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# Heterogeneity in Regional Consumption due to Current Income Shocks: Testing the Permanent Income Hypothesis for German Panel Data

Timo Mitze\*

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## Abstract

In this paper we analyse the role of current income shocks on regional consumption patterns for a panel of German regions (NUTS1 level) between 1970 and 2007. The basic workhorse model of the analysis is the Permanent Income Hypothesis, which has recently been subject to new methodological improvements regarding the specification of expectation formation. Using both a short-run as well as cointegration perspective we tackle earlier findings for German regional data, which find a significant degree of "excess sensitivity" of consumption to income shocks. However, our results do not support these findings. Instead, in the short-run approach the influence of past and current income changes turns out to be significant if we control for potentially omitted variables. Using also long-run information in a Panel Error Correction Model (ECM) approach we indeed find a significant share of liquidity constrained households. However its share is below the fraction recently found for German regional data. These results may give new insights with respect to the effectiveness of policy measures for short-term consumption stabilization as being raised in the policy debate to combat the global financial and economic crisis (e.g. so-called consumption vouchers or tax rebates). Additionally, by testing for the asymmetry in the consumption patterns for German regions both in the short and long run, we are able to give an initial answer to the question dealing with the regional effects of the world economic and financial crisis. Here we apply different testing approaches for slope homogeneity of the regions in the panel and aim to identify different long- and short-run adjustment regimes in Germany. We finally also account for the likely role of spatial autocorrelation when dealing with territorial data.

*JEL:* E21, C23

*Keywords:* Consumption, PIH, Panel ECM

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# 1 Introduction

Though immense research efforts have been devoted to the analysis of the likely sources and impacts of the global economic and financial crises, little is still known about the distribution of its regional effects. One important question is whether current income shocks hit regions symmetrically or asymmetrically. In this paper we take a closer look at the general role of current income shocks in explaining the regional patterns of private consumption spending. Although due to data restriction we do not have any post-crisis data at the regional level yet, a structural analysis with historical data may provide guidance. We take the Permanent Income Hypothesis (henceforth PIH) as a basic workhorse model and test whether German consumers significantly depart from the model's predictions in terms of adapting the level of consumption spending to past and current income shocks. This may also guide policymakers with respect to the effectiveness of policy measures intended to stabilize short term consumption. Though ultimately not implemented by the German government, one prominently advocated option to combat the economic breakdown is to propose the distribution of so-called consumption or tax vouchers (rebates) in the spirit of earlier policy experiments in the USA and Japan (see e.g. Seidman, 2003). According to macroeconomic theory, for such a temporary expenditure shock to work, consumers (or at least a large part of them) need to be either myopic or liquidity constrained.

Recent empirical findings on German (regional) consumption data indeed found a substantial degree of "excess sensitivity" to income, which can be associated with liquidity constrained consumers (see Dreger & Kosfeld, 2003, and DeJuan et al., 2006). This latter result also implies a rejection of the PIH, which rests on the assumption that representative agents plan consumption expenditures on the basis of their lifetime income expectations rather than period-by-period income. According to the PIH, agents should therefore not react to temporary income shocks if their expectations about lifetime (or permanent) income remain unchanged. In the empirical literature therefore a huge stock of testing approaches for the appropriateness of the Permanent Income model with rational expectations has been built up. Besides the problem of liquidity constraints the notion of loss aversion and myopic consumers has also been raised (see e.g. van Treeck, 2008, for a recent overview). Likewise habit formation, rule-of-thumb consumers or social norms may motivate a deviation from the strong rational expectation assumption of the PIH framework.

In this paper we therefore tackle the above empirical findings for Germany in light of new theoretical and methodological work on the PIH approach. As argued we explicitly adopt a regional perspective and use German state-level data for the period 1970 (or 1991)

to 2007. Starting with a short-run approach as the workhorse model in the empirical PIH literature (see e.g. Hall, 1978, Flavin, 1981, Campell & Mankiw, 1990), we do not find evidence for "excess sensitivity" and "liquidity constraints" in the context of a dynamic panel data model. Our preferred specification relates changes in consumption to a "surprise" term in permanent income, proxied by the residual of an autoregressive income process, as well as past values of consumption growth. Following Malley & Molana (2006) we can interpret this augmented specification as solution to a optimization problem with habit persistence.

Next to the short run approach typically used in testing for the validity of the PIH framework we also combine the long- and short-run perspective in a panel cointegration model. A stable long-run cointegration relationship between (real) income and consumption is a necessary (but not sufficient) condition for the validity of the PIH. Using panel cointegration tests in the framework of Westerlund (2007) we could clearly reject the null hypothesis of no cointegration between income and consumption. We are therefore able to specify a Panel ECM and look at the short run adjustment coefficients in the consumption equation to judge the share of liquidity constraint households. Compared to the short-run approach here we find at the aggregate level a significant but much lower fraction of constrained agents as in Dreger & Korfeld (2003). Our findings thus come close to earlier results with respect to German data in Wolters (1992) and for the US in Fuhrer (2000).

