information on income, consumption, and savings, very detailed by single components, at the household level. Households are recruited voluntarily for reports every year, according to stratified quota samples from Germany's current population survey (*Mikrozensus*). Six waves are available covering the time period from 2002 to 2007, where we face a break in data structure since 2005.

In the first three waves, the sampled households are observed for a time of four months (one month out of each quarter of the year). Since 2005, recruited households stem from a subsample of the Income and Consumption Survey for Germany (*Einkommens- und Verbrauchsstichprobe*, EVS) and are observed for an entire quarter. This means that we have monthly data from 2002 to 2004 and quarterly data from 2005 to 2007. This structural break in the data has been accounted for in the analysis and it is addressed in further detail in Section 4. Altogether, the pooled data set contains 91,359 observation at the household level.

[10] Note that the variables in K_{it} that actually identify the selection may vary from those in Eq. (7) for stability reasons in the iterations of the switching regression.

[11] The LWR data were provided by the Research Data Centre of the Statistical Offices of the Länder *(Forschungsdatenzentrum der Statistischen Landesämter, FDZ).*

Construction of the Pseudo-Panel

In order to eliminate a cluster specific fixed-effect and to be able to model the dynamics in our approach, we construct a pseudo-panel from the repeated cross sections. This is done by forming 17 clusters, which are observed over 48 time periods (3×12 in the waves 2002-2004 and 3×4 in the waves 2005-2007). The 91,359 observations at the household level are organized into these 17 clusters by three dimensions: birth cohort group, gender, and level of education. This results in a volume of 816 cells at the pseudo-panel level where the criteria, especially the size of the birth cohort group, are set in such a way that we have a similar number of observations in each cell. In the end, the average number of observations per cell is about 112, where the smallest cell has 77 and the biggest cell reaches 175 observations. More details can be found in the next subsection. All variables that are available at the household level are averaged over all observations in a cell.[12] These averaged variables become the pseudo-panel variables, which are measured with measurement error if the composition of the cells varies over time in the selected criteria.

Due to the fact that the measurement error, and thus the potential bias in the estimated coefficients, diminishes with the number of observations per cell, Verbeek and Nijman (1992) suggest that at least about 100 observations per cell should be reached to fulfill consistency. While the population per cell in our pseudo-panel is slightly above 100 on average, and additionally there is nearly the same composition per cluster within the waves 2002, 2003 and 2004, we argue that consistency is given in our constructed pseudo-panel. As mentioned above, the specific characteristic of the data structure, which consists of half monthly and half quarterly data, has been accounted for in the dynamic estimations (see Section 4 for details).

Descriptives

Table 1 displays descriptive statistics on cluster composition and on income and total consumption on cluster averages, as well as over all clusters. Total consumption contains non-durable consumption as well as generated user costs for durable consumption (also see Appendix A). In the regression analysis, we focus on total consumption as well as non-durable consumption.

The composition on the clusters reveals that households are distributed sufficiently evenly across the 17 cluster. Clusters are composed of between 4,699 and 6,726 observations over all points in time (which corresponds to 5.1% and, respectively, 7.4% of all observations). As a

[12] Note that dummy variables at the household level only form dummies in the pseudo-panel if all observations face the same outcome, which is the case e.g. for time dummies and gender. If the observations in a cell have heterogeneous outcomes for a dummy the averaging will lead to a proportion variable at the pseudo-panel level, which can be treated as a normal variable in the regressions.

result, clusters are filled, on average over all points in time, with about 112 households (ranging between 98 and 140).

Table 1: Income and Consumption by Clusters

Cluster[a]	N	N_j/N	N_j/T	age	$\bar{y}_j^{curr}$	$\bar{c}_j(\%)$
All clusters	91,359	100.0	1,903	55.0	2,791	88.8
(m, h, <1937)	6,726	7.4	140	74.4	2,870	90.9
(m, h, 1937-1942)	5,547	6.1	116	64.8	3,287	91.7
(m, h, 1943-1950)	5,654	6.2	118	58.1	4,193	87.9
(m, h, 1951-1956)	5,584	6.1	116	51.2	4,483	82.8
(m, h, 1957-1962)	5,180	5.7	108	44.8	4,532	77.5
(m, h, >=1963)	5,187	5.7	108	36.8	4,088	76.1
(m, l, <1938)	5,207	5.7	108	74.2	2,124	92.3
(m, l, 1938-1945)	4,699	5.1	98	62.9	2,782	94.7
(m, l, 1946-1954)	4,985	5.5	104	54.3	3,205	88.7
(m, l, 1955-1962)	5,129	5.6	107	46.1	3,254	87.8
(m, l, >=1963)	4,895	5.4	102	36.6	2,964	82.2
(f, h, <1948)	5,618	6.1	117	68.5	1,959	94.6
(f, h, 1948-1958)	5,577	6.1	116	51.6	2,615	89.5
(f, h, >=1959)	5,431	5.9	113	38.2	2,605	83.4
(f, l, <1942)	5,482	6.0	114	72.8	1,379	96.5
(f, l, 1942-1958)	5,581	6.1	116	54.4	2,007	91.7
(f, l, >=1959)	4,877	5.3	102	38.4	1,959	90.9

Notes: $\bar{y}_j^{curr}$ is current household income in monthly averages, in real terms, and weighted by population weights. The pseudo-panel weight for permanent income is the average household weight in each cell. N_j/T is the average number of observations in the cluster over the 48 points in time. $\bar{c}_j(\%)$ is the average consumption rate, as a share from current income, in percent and weighted.
[a]: Clusters defined by gender of household head, education of household head, and year of birth of household head. E.g., (m, h, 1937-1942) is for males, highly educated, born between 1937 and 1942.

Source: Own calculations using the LWR data (2002-2007), provided by the FDZ.

For an average household, current income in monthly terms amounts to 2,791 euros, on average over the 48 points in time. The average household consumes 88.8% of current income and saves the rest. Its head is 55.0 years of age on average. When current household income in monthly terms and consumption as a share of current income (in %) are broken down by the 17 clusters, between-cluster heterogeneity is revealed. Generally speaking, current income is around average, or lower than average, for all clusters with a female head. As would be expected, income is relatively greater for the highly educated than for the less educated. Incomes are also greater for the younger cohorts than for older cohorts. The greatest incomes are found for the clusters of highly educated male heads and the lowest among households headed by less educated females.

There is also between-cluster variation in the total consumption share, averaged over the 48 time periods. The main apparent pattern is that consumption shares are relatively lower for younger cohorts. While they are above 90% for cohorts born before 1943, they are between 85% and 90% for most of the cohorts born between 1943 and 1960, and they are lower than 80% for some of the cohorts born after 1960. This pattern is, of course, a mixture of an age effect and potential cohort and time effects, which are not be separable in this interpretation. As we would expect the consumption share to increase in old age when agents run down their assets according to the life-cycle hypothesis, and we only observe cohorts for a period of seven years here, we would expect the consumption shares of the older cohorts to be relatively greater.

Consumption shares are furthermore slightly greater for the clusters of the less educated household heads (82.2%-96.5%) than for those of highly educated ones (76.1%-94.6%). They are also slightly greater for clusters with a female household head (83.4%-96.5%) than for those with a male head (76.1%-94.7%). This descriptive consumption evidence is probably contaminated by an income effect that is not controlled for here.

4. Results

This section is divided into three parts: 1) results for the income process, 2) for the consumption growth equation, and 3) for the excess sensitivity test. Due to the two different data structures underlying our constructed pseudo panel (see Section 3), a differentiated treatment is required at some points. Firstly, some conceptual points resulting from these structural differences for the estimation of the income process are discussed. Results from this first stage are used to estimate the consumption growth equation and the excess sensitivity test, by both OLS and the switching regression technique. Results for the main specification as well as for some alternative specifications and robustness checks are presented and discussed in the second and the third subsection.

Results for the Income Process

While estimation of Eq. (1), the fixed effect model, is quite standard and leads to plausible results if it is estimated with the whole pseudo-panel (see Table B.1 in Appendix B for results), we shall further elaborate somewhat on the results for the dynamic specification. We estimated an autocorrelation coefficient ρ in Eq. (2) of -0.1, which indicates a relatively small persistence in the error-term. This result gives us faith that we can probably assume that the household preceives this part of the income variation as truely transitory so that we can interpret this residual effect as a proxy for transitory income shocks.

As already mentioned in Section 2, we are interested in the *non*-anticipated permanent income shocks. Thus, we investigate the dynamics of the proxy for permanent income, π_{it}^P. To serve as a good proxy for non-anticipated permanent changes, the residual of the dynamics should be free of any autocorrelation structure and should have an expected value of zero. Thus, if we would see an estimate for the autocorrelation term in the dynamics of π_{it}^P that is significantly smaller than 1, this would result in a non-zero expected value for the residual and we could not interpret the residual as a shock to permanent income, which has not been anticipated by the household.

In the estimation of the dynamics of permanent income, we have to account for the structural break from monthly to quarterly waves in the survey data. Firstly, we analyse the autocorrelation structure of π_{it}^P for the quarterly waves and estimate an AR(1)-process for $\widehat{\ln(y_{it})}$. Applying the Arellano-Bover system GMM estimator, using the second lag as an instrument for the endogenous first lag, to estimate the AR(1)-process for the waves 2005, 2006 and 2007 separately, we cannot reject the null hypothesis of an autocorrelation coefficient of 1 for each wave.[13] This result provides evidence for the assumption that the residual of the permanent income dynamics can be interpreted as a proxy for true shocks that are non-anticipated. Conclusively, we expect the first differences in permanent income to be stationary and apply them as a proxy for permanent shocks in the consumption growth equation Eq. (6).

On the contrary, the monthly waves (2002 to 2004) are autocorrelated per construction within a year because a household is observed for one month of each quarter. This break in the data structure requires an alternative treatment. We therefore regress the first-difference of permanent income in t on the respective first-difference in $t-3$. The autocorrelation coefficient is estimated at 0.92 in 2002, not significantly different from 1, which implies that a unit root cannot be rejected. However, in 2003 and 2004, the coefficient is found to be significantly smaller than one (0.88 and 0.78). To correct the first-difference for the quarterly autocorrelation, we take the residual of this regression as our estimate for $\Delta\pi_{it}^P$ in the waves 2002 to 2004.[14]

Results for the Consumption Growth Equation

Table 2 displays the results from the estimation of the consumption growth equation, Eq. (6), where firstly only non-durable consumption is considered as the dependent variable. Coefficient

[13]This implies a unit root for permanent income. The Sargan test statistic suggests valid instruments in all cases with p-values of 0.33 (2005), 0.75 (2006) and 0.50 (2007).

[14]We also checked a specification where quarterly pseudo panel observations are constructed from the monthly micro data. While this reduces the number of pseudo panel observations drastically, the OLS results do not vary much. However, switching regression estimation was not feasible for this specification due to non-convergence of the EM algorithm. Also see footnote 15.

estimates for the permanent shock (ϕ) and the transitory shock (ψ) from the OLS estimation of the pooled model, as well as the switching regression for the regime model, are presented. Effects of covariates are left out here; full results are relegated to Table B.2 in Appendix B. The effects in Table 2 can be interpreted as the marginal propensity to consume (MPC) out of permanent income, respectively transitory income. Additionally, test results for hypotheses on the effects from theory are presented.