Since we employ different estimators for the Panel ECM including dynamic fixed effects (DFE), mean group (MG) and pooled mean group (PMG) estimation we are also able to check for the asymmetry in the income-consumption path for German regions both with respect to the long- and short-run adjustment dynamics. From a methodological point of view testing for the asymmetry of the different cross-sections (regions) in focus boils down to the question of equal slope coefficients in the short- and long-run coefficients. The design of the MG, PMG and DFE estimators allows for sequential testing for the validity of different cross-section restrictions. That is, starting from the consistent but potentially inefficient MG estimator we first test for the equality of the long-term coefficients using a standard Hausman test. Both for West Germany as well as total Germany our Hausman test statistic does not reject the null hypothesis of equal coefficient signs for the different states and thus a homogeneous long run co-integration path - both for the PMG as well as the DFE.

However, when testing also for short-run equality in the latter two restricted panel data models, different versions of the F-Test (or accordingly Wald tests) for poolability - namely the standard Chow and Roy-Zellner type test in its asymptotic normal as well as

bootstrapped version - reject slope homogeneity especially for the total German sample since 1991. Aiming to get a deeper understanding of the heterogeneity of short-run adjustment of German states in their regional consumption path to a common long-run solution, we additionally borrow a testing approach from the "club convergence"-literature (e.g. in Phillips & Sul, 2007). Here we apply a clustering algorithm for club identification based on a series of F-tests which starts from a cross-section ordering of estimated short-run coefficients and sieves the data for club members starting from the formation of an initial core group. This allows us to identify different short-run regimes within Germany. Since spatial income and consumption correlations may be in order at the regional level, we run the same type of testing approach based on spatially filtered variables. The latter may thus be seen as a robustness analysis to the aspatial benchmark estimation.

The remainder of the paper is organised as follows: In section 2 we briefly review the theoretical underpinnings of the PIH framework and derives testable empirical specifications. Section 3 presents the data and some stylized facts of the income-consumption linkage. In this section also the time-series properties of the variables in focus are tested. Section 4 presents the empirical estimation strategy, both focusing on a short run specification as well as modelling a cointegrated consumption equation. Section 5 puts special emphasis on testing for long-run and short-run slope heterogeneity. Section 6 presents some sensitivity analyses, especially with respect to spatial autocorrelation in the data. Section 7 finally concludes.

## 2 The Permanent Income Hypothesis

The Permanent Income Hypothesis models the optimal intertemporal behaviour of a representative agent with an infinite time horizon.<sup>1</sup> and was first proposed by Milton Friedman (1957) to establish a micro founded relationship between income and consumption. The main innovation to earlier consumption models such as the 'absolute' and 'relative' income hypotheses is that agents are assumed to plan expenditures on the basis of lifetime income expectations rather than income received period-by-period. That is, using a discrete time framework for any period  $t$  the agent chooses  $C_{t+j}$  for all  $j \geq 0$  to maximize the expected value of objective function  $E_t[U_t]$  with:

$$U_t = U(C_t, C_{t+1}, \dots, C_{t+j}, \dots), \quad (1)$$

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<sup>1</sup>Closely related, the Life Cycle Hypothesis (LCH) assumes that individuals consume a constant percentage of the present value of their life income, where the latter is based on a finite lifetime perspective

subject to a sequence of budget constraints (again  $j \geq 0$ )

$$W_{t+j+1} = (1 + r_{t+j})W_{t+j} + Y_{t+j} - C_{t+j}. \quad (2)$$

$W$  is the real value of the stock of non-human wealth,  $r$  is real (after tax) interest rate and  $Y$  is real (after tax) labour income. As Malley & Molana (2006) point out, the solution to this problem yields a smoothing rule for the expected marginal utility of consumption

$$E_t \left( \frac{\partial U_t}{\partial C_{t+j+1}} - (1 + r_{t+j}) \frac{\partial U_t}{\partial C_{t+j}} \right) = 0 \quad (3)$$

for  $j \leq 1$ . Given that the underlying utility function is time separable and agents assume a constant real interest rate to discount both future income and future utility of consumption eq.(3) implies that agent's expected consumption remains constant over times as

$$E_t C_{t+j} = C_t = Y_t^P, \quad (4)$$

where  $Y_t^P$  is defined as constant annuity income stream associated with the present value of agent's human and non-human wealth. The latter variable can be derived from the budget constraint re-written in terms of its infinite lifetime version as

$$\sum_{j=0}^{\infty} \rho^{j+1} C_{t+j} = W_t + \sum_{j=0}^{\infty} \rho^{j+1} Y_{t+j} \quad (5)$$

with  $\rho = 1/(1 + r)$ . Then,

$$Y_t^P = r \left( A_t + \sum_{j=0}^{\infty} \rho^{j+1} E_t Y_{t+j} \right) = r \sum_{j=0}^{\infty} \rho^{j+1} E_t C_{t+j} \quad (6)$$

and hence

$$Y_t^P = (1/\rho)Y^{P_{t-1}} - ((1 - \rho)/\rho) C_{t-1} + V_t, \quad (7)$$

where the only revisions  $V_t$  in the previously formulated plan are due to unexpected factors affecting the agents' income as

$$V_t = r \sum_{j=0}^{\infty} \rho^{j+1} (E_t Y_{t+j} - E_{t-1} Y_{t+j}). \quad (8)$$