Table 2: Marginal Effects for the Consumption Growth Equation (**Non-durable** Cons.)

	OLS	**Switching Regression**	
Dep. var.: $\Delta \ln(c_{it})$	Pooled Model	Regime 1 (Constrained)	Regime 2 (Unconstrained)
Permanent shock (ϕ)	0.667***	0.626***	0.787***
	(0.058)	(0.055)	(0.133)
Transitory shock (ψ)	0.204***	0.427***	0.048
	(0.046)	(0.040)	(0.049)
Probability of regime 1 (P_{it})	-	0.422	
N (cells)	748	748	748
R^2	0.215	0.586	0.146
Tests (χ^2-statistic):			
$\phi = \psi$	40.12	11.05	24.79
$\phi = 1$	32.99	46.72	2.56
$\psi = \kappa = 0.05$	11.41	86.94	0.00
$\phi_1 = \phi_2$	-	1.26	
$\psi_1 = \psi_2$	-	35.58	
$\phi_{SWR} = \phi_{OLS}$	-	0.27	0.68
$\psi_{SWR} = \psi_{OLS}$	-	13.33	5.44

Notes: See Table B.2 in Appendix B for complete estimation results and full list of covariates. Significance levels: $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$; robust standard errors in parentheses, adjusted for clustering, not adjusted for two-step estimation.

Source: Own calculations using the LWR data (2002-2007), provided by the FDZ.

In the pooled model, the MPC out of permanent income is estimated at 0.667 and the MPC out of transitory income at 0.204. If permanent income increases by 10% consumption increases by 6.7%, on average for the entire sample. However, if the 10%-increase in income is of transitory nature consumption increases only by 2.0%. Although the reaction to transitory shocks is found to be significantly smaller than to permanent changes (χ^2-statistic = 40.12), these effects do not correspond to the PIH. The hypothesis that the reaction to a permanent shock is unity must be rejected at the 1% level (32.99). For a transitory shock, it shall be tested whether the reaction equals $\kappa = 0.05$, which results from the approximation in Section 2. This hypothesis must also be rejected at the 1% level (11.41).

In the regime model, estimated by the switching regression, the results are generally in the same range as for the linear model. For the constrained regime, which is expected to contain households that are identified to be liquidity constrained, the MPC out of permanent income is estimated at 0.626. This is significantly lower than unity (χ^2-statistic = 46.72), but it is not significantly different from both the unconstrained regime (1.26) and the pooled model (0.27). Most remarkably, the MPC out of transitory income is estimated at 0.427, which is notably great and in particular greater than for the other groups. It is significantly different from 0.05 (86.94) and from the coefficient for the pooled sample (13.33). The initial probability of being in the constrained regime is estimated at 42.2%.

Considering the coefficients for the unconstrained regime, the MPC out of permanent income is estimated at 0.787, which is not significantly lower than unity at the 5% level (χ^2-statistic = 2.56), but not significantly different from the result for the pooled model either (0.68). The standard error here is quite large compared to the other ones found in the OLS equation and in Regime 1. The effects for the *transitory* shock differ significantly over the regimes. The MPC out of transitory income is estimated at 0.048 for the unconstrained regime, which is neither significantly greater than 0 nor than 0.05, but it is significantly lower than for the constrained regime (35.58), and also lower than for the pooled model, at least at the 5% level (5.44). Furthermore, a test on equality of the two parameters within the regime is strongly rejected (24.79).

We have also estimated the switching regression model again, this time allowing switching to be determined endogenously. We do not find significant differences in the coefficients, although the selection term in the constrained regime is significant (see Table B.4 in Appendix B for details). A remarkable difference is that the standard error of $\widehat{\phi}$ in the unconstrained regime is clearly smaller with 0.077 than in the model with exogenous switching. Here, the test for $\widehat{\phi} = 1$ is rejected at the 1% level. Also, it is significantly different from the coefficient in the constrained equation, at least at the 10% level (χ^2-statistic = 3.29), which means that the two regimes are significantly different in both coefficients, the permanent and the transitory one. A limitation that we have to take into account when applying the switching regression with unknown sample separation, besides converging issues, is that the standard errors can only be approximated. We therefore interpret the results based on this smaller standard error in the model with endogenous switching only as a robustness check.[15]

[15] Furthermore, we re-estimated the consumption growth equation by constructing quarterly observations from the monthly waves for 2002-2004 as a robustness check. We ended up with 374 pseudo panel observations, and the OLS results turn our to be not significantly different from the ones shown in Table 2: $\widehat{\phi} = 0.643\,(0.083)$ and $\widehat{\psi} = 0.265\,(0.064)$. However, the switching regression analysis could not have been applied here due to non-convergence of the EM algorithm.

In an alternative specification, we extend our measure for consumption from non-durable to total consumption. We test whether there are differences in the reaction to income changes if we allow consumption to consist of non-durable commodities as well as a constructed depreciation for durable consumption goods, in terms of user costs. In Table 3, the results for the consumption growth equation on *total* consumption are shown.

Table 3: Marginal Effects for the Consumption Growth Equation (**Total** Cons.)