Following Hall (1978) and assuming rational expectations  $V_t$  will behave as an unpredictable error term with  $E_{t-1} V_t = 0$ . This rational expectation interpretation of the PIH

has been to subject to extensive empirical testing in the recent past. The underlying testable hypothesis based on a 'surprise' consumption function typically implies estimating the following two-equation system (Flavin, 1981) with variables in logs as

$$Y_t = \alpha_0 + \sum_{i=1}^k \alpha_i Y_{t-i} + \gamma t + \epsilon_t, \quad (9)$$

$$\Delta C_t = \beta_0 + \beta_1 \Delta Y_{t-1} + \beta_2 \epsilon_t + u_t \quad (10)$$

where for the first equation of the system income is assumed to be a linear stochastic Autoregressive (AR) process. In the consumption equation (in first differences) the surprise in permanent income ( $\lambda$ ) is modelled in terms of observable income  $\epsilon_t = (Y_t - E_{t-1} Y_t)$ . Accordingly,  $\beta_1$  measures excess sensitivity of changes in consumption to income changes and  $\beta_2$  is the warranted change in consumption given the proxy for innovation in the income process ( $\epsilon_t$ ). Under the Rational Expectations Hypothesis (REH) innovation in labour income is proportional to the surprise in permanent income ( $Y_t^P - E_{t-1} Y_t^P$ ). If the PIH according to the system eq.(9) to eq.(10) is valid  $\beta_1$  should be equal to zero. According to the "excess sensitivity" hypothesis  $\beta_1 \neq 0$  may reflect liquidity constraints (e.g. credit rationing). A modification of system estimation from above may include substituting eq.(10) by

$$\Delta C_t = \beta_0 + \beta_1 \Delta C_{t-1} + \beta_2 \epsilon_t + u_t, \quad (11)$$

As Fuhrer (2000) and Malley & Molana (2006) show, eq.(11) can be interpreted as the solution to a life-cycle optimisation problem with habit persistence (rule-of-thumb smoothing). In this model  $\beta_1$  measures the impact of habit formation as a gradual adjustment of current consumption to a reference value. In a very influential empirical paper Campbell & Mankiw (1990) finally propose a generalisation of the above described equation system, which allows to directly estimate the fraction of liquidity constrained households on the one hand, and those households which behave in line with the PIH on the other hand. Following Campbell & Mankiw (1991) we assume that constrained agents set  $\Delta C_t = \Delta Y_t$ . We denote  $\rho$  as proportion of agents for whom the constraints are binding, and augment the ARIMA(1,1,0) model of Malley & Molana (2006) as:

$$\Delta C_t = \beta_0 + (1 - \rho) \beta_1 \Delta C_{t-1} + (1 - \rho) \beta_2 \epsilon_t + \rho \Delta Y_t + u_t. \quad (12)$$

An alternative way to model the income-consumption system from above is to start from a cointegration perspective. As Dreger & Kosfeld (2003) point out, the PIH frame-

work implies cointegration between consumption and disposable income or a stationary saving rate. However, here one has to note that cointegration analysis can only provide a weak test for the validity of the PIH since stationarity of the saving rate is also consistent with several alternative specifications such as the Keynesian absolute income hypothesis. In this sense besides testing for a cointegrated long-run relationship between income and consumption special attention should also be devoted to the interpretation of the short-run coefficients. As discussed above, the parameter coefficient of lagged and actual income growth may be interpreted as the income share earned by liquidity constrained households. The consumption function in a cointegration perspective can be written in the form of a error correction model (ECM) as:

$$\Delta C_t = \gamma_0 + \gamma_1 \Delta C_{t-1} + \gamma_2 \Delta Y_t + \phi \xi_t + e_t, \quad (13)$$

where  $\xi$  is an estimate of the long-run cointegration relationship as

$$\xi_t = C_t - \kappa_0 + \kappa_1 Y_t \quad (14)$$

and  $\phi$  in eq.(13) is the speed of adjustment parameter. The coefficients  $\gamma_1$  and  $\gamma_2$  measure the influence of short-run movements in the dependent and exogenous variables in the error correction presentation. Using the long-run cointegration relationship can be seen as an alternative to the commonly used AR based surprise income term in eq.(11) when testing for the validity of the PIH approach. In the following we apply subsequently the short run and cointegration approach to test for liquidity constrained households in German regional data.