	OLS	**Switching Regression**	
Dep. var.: $\Delta \ln(c_{it})$	Pooled Model	Regime 1 (Constrained)	Regime 2 (Unconstrained)
Permanent shock (ϕ)	0.648***	0.610***	0.660***
	(0.058)	(0.060)	(0.065)
Transitory shock (ψ)	0.246***	0.410***	0.071
	(0.043)	(0.035)	(0.053)
Probability of regime 1 (P_{it})	-	0.435	
N (cells)	748	748	748
R^2	0.303	0.487	0.593
Tests (χ^2-statistic):			
$\phi = \psi$	38.18	12.10	58.46
$\phi = 1$	36.97	42.99	27.44
$\psi = \kappa = 0.05$	20.30	108.12	0.15
$\phi_1 = \phi_2$	-	0.32	
$\psi_1 = \psi_2$	-	28.39	
$\phi_{SWR} = \phi_{OLS}$	-	0.21	0.02
$\psi_{SWR} = \psi_{OLS}$	-	8.73	6.47

Notes: See Table B.3 in Appendix B for complete estimation results and full list of covariates. Significance levels: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$; robust standard errors in parentheses, adjusted for clustering, not adjusted for two-step estimation.

Source: Own calculations using the LWR data (2002-2007), provided by the FDZ.

Although the point estimates slightly differ, we find no significant differences in the coefficients compared to the analysis on non-durables. The point estimate in the unconstrained regime for the permanent shock is a bit smaller with 0.660 and estimated more robustly. Now, as in the model for non-durables with endogenous switching, it is significantly different from 1 (27.44) and the PIH would be rejected even for the unconstrained households. The transitory reaction is with an MPC of 0.071 found to be only slightly greater. Importantly, the difference in the effect on transitory income between the two regimes is actually significant at the 1%

level (6.47) because of more efficient estimates.[16]

This is the central finding of our analysis. On the one hand, the reaction of average households to non-anticipated changes in income do not correspond to the PIH. Their reaction to permanent shocks is lower than theory would predict, and transitory shocks are perceived more sensitively than the model would tell. On the other hand, two groups could have been identified according to indicators for presence of liquidity constraints. Households in the group that is identified as constrained react significantly stronger to transitory shocks to income than households in the unconstrained group. The findings for the unconstrained regime tend to confirm the validity of the PIH for this group. The results are in line with findings from the relevant literature, where rejections of the PIH are interpreted in terms of liquidity constraints (see inter alia Blundell et al., 2008b).

Looking at the distribution of the probability for being in the *un*constrained regime across the pseudo panel clusters, we see above-average probabilities for the older birth cohorts (<1943 high skilled men, <1948 high skilled women, and <1942 low skilled women), except for the oldest birth cohorts of low skilled men (<1946). The latter group in fact also face the highest probabilities compared to the younger birth cohorts within this group, but they are only at about the overall mean of 57.8%. The probability for high skilled men in the oldest birth cohorts, averaged over all time periods, is about 69%; for high skilled women, it is 76%, and for low skilled women 66%. Among the oldest birth cohorts, conditional on gender, the high skilled birth cohorts have clearly higher probabilities to be unconstrained than the lower ones, as expected, but this relationship does not hold over all birth cohorts. Interestingly, women tend to have slightly higher probabilities than men in the oldest birth cohorts, and the youngest birth cohorts (>=1963 for men and >=1959 for women), surprisingly, never have the lowest probabilities within a skill level.[17]

We find our results for the consumption growth equation to be robust to a couple of alternative specifications. In the regime equation, an initial guess is needed to determine the two regimes according to the presence of liquidity constraints. This initial guess has been made for three indicators of liquidity constraints, namely the unemployment rate in the cell, the ratio of financial wealth to permanent income, and the ratio of loan repayments to the level of outstanding debt (also see Section 2). We have re-estimated the model for several alternative guesses on these indicators: the 25%, 50%, and 75% quantiles of the unemployment rate, respectively of the wealth ratio, and respectively of the loan repayment ratio. The results were robust for all these nine estimations.

[16] Again, we also estimated the model with endogenous selection and do not find any significant differences.

[17] Another relevant factor determining the probability of being unconstrained, which has not been separated from the probabilities interpreted here, is household composition.

As it has already been mentioned in the introduction of the model (Section 2), due to its construction, the effect of the transitory shock (ϕ) potentially varies by age of the household head. Therefore, we have estimated an interaction switching regression to test the hypothesis that, if the PIH holds, the relevance of transitory shocks should increase in age. This is because transitory shocks gain more weight related to the lifetime income with increasing age. Table 4 shows the results for the interaction model with total consumption as dependent variable.[18]

Table 4: Effects for Consumption Growth with **Age Interaction** (Total consumption)

	OLS	**Switching Regression**	
Dep. var.: $\Delta \ln(c_{it})$	Pooled Model	Regime 1 (Constrained)	Regime 2 (Unconstrained)
Permanent shock (ϕ)	0.646***	0.609***	0.692***
	(0.061)	(0.062)	(0.069)
Transitory shock (ψ)	0.296*	0.480***	−0.339**
	(0.152)	(0.158)	(0.149)
Age of household head	−0.000	0.000*	0.000
	(0.000)	(0.000)	(0.000)
Transitory shock * Age	−0.001	−0.001	0.008**
	(0.003)	(0.003)	(0.003)
Quarter dummies	Yes	Yes	Yes
Year dummies	Yes	Yes	Yes
Constant	−0.043***	−0.062***	−0.121***
	(0.004)	(0.012)	(0.010)
Probability of regime 1 (P_{it})	–	0.433	
N (clusters)	748	748	748
R^2	0.303	0.481	0.602

Notes: Significance levels: $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$; robust standard errors in parentheses, adjusted for clustering, not adjusted for two-step estimation.

Source: Own calculations using the LWR data (2002-2007), provided by the FDZ.

In the OLS equation, the interaction term for age is not significantly different from zero, while the effect on transitory income is slightly higher than in Table 3. The switching regression reveals an interesting result. While there is no age varying effect found for the constrained households, we find a positive significant interaction term for the unconstrained regime. This means that the transitory income shocks become more important with higher age. This result is consistent with the PIH. Evaluating the marginal effect $\widehat{\psi}$ at the lowest cluster mean age of

[18] We could not check the results for non-durable consumption because the switching regression computation was not feasible for this model due to non-convergence of the EM algorithm.

36.6 gives still a negative but insignificant effect of -0.037 (standard error of 0.054), which is due to the linear unrestricted specification. But at the age of 55, we get a significant effect of 0.115 (0.057) and at the statutory retirement age of 65, the effect is 0.197 (0.079). Following the strict PIH, the reaction to transitory shocks should approach 1 when agents reach the end of the life-cycle. However, there are several explanations for an MPC smaller than 1 at old ages. The exact length of the life-cycle is in fact unknown to the households and there are several motives for savings that become increasingly relevant towards the end of the life-cycle, such as the bequest motive, that give reasonable explanations and that have been supported with evidence in the literature (Lusardi, 1997).

Results for the Excess Sensitivity Test

The results from the consumption growth equation suggest an appearance of two kinds of households. The first ones are unable to smooth consumption over the life-cycle and depend on changes in transitory income. For them, we thus observe reactions to non-anticipated changes in transitory income. However, the other ones do not respond to non-anticipated changes in transitory income, but also do not fully consume non-anticipated increases in permanent income.

Another question related to the PIH is whether households respond to income changes that are *anticipated*. We would expect that for a liquidity constrained household the realization of an income change matters, rather than its anticipation. Table (5) presents the main results of estimating the Euler equation in Eq. (11) by OLS and by the switching regression approach. In the right column, for means of robustness check, we show OLS results for the model based on non-durable consumption.[19]

In the pooled model, the coefficient δ is estimated at -0.019 and significantly different from zero on the 5% level. This result suggests excess sensitivity to anticipated income for the whole sample. An estimate for δ that is significantly lower than zero is not consistent with the PIH. It is, however, consistent with the empirical evidence on excess sensitivity in the relevant literature (e.g. Zeldes, 1989a). It implies that households do respond to changes in income that have been anticipated. The negative coefficient implies that consumption increases at a slower rate than income has increased in the previous period because anticipated income changes have been proxied by observed income (also see Section 2).

Two regimes are again identified through the selection covariates, where only the ratio of financial assets to permanent income has been applied to identify liquidity constraints, besides

[19] In the switching regression, we do not find two stable regimes for non-durable consumption.

Table 5: Marginal Effects for the **Excess Sensitivity** Test (Total Cons.)

Dep. var.: $\Delta \ln(c_{it+1})$	**OLS** Pooled Model	**Switching Regression** Regime 1 (Constr.)	Regime 2 (Unconstr.)	**OLS** Pooled Model (**Nondur.** Cons.)
Anticipated Income[a] (δ)	-0.019^{**}	-0.039^{***}	-0.015	-0.014^{*}
	(0.008)	(0.008)	(0.010)	(0.008)
$\Delta\text{adults}_{it+1}$	0.210^{***}	0.244^{***}	0.208^{***}	0.276^{***}
	(0.027)	(0.027)	(0.031)	(0.026)
Δkids_{it+1}	0.093^{***}	0.091^{***}	0.085^{***}	0.081^{***}
	(0.024)	(0.031)	(0.028)	(0.027)
Probability of regime 1 (P_{it})	-	0.482		-
N (cells)	799	782[b]	782[b]	799
R^2	0.191	0.597	0.223	0.154
Tests (χ^2-statistic):				
$\delta_1 = \delta_2$	-	3.87		-
$\delta_{SWR} = \delta_{OLS}$	-	3.44	0.09	-

Notes: See Table B.5 and Table 4 in Appendix B for complete estimation results and full list of covariates. Significance levels: $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$; robust standard errors in parentheses, adjusted for clustering, not adjusted for two-step estimation.
[a]: The anticipated income term $E_{it}[\Delta \ln(y_{it+1})]$ is approximated by $\ln(y_{it})$.
[b]: Some 17 observations are dropped from the EM algorithm two-step procedure due to a zero in the denominator of Eq. (9).

Source: Own calculations using the LWR data (2002-2007), provided by the FDZ.

the usual socio-demographics.[20] While the first regime is found to react more strongly to anticipated income changes than in the pooled model, the reaction for the second regime is slightly smaller, and it is estimated with a greater standard error.[21]

This is the central finding of this additional test for excess sensitivity, and it supports the excess sensitivity hypothesis. Households that are identified to be liquidity constrained are found to respond more strongly to changes in *anticipated* income changes than households that are not liquidity constrained. This result is also in line with the findings from the relevant literature on excess sensitivity (see e.g. Zeldes, 1989a; Garcia et al., 1997; Jappelli et al., 1998).

The estimated effects for the control variables of changing household composition suggest that an additional adult in the household increases consumption by around 20% while an additional child increases consumption only by around 9%. These results vary only slightly over the regimes, suggesting a slightly higher reaction among constrained households. The

[20]This is done to avoid endogeneity issues by having the current and lagged income in both the selection and the main equation. Another reason is to have a stable switching regression.

[21]We find slightly different coefficients for some alternative initial guesses. However, these do not affect the validity of these results.

hypothesis of equality of the excess sensitivity coefficients over the regimes is marginally rejected on the 5% level (χ^2-statistic = 3.87). Furthermore, equality compared to OLS cannot be rejected at the 5% level, for both regimes. Further considering the OLS results for non-durable consumption, there is less evidence for excess sensitivity than in the unconstrained regime for total consumption. We also checked the response on only food consumption and found an even smaller and insignificant coefficient. This could indicate that consumption reactions to anticipated income changes do not affect convenience goods or necessities.