### 3 Database and Variable Description

For empirical estimation we use federal state-level data at the NUTS1 classification for the 10 West German states (Bundeslaender) between 1970 and 2007, as well as data for all 16 German states starting from 1991. We construct per capita time series data in real terms for consumption, GDP and disposable income. The data is taken exclusively from the German statistical office (*Volkswirtschaftliche Gesamtrechnung der Laender, VGRdL*). Before we turn to the estimation exercise, we first present some stylised facts. Over the sample consumption accounts for an average 55 % of GDP (in real per capita terms) with a minimum of 36 % for Hamburg, maximum of 65 % in Lower Saxony. As Table 1 shows annual growth rates of GDP are found to be more volatile than consumption changes for different sub-periods. Figure 1 additionally plots for each federal state the ratio of



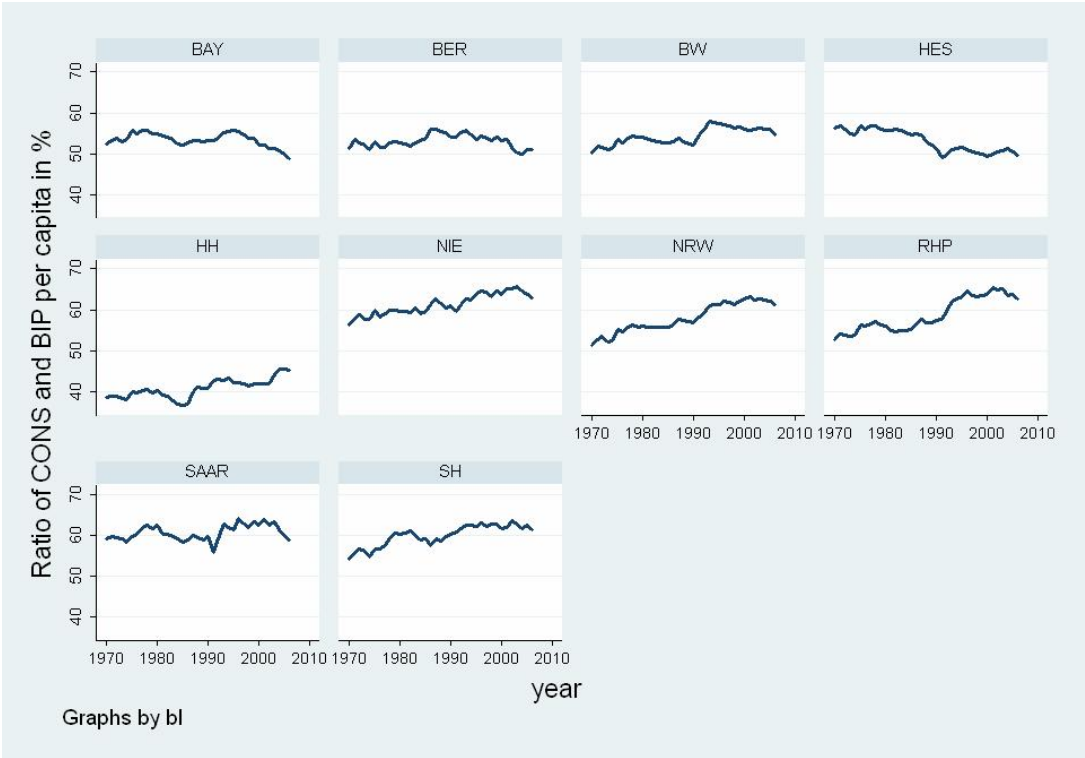
consumption and GDP in per capita levels. Figure 2 compares the annual growth rates of real per capital income and consumption.

**Table 1: Volatility of consumption and income**

|            | 1970-80 | 1981-90 | 1991-00 | 2001-07 |
|------------|---------|---------|---------|---------|
| $\Delta C$ | 0.015   | 0.017   | 0.016   | 0.013   |
| $\Delta Y$ | 0.021   | 0.018   | 0.021   | 0.015   |

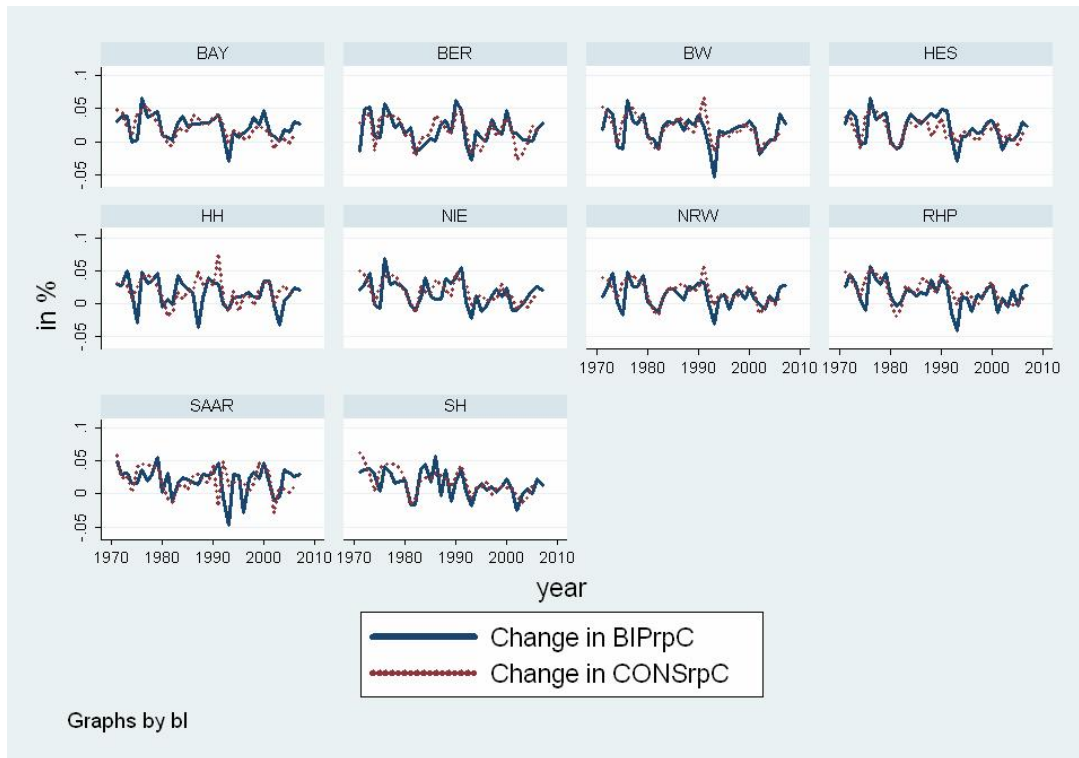
Source: Data from VGRdL(2009).