5. Conclusion

We have analyzed empirically the relevance of the permanent income hypothesis (PIH) in German consumption survey data. We found evidence for deviation from theory predictions and have investigated to which extent these deviations from the PIH can be traced back to the presence of liquidity constraints in household consumption. We made use of a pseudo panel constructed on repeated cross-sections of consumption survey data for Germany to investigate the consumption effects of income shocks in the context of liquidity constraints. This data set has proven to be rich in the sense that it provides relatively precise measures of the individual income and consumption dynamics, for durable and non-durable consumption, from which we have utilized their joint evolution over time to disentangle the consumption effects of income shocks into transitory and permanent elements. In a switching regression approach with unknown sample separation we have identified households that can be assumed to be affected by liquidity constraints and those that seem to be rather unconstrained.

We find that for households in the constrained regime, reactions to changes in transitory income are significantly greater than for households in the unconstrained regime. We contribute to the literature evidence on liquidity constraints, based on a pseudo panel of rich German consumption survey data, which has not been exploited for Germany so far in this context to the best of our knowledge.

We find, on the one hand, that households' responses to non-anticipated changes in income are at odds with the PIH. Their reaction to permanent shocks is lower than theory predicts and transitory shocks are perceived more sensitively than the model would tell. On the other hand, we have identified two groups according to indicators for presence of liquidity constraints. Households identified as constrained react significantly stronger to transitory income shocks than households in the unconstrained group. These results are in line with findings from the relevant literature, where relevance of liquidity constraints has been found. These results have been found to be robust with respect to various model specifications as well as different consumption concepts.

Furthermore, we find evidence for excess sensitivity to anticipated income changes for households in the constrained regime if total consumption, durable as well as non-durable, is considered. Households that are identified to be liquidity constrained are found to respond more strongly to *anticipated* income changes than households that are not liquidity constrained. We conclude that there seems to be a different reaction to anticipated income changes due to the presence of liquidity constraints among the two groups, at least if one considers total consumption, but the two different types of households have proven to be more difficult to identify.

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A. Appendix - Data and Definition of Income and Savings

Data

For the LWR consumption data, households are recruited voluntarily for reports every year, according to stratified quota samples from Germany's current population survey (Mikrozensus), and report for a time of four months (one month out of each quarter of the year). Since 2005, recruited households stem from a subsample of the Income and Consumption Survey for Germany (Einkommens- und Verbrauchsstichprobe, EVS). They are aggregated to the population according to a marginal distribution of demographic variables. The entire population covered by the LWR is restricted, as there are groups that are not covered: self-employed, institutionalized people (i.e. military people in caserns, students in dormitories, elderly and disabled people in nursery homes or hospitals, nurses or migrant workers in residences, people in jails), homeless people, and households with monthly net household income greater than 18,000 euros. When descriptive statistics on the LWR data are presented (see Section 3), data are weighted by population weights. Population weights for the LWR are constructed w.r.t. the marginal distribution of households in the Mikrozensus-population by strata of household composition, social status, and net household income. For further details on the LWR data, see Statistisches Bundesamt (2007).

Treatment of Durables

The investment character of the consumption of durables goods is accounted for by calculating user costs or depreciation rates for these goods for current consumption, and the residual of actual expenditures and user costs is interpreted as savings. For most of the "relevant" durable goods, user costs are computed by mean imputation. A durable good is considered "relevant" if nearly every household can be assumed to consume at least a small amount of the good every period and the macroeconomic expenses on that good are above an arbitrary threshold. These goods include e.g. furniture, electric devices, entertainment electronics, clothes, shoes, and carpets.

In performing the mean imputation, household clusters are constructed depending on six age groups, seven income groups, and six household types. Then, the expenditures are summed up for a durable good in each cluster, and the sum is reallocated equally among all observations in the cluster. Afterwards, an estimated quarter effect is added to every adjusted category of expenditure to avoid a bias in the quarter dummies of the main equation. This is necessary because non-durable consumption is not adjusted for quarter effects.

Expenditures for car purchases form the most significant durable good related to the macroeconomic expenditures, except for housing expenditures. Cars have been treated a little differently from the described mean imputation. Firstly, a tobit-regression is estimated for households owning exactly one car with the reported expenditures for leasing as dependent variable and the disposable income and household characteristics as explanatory variables. Then, the unconditional value is predicted for each household owning at least one car assuming that 90% of the leasing rate is depreciation and 10% is interest payment.[22] The depreciation is calibrated dependent on the number of cars in the household and their characteristics (newly or second-hand bought). If the household reports expenditures for car purchases, 15% of this value is taken directly as depreciation for the first year (5% in case of second-hand purchase). Furthermore, if there are expenditures reported for preventive maintenance or spare parts then these are taken into account in calculating the depreciation. Finally, it has been guaranteed that the population-aggregate sum of expenditures for all the relevant durable goods is roughly conserved after adjustment.[23]

Following Garner and Verbrugge (2009), the market rental value approach has also been applied to the measurement of services from owner-occupied housing. For owner-occupied housing, rents that are provided with the data have been applied and imputed both in current income as well as in consumption. The rents applied have been computed by the Federal Statistical Office as follows: an average gross rent (excluding heating and maintenance) per square meter differentiated by federal states is applied to the reported size of the house or flat, and this is added to the reported expenditures for heating and maintenance (Statistisches Bundesamt, 2005).

[22] In case positive leasing payments are reported, they are applied here.

[23] On arguments for this market rental value approach for the measurement of services from durables, see Garner and Verbrugge (2009). For a survey on various approaches for the measurement of durable service flows, see Katz (1983).