**Fig. 2: State-level ratio of consumption and GDP per capita in %**



Source: Data from VGRdL(2009).

**Fig. 3: Annual Growth Rate of Consumption and GDP per Capita in %**



*Note:* Data from VGRdl(2009).

We also test for the time series properties of the variables. Here we apply the panel unit root tests proposed by Im et al. (2003) as well as Pesaran's CADF test (see Pesaran, 2005). The advantage of the latter test is that it has been found to be more powerful in the case of spatial dependence (see e.g. Baltagi et al., 2007). Taking the longer West German time series as benchmark the results in table 2 and table 3 show that both income and consumption are integrated series of order  $I(1)$ , and thus turn stationary if transformed into first differences.<sup>2</sup> We apply both panel unit root tests both with and without lag structure as indicated in table 2 and table 3 respectively. The results thus support our hypothesis that income and consumption are non-stationary and potentially co-integrated. We may thus proceed both with our short-run stationary and long-run cointegration estimation strategy.

<sup>2</sup>Results for the panel unit root tests for all 16 states can be obtained from the authors upon request.

**Table 2: IPS Panel unit root test for variables in levels and 1.diff.**

|                  | IPS t-bar test N,T=(10,38(l.); 10,37(d.)) |          |           |          |
|------------------|---|----------|-----------|----------|
|                  | $H_0$ : Series non-stationary             |          |           |          |
| Specification    | Var. 1                                    |          | Var. 2    |          |
|                  | W[t-bar]                                  | (p-val.) | W[t-bar]  | (p-val.) |
| $Y_{i,t}$        | 1.831                                     | (0.96)   | 0.113     | (0.54)   |
| $\Delta Y_{i,t}$ | -15.11***                                 | (0.00)   | -8.820*** | (0.00)   |
| $C_{i,t}$        | -2.089**                                  | (0.02)   | -0.046    | (0.48)   |
| $\Delta C_{i,t}$ | -17.71***                                 | (0.00)   | -8.820*** | (0.00)   |

Note: \*\*\*, \*\*, \* = denote significance levels at the 1%, 5% and 10% level respectively. Variant 1 = lag(0), constant; variant 2 = lag(1), constant. The tests have been performed using the *ipshin* Stata-routines written by Bornhorst & Baum (2007).

**Table 3: Pesaran CADF unit root test for variables in levels and 1.diff.**

|                  | CADF t-bar test N,T=(10,38(l.); 10,37(d.)) |          |           |          |
|------------------|--|----------|-----------|----------|
|                  | $H_0$ : Series non-stationary              |          |           |          |
| Specification    | Var. 1                                     |          | Var. 2    |          |
|                  | Z[t-bar]                                   | (p-val.) | Z[t-bar]  | (p-val.) |
| $Y_{i,t}$        | -0.052                                     | (0.48)   | 0.462     | (0.68)   |
| $\Delta Y_{i,t}$ | -12.55***                                  | (0.00)   | -6.424*** | (0.00)   |
| $C_{i,t}$        | -1.276                                     | (0.10)   | 1.327     | (0.91)   |
| $\Delta C_{i,t}$ | -13.16***                                  | (0.00)   | -7.957*** | (0.00)   |

Note: \*\*\*, \*\*, \* = denote significance levels at the 1%, 5% and 10% level respectively. Variant 1 = lag(0), constant; variant 2 = lag(1), constant. The tests have been performed using the *pescadf* Stata-routines written by Lewandowski (2007).

## 4 Empirical Estimation

### 4.1 Short Run Analysis

Since the short run approach according to eq.(9) to eq.(12)) includes a lagged endogenous regressor in the respective equations, we have to estimate both the AR-type income equation as well as the equation for consumption growth by means of dynamic panel data (DPD) techniques. The econometric literature proposes different IV and non-IV estimators to deal with the likely problem of endogeneity between the lagged endogenous variable and the error term (see e.g. Baltagi, 2008, for an overview). As table 4 shows for the (auxilliary) income equation different estimators including simple Pooled OLS (POLS), corrected Fixed Effects estimation (LSDVC) and IV-based GMM approaches (both Arelland-Bond, 1991, and Blundell-Bond, 1998) yield similar parameter estimates. Given the performance found in various Monte Carlo simulation studies (see e.g. Soto. 2007) we give preference to the Blundell-Bond System GMM (SYS-GMM) in the following.

**Table 4: Estimation results for the (auxilliary) AR income equation**

|             | <b>POLS</b>       | <b>LSDVC</b>      | <b>AB-GMM</b>     | <b>SYS-GMM</b>    | <b>SYS-GMM</b>    |
|-------------|-------------------|-------------------|-------------------|-------------------|-------------------|
| $Y_{i,t-1}$ | 0.98***<br>(0.00) | 0.96***<br>(0.00) | 0.96***<br>(0.00) | 0.94***<br>(0.00) | 0.86***<br>(0.00) |
| Constant    | 0.06***<br>(0.00) | 0.11***<br>(0.00) |                   | 0.19<br>(0.53)    | 0.41<br>(0.51)    |
| Trend       |                   |                   |                   |                   | 0.001<br>(0.67)   |

Note: \*\*\*, \*\*, \* = denote significance levels at the 1%, 5% and 10% level respectively.

The error term from the AR income equation is then used as right hand side regressor in the specification of short-run consumption changes. Here we start from a standard specification typically used in testing for the validity of the PIH (see e.g. Malley & Molana, 2008, for an overview) and subsequently augment the specification by lagged values of  $\Delta C$ . The results are shown in table 5.

**Table 5: Short-run estimates of  $\Delta C_{i,t}$  using SYS-GMM**

|                    |                   |                   |                   |                   |                   |                   |
|--------------------|-------------------|-------------------|-------------------|-------------------|-------------------|-------------------|
| Constant           | 0.01***<br>(0.00) | 0.01***<br>(0.00) | 0.01***<br>(0.00) | 0.01***<br>(0.00) | 0.01***<br>(0.00) | 0.01<br>(0.17)    |
| $\Delta Y_{i,t-1}$ | 0.17***<br>(0.00) | 0.08<br>(0.938)   |                   |                   | 0.01<br>(0.72)    |                   |
| $\Delta Y_{i,t}$   |                   |                   |                   |                   |                   | 0.16<br>(0.68)    |
| $\Delta C_{i,t-1}$ |                   | 0.22**<br>(0.03)  | 0.25***<br>(0.00) | 0.18**<br>(0.06)  | 0.19*<br>(0.05)   | 0.45***<br>(0.00) |
| $\epsilon_{i,t}$   | 0.46***<br>(0.00) | 0.43***<br>(0.00) | 0.43***<br>(0.00) | 0.43***<br>(0.00) | 0.43**<br>(0.04)  | 0.25<br>(0.23)    |
| $\Delta C_{i,t-2}$ |                   |                   |                   | 0.03<br>(0.53)    | 0.04<br>(0.47)    | 0.03<br>(0.57)    |
| $\epsilon_{i,t-1}$ |                   |                   |                   | 0.06<br>(0.42)    | 0.06<br>(0.45)    |                   |
| Obs.               | 360               | 360               | 360               | 350               | 350               | 350               |
| m <sub>1</sub>     | (0.02)            | (0.02)            | (0.02)            | (0.02)            | (0.02)            | (0.04)            |
| m <sub>2</sub>     | (0.07)            | (0.88)            | (0.36)            | (0.92)            | (0.96)            | (0.16)            |
| J-Stat.            | (0.28)            | (0.69)            | (0.34)            | (0.17)            | (0.51)            | (0.25)            |

Note: \*\*\*, \*\*, \* = denote significance levels at the 1%, 5% and 10% level respectively. We include collapsed IVs up to lag(4) in each regression equation. We apply two-step efficient, heteroscedasticity robust GMM estimation.

The results show that the further addition of  $\Delta C_{t-1}$  renders  $\Delta Y_{t-1}$  insignificant. This indicates that  $\Delta Y_{t-1}$  was merely capturing the omitted effect of habit formation rather than excess sensitivity. In line with Malley & Molana (2006) we find that consumption may well be explained by an ARIMA(1,1,0) model. We finally also check whether the ARIMA(1,1,0) remains robust against a more general specification of Campbell & Mankiw

(1990) with a fraction of agents ( $\rho$ ) being liquidity constrained. The results are shown in the last column of table 5. As before we do not find significant "liquidity constraints" in the above PIH framework. This result contrasts Dreger & Kosfeld (2003), DeJuan et al. (2006) who find "excess sensitivity" and "liquidity constraints" for similar samples. To get further insights we now augment the specification to a cointegration perspective.