B. Appendix - Results

Table B.1: Estimation Results for the **Income** Equation

Dep. var.: $\ln(y_{it})$	**Coefficient**	**Standard Error**
Age of household head	0.0013	(0.0175)
Age squared of household head	−0.0000	(0.0002)
Interactions:		
Singles × age (ref.)		
Single parents × age	0.0377**	(0.0173)
Couples, no kids × age	0.0136**	(0.0054)
Couples, one kid × age	0.0304**	(0.0120)
Couples, two and more kids × age	0.0255**	(0.0118)
Large households[a] × age	−0.0074	(0.0082)
Singles × age^2 (ref.)		
Single parents × age^2	−0.0007*	(0.0004)
Couples, no kids × age^2	−0.0002**	(0.0001)
Couples, one kid × age^2	−0.0005*	(0.0002)
Couples, two and more kids × age^2	−0.0004*	(0.0002)
Large households[a] × age^2	0.0002	(0.0001)
Unemployed × age (ref.)		
Civil servants × age	0.0471***	(0.0109)
White collar × age	0.0362***	(0.0085)
Blue collar × age	0.0158	(0.0107)
Pensioners × age	−0.0038	(0.0091)
Civil pensioners × age	0.0166	(0.0124)
Unemployed × age^2 (ref.)		
Civil servants × age^2	−0.0006***	(0.0002)
White collar × age^2	−0.0004***	(0.0002)
Blue collar × age^2	−0.0001	(0.0002)
Pensioners × age^2	0.0001	(0.0001)
Civil pensioners × age^2	−0.0001	(0.0002)
Other Effects:		
High skill (ref.)		
Medium skill	−0.2268***	(0.0632)
Low skill	−0.3067***	(0.1119)
Number of adults in household	0.2314***	(0.0827)
Year effects	yes	
Quarter effects	yes	
Fixed effects	yes	
ρ	-0.098	
N (cells)	799	
R^2	0.567	

Notes: Significance levels: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$; robust standard errors in parentheses.
[a]: The group "large households" is the residual group of all remaining households. It mainly consists of households with more than two adults.

Source: Own calculations using the LWR data (2002-2007), provided by the FDZ.

Table B.2: Estimation Results for the Consumption Growth Eq. (**Non-durable** Cons.)

	OLS	**Switching Regression**	
	Pooled Model	Regime 1	Regime 2
Dep. var.: $\Delta \ln(c_{it})$		(Constrained)	(Unconstrained)
Permanent shock (ϕ)	0.667***	0.626***	0.787***
	(0.058)	(0.055)	(0.133)
Transitory shock (ψ)	0.204***	0.427***	0.048
	(0.046)	(0.040)	(0.049)
Age of household head	−0.000	−0.000	−0.001***
	(0.000)	(0.000)	(0.000)
Year 2002	0.002	ref.	0.035***
	(0.003)		(0.008)
Year 2003	0.006**	−0.005	0.055***
	(0.002)	(0.009)	(0.008)
Year 2004	0.003	−0.021***	0.051***
	(0.002)	(0.007)	(0.009)
Year 2005	0.019***	−0.021***	0.033***
	(0.003)	(0.007)	(0.008)
Year 2006	0.007**	−0.032***	0.050***
	(0.003)	(0.005)	(0.006)
Year 2007	ref.	−0.044***	ref.
		(0.006)	
Quarter 1 (ref.)			
Quarter 2	0.016***	0.052***	0.037***
	(0.005)	(0.012)	(0.007)
Quarter 3	0.008**	0.005	0.033***
	(0.003)	(0.005)	(0.009)
Quarter 4	−0.019***	−0.053***	0.028***
	(0.006)	(0.010)	(0.007)
Constant	−0.005	0.043***	−0.059***
	(0.003)	(0.009)	(0.010)
Probability of regime 1 (P_{it})	–	0.422	
N (clusters)	748	748	748
R^2	0.215	0.586	0.146

Notes: Significance levels: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$; robust standard errors in parentheses, adjusted for clustering, not adjusted for two-step estimation.

Source: Own calculations using the LWR data (2002-2007), provided by the FDZ.

Table B.3: Estimation Results for the Consumption Growth Equation (**Total** Cons.)

	OLS	**Switching Regression**	
	Pooled Model	Regime 1	Regime 2
Dep. var.: $\Delta \ln(c_{it})$		(Constrained)	(Unconstrained)
Permanent shock (ϕ)	0.648***	0.610***	0.660***
	(0.058)	(0.060)	(0.065)
Transitory shock (ψ)	0.246***	0.410***	0.071
	(0.043)	(0.035)	(0.053)
Age of household head	−0.000	0.000**	0.000
	(0.000)	(0.000)	(0.000)
Year 2002	0.004*	0.031***	−0.066***
	(0.002)	(0.009)	(0.004)
Year 2003	0.011***	0.051***	−0.050***
	(0.002)	(0.009)	(0.005)
Year 2004	0.010***	0.051***	−0.059***
	(0.002)	(0.008)	(0.005)
Year 2005	0.008**	0.012*	−0.007*
	(0.003)	(0.007)	(0.004)
Year 2006	0.016***	0.017**	ref.
	(0.003)	(0.008)	
Year 2007	ref.	ref.	−0.016***
			(0.004)
Quarter 1 (ref.)			
Quarter 2	0.057***	0.048***	0.166***
	(0.005)	(0.007)	(0.008)
Quarter 3	0.027***	0.023***	0.089***
	(0.004)	(0.005)	(0.005)
Quarter 4	0.055***	0.068***	0.175***
	(0.005)	(0.010)	(0.006)
Constant	−0.043***	−0.075***	−0.116***
	(0.004)	(0.013)	(0.012)
Probability of regime 1 (P_{it})	–	0.435	
N (clusters)	748	748	748
R^2	0.303	0.487	0.593

Notes: Significance levels: $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$; robust standard errors in parentheses, adjusted for clustering, not adjusted for two-step estimation.