## 4.2 Combining Short- and Long-run Information in Cointegration Analysis

To further check for the robustness of these findings we move on to combine the long- and short-run perspective. As Dreger & Kosfeld (2003) note, the PIH implies cointegration between consumption and income in the long-run. We first check for the long run comovement of  $C$  and  $Y$  based on Westerlund's (2007) panel cointegration tests, again starting with the long panel for the West German states between 1970 and 2007:

**Table 6: Panel Cointegration tests for income and consumption for West Germany**

| Test       | Value    | P-Val. | Robust P-Val. |
|------------|----------|--------|---------------|
| $G_\tau$   | -3.04*** | (0.00) | (0.00)        |
| $G_\alpha$ | -7.74*** | (0.00) | (0.01)        |
| $P_\tau$   | -8.28*** | (0.00) | (0.00)        |
| $P_\alpha$ | -5.55*** | (0.00) | (0.00)        |

*Note:* \*\*\*, \*\*, \* = denote significance levels at the 1%, 5% and 10% level respectively. Automatic Lag-selection based on the AIC, Robust P-Values are obtained on the bootstrapped distribution with 500 reps. Calculations based on the *xtwest* Stata routine by Persyn & Westerlund (2008).

We then specify a Panel Error Correction Model (Panel ECM) of the following general form:

$$\begin{aligned} \Delta C_{i,t} = & -\phi(C_{i,t-1} - \gamma_{0,i} - \gamma_{1,i}t - \kappa_{1,i}Y_{i,t}) \\ & + \sum_{i=1}^k b_{0,i}\Delta C_{i,t-i} + \sum_{i=0}^k b_{1,i}\Delta Y_{i,t-i} \end{aligned} \quad (15)$$

We estimate the model by dynamic Fixed Effects (DFE) and Pooled Mean Group (see Pesaran et al., 1999). The DFE model assumes homogeneity of short and long-run parameters in the panel, while the PMG estimator allows for short-run heterogeneity. As in the short-run analysis we can test for the significance of liquidity constrained agents as a fraction of all agents ( $\rho$ ). To take a closer look at the different estimators they can be summarized as follows:

- MG estimator calculates short and long run parameters as unweighted means of

individual coefficients such as:

$$\hat{\psi} = \sum_{i=1}^N \frac{1}{N} \hat{\psi}_i \quad \hat{\Delta}_{\hat{\phi}} = \frac{1}{N(N-1)} \sum_{i=1}^N (\hat{\phi}_i - \hat{\phi})^2 \quad (16)$$

$$\hat{\kappa} = \sum_{i=1}^N \frac{1}{N} \hat{\kappa}_i \quad (17)$$

- PMG constrains long-run parameters to be identical, when restriction holds more efficient than MG:

$$\hat{\psi} = \sum_{i=1}^N \frac{1}{N} \hat{\psi}_i \quad (18)$$

$$\hat{\kappa} = \hat{\kappa}_i \quad \forall i \quad (19)$$

- DFE finally constrains short- and long-run parameters to be identical for all cross-sections

$$\hat{\psi} = \hat{\psi}_i \quad \forall i \quad (20)$$

$$\hat{\kappa} = \hat{\kappa}_i \quad \forall i \quad (21)$$

Since the MG estimator is a consistent (although potentially inefficient) estimator we can use it as a benchmark to test for the validity of cross-section restrictions in a series of Hausman tests. Additionally we borrow from the growing literature on 'poolability tests' or 'tests on slope homogeneity' (see e.g. Bun, 2004, Pesaran & Yamagata, 2008, as well as Baltagi, 2008, for recent surveys.) To test for short-run homogeneity we may thus apply a set of F- or Wald tests in the spirit of the Roy-Zellner poolability test of the form:

$$H_0 : b_{j,1} = b_{j,2} = \dots = b_{j,k} \text{ with: } i = 1, \dots, k \quad (22)$$

For the latter testin approach we use both standard asymptotic as well as bootstrapped versions of the test as recently proposed by Bun (2007). Table 7 displays the estimation results of the Panel ECM for the West German states using the different estimators discussed above. We also present a set of postestimation tests to guide statistical inference.

The results are as follows: The coefficient measuring the speed of adjustment from short- to long-run is significant in all cases and underlines a cointegration relationship of income and consumption. As in the short-run estimation setup from above lagged income turns out to be insignificant in PMG and DFE specification. The more rigorous form using current rather than lagged income changes as proposed by Campbell & Mankiw (1990)

**Table 7: Estimation results of the Panel ECM for  $\Delta C_{i,t}$**

|                               | WG 70-07           | WG 70-07           | WG 70-07           | WG 70-07           | WG 70-07           | WG 70-07           |
|-------------------------------|--------------------|--------------------|--------------------|--------------------|--------------------|--------------------|
|                               | MG                 | PMG                | DFE                | MG                 | PMG                | DFE                |
| <b>Long run coefficient</b>   |                    |                    |                    |                    |                    |                    |
| $Y_{i,t}$                     | 0.93***<br>(0.00)  | 0.78***<br>(0.00)  | 0.83***<br>(0.00)  | 0.90***<br>(0.00)  | 0.74***<br>(0.00)  | 0.74***<br>(0.00)  |
| <b>Short run coefficients</b> |                    |                    |                    |                    |                    |                    |
| $\phi$                        | -0.30***<br>(0.00) | -0.16***<br>(0.00) | -0.12***<br>(0.00) | -0.16***<br>(0.00) | -0.09***<br>(0.00) | -0.08***<br>(0.00) |
| $\Delta C_{i,t-1}$            | 0.03***<br>(0.00)  | 0.02***<br>(0.00)  | 0.02***<br>(0.00)  | 0.01***<br>(0.00)  | 0.01***<br>(0.00)  | 0.01***<br>(0.00)  |
| $\Delta Y_{i,t}$              |                    |                    |                    | 0.28***<br>(0.00)  | 0.34***<br>(0.00)  | 0.35***<br>(0.00)  |
| $\Delta Y_{i,t-1}$            | -0.09**<br>(0.04)  | -0.03<br>(0.56)    | 0.05<br>(0.15)     |                    |                    |                    |
| $ m $ -stat.                  |                    | 1.73<br>(0.19)     | 0.25<br>(0.61)     |                    | 1.71<br>(0.19)     | 0.01<br>(0.98)     |
| Wald 1                        |                    |                    |                    |                    |                    | 4.45<br>(0.87)     |
| Wald 2                        |                    |                    |                    |                    |                    | 8.37<br>(0.49)     |