Source: Own calculations using the LWR data (2002-2007), provided by the FDZ.

Table B.4: Estimation Results for the Consumption Growth Equation (Non-durable Cons. with **Endogenous** Selection)

	OLS	**Switching Regression**	
	Pooled Model	Regime 1	Regime 2
Dep. var.: $\Delta \ln(c_{it})$		(Constrained)	(Unconstrained)
Permanent shock (ϕ)	0.667***	0.608***	0.783***
	(0.058)	(0.058)	(0.077)
Transitory shock (ψ)	0.204***	0.401***	0.056
	(0.046)	(0.044)	(0.059)
Age of household head	−0.000	−0.000	−0.000***
	(0.000)	(0.000)	(0.000)
Year 2002	0.002	0.013	0.054**
	(0.003)	(0.010)	(0.025)
Year 2003	0.006**	ref.	0.080***
	(0.002)		(0.026)
Year 2004	0.003	−0.009	0.072**
	(0.002)	(0.009)	(0.025)
Year 2005	0.019***	−0.000	0.058***
	(0.003)	(0.015)	(0.020)
Year 2006	0.007**	−0.014	0.069***
	(0.003)	(0.014)	(0.020)
Year 2007	ref.	−0.021	ref.
		(0.017)	
Quarter 1 (ref.)			
Quarter 2	0.016***	0.044**	0.090***
	(0.005)	(0.016)	(0.012)
Quarter 3	0.008**	0.000	0.065***
	(0.003)	(0.006)	(0.011)
Quarter 4	−0.019***	−0.047**	0.080***
	(0.006)	(0.016)	(0.012)
Selection Term	–	−0.032***	0.014
		(0.008)	(0.009)
Constant	−0.005	0.032**	−0.135***
	(0.003)	(0.008)	(0.032)
Probability of regime 1 (P_{it})	–	0.395	
N (clusters)	748	748	748
R^2	0.215	0.564	0.212

Notes: Significance levels: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$; robust standard errors in parentheses, adjusted for clustering, not adjusted for two-step estimation.

Source: Own calculations using the LWR data (2002-2007), provided by the FDZ.

Table B.5: Estimation Results for the Excess Sensitivity Test (**Total** Cons.)

	OLS	**Switching Regression**	
Dep. var.: $\Delta \ln(c_{it+1})$	Pooled Model	Regime 1 (Constrained)	Regime 2 (Unconstrained)
Anticipated Income δ	-0.019^{**}	-0.039^{***}	-0.015
	(0.008)	(0.008)	(0.010)
Age of household head	-0.000	0.000^{*}	0.001^{**}
	(0.000)	(0.000)	(0.000)
Year 2002	-0.000	-0.068^{***}	0.024^{***}
	(0.008)	(0.009)	(0.008)
Year 2003	-0.007	-0.048^{***}	0.019^{**}
	(0.009)	(0.008)	(0.008)
Year 2004	-0.006	-0.058^{***}	0.031^{***}
	(0.009)	(0.009)	(0.009)
Year 2005	-0.002	ref.	-0.010
	(0.011)		(0.010)
Year 2006	ref.	0.001	-0.008
		(0.009)	(0.011)
Year 2007	-0.016	-0.013	ref.
	(0.010)	(0.009)	
Quarter 1 (ref.)			
Quarter 2	0.043^{***}	0.152^{***}	0.029^{***}
	(0.007)	(0.008)	(0.009)
Quarter 3	0.012^{*}	0.066^{***}	-0.005
	(0.007)	(0.009)	(0.006)
Quarter 4	0.054^{***}	0.167^{***}	0.033^{***}
	(0.007)	(0.008)	(0.008)
$\Delta \text{adults}_{it+1}$	0.210^{***}	0.244^{***}	0.208^{***}
	(0.027)	(0.027)	(0.031)
$\Delta \text{kids}_{it+1}$	0.093^{***}	0.091^{***}	0.085^{***}
	(0.024)	(0.031)	(0.028)
Constant	0.160^{**}	0.271^{***}	0.088
	(0.079)	(0.076)	(0.095)
Probability of regime 1 (P_{it})	–	0.482	
N (clusters)	799	782	782
R^2	0.191	0.597	0.223

Notes: Significance levels: $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$; robust standard errors in parentheses, adjusted for clustering, not adjusted for two-step estimation.

Source: Own calculations using the LWR data (2002-2007), provided by the FDZ.

Table B.6: Estimation Results for the Excess Sensitivity Test (**Non-durable** Cons.)

	OLS
	Pooled Model
Dep. var.: $\Delta \ln(c_{it+1})$	
Anticipated Income δ	-0.014^{**}
	(0.008)
Age of household head	-0.000
	(0.000)
Year 2002	0.001
	(0.009)
Year 2003	-0.002
	(0.009)
Year 2004	-0.003
	(0.009)
Year 2005	0.020^{*}
	(0.011)
Year 2006	ref.
Year 2007	-0.006
	(0.011)
Quarter 1 (ref.)	
Quarter 2	0.011^{*}
	(0.006)
Quarter 3	0.002
	(0.006)
Quarter 4	-0.014^{**}
	(0.006)
$\Delta \text{adults}_{it+1}$	0.276^{***}
	(0.026)
$\Delta \text{kids}_{it+1}$	0.081^{***}
	(0.027)
Constant	0.140^{*}
	(0.076)
Probability of regime 1 (P_{it})	-
N (clusters)	799
R^2	0.154

Notes: Significance levels: $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$; robust standard errors in parentheses, adjusted for clustering, not adjusted for two-step estimation.

Source: Own calculations using the LWR data (2002-2007), provided by the FDZ.