*Note:* \*\*\*, \*\*, \* = denote significance levels at the 1%, 5% and 10% level respectively. Wald 1 = Wald test on equality of error correction term, Wald 2 = Wald test on equality of  $\Delta Y_{i,t}$ .

shows however a significant fraction of liquidity constrained households. The obtained results match earlier findings ( $\rho = 0.29$  for West Germany in Wolters, 1992;  $\rho = 0.26 - 0.29$  for US in Fuhrer, 2000). In all specifications Hausman m-statistic favours long-run restriction in PMG and DFE. Short-run tests Wald1 ( $\psi$ ) and Wald2 ( $\Delta y$ ) favour slope homogeneity in terms of DFE.

We additionally also estimate the Panel ECM for the sample of all 16 German states from 1991. One motivation for doing so is to account for a potential structural break in the variables due to German Re-unification and the question is: How does it affect regional consumption paths? Another research question to answer given the huge macro regional differences between East and West Germany is: Are the less wealthy regions in East Germany more liquidity constrained than their Western counterparts? Earlier results in Dreger & Kosfeld (2003) for West Germany indeed find a positive correlation between the share of liquidity constrained household and regional unemployment rates.

We again use Panel cointegration tests to check for the cointegration of income and consumption. The results in table 8 are less evident. However we are not sure about the power of the test for shorter time dimensions. Table 9 then presents the results of the Panel ECM estimation.

**Table 8: Panel Cointegration tests for income and consumption for West Germany**

| Test       | Value   | P-Val. |
|------------|---------|--------|
| $G_\tau$   | -1.33*  | (0.08) |
| $G_\alpha$ | -3.08   | (0.73) |
| $P_\tau$   | -3.88** | (0.05) |
| $P_\alpha$ | -1.99*  | (0.09) |

Note: \*\*\*, \*\*, \* = denote significance levels at the 1%, 5% and 10% level respectively. Lag selection based on the AIC. Calculations based on the *xtwest* Stata routine by Persyn & Westerlund (2008).

**Table 8: Estimation results of the Panel ECM for  $\Delta C_{i,t}$**

|                               | G 91-07            | G 91-07            | G 91-07            | G 91-07            | G 91-07            | G 91-07            |
|-------------------------------|--------------------|--------------------|--------------------|--------------------|--------------------|--------------------|
|                               | MG                 | PMG                | DFE                | MG                 | PMG                | DFE                |
| <b>Long run coefficient</b>   |                    |                    |                    |                    |                    |                    |
| $Y_{i,t}$                     | 0.59***<br>(0.00)  | 0.44***<br>(0.00)  | 0.53***<br>(0.00)  | 0.22<br>(0.47)     | 0.37***<br>(0.00)  | 0.53***<br>(0.00)  |
| <b>Short run coefficients</b> |                    |                    |                    |                    |                    |                    |
| $\phi$                        | -0.42***<br>(0.00) | -0.31***<br>(0.00) | -0.24***<br>(0.00) | -0.31***<br>(0.00) | -0.25***<br>(0.00) | -0.22***<br>(0.00) |
| $\Delta C_{i,t-1}$            | 0.01**<br>(0.05)   | 0.01<br>(0.25)     | 0.01***<br>(0.00)  | 0.01***<br>(0.00)  | 0.01***<br>(0.00)  | 0.02***<br>(0.00)  |
| $\Delta Y_{i,t}$              |                    |                    |                    | 0.10*<br>(0.08)    | 0.15***<br>(0.00)  | 0.12***<br>(0.00)  |
| $\Delta Y_{i,t-1}$            | -0.01<br>(0.84)    | 0.08*<br>(0.08)    | 0.04<br>(0.29)     |                    |                    |                    |
| $ m  - stat.$                 |                    | 3.32<br>(0.07)     | 0.01<br>(0.99)     |                    | 0.25<br>(0.62)     | 1.12<br>(0.29)     |
| Wald 1                        |                    |                    | 18.27<br>(0.24)    |                    |                    | 18.75<br>(0.23)    |
| Wald 2                        |                    |                    | 22.05<br>(0.11)    |                    |                    | 37.65<br>(0.00)    |

Note: \*\*\*, \*\*, \* = denote significance levels at the 1%, 5% and 10% level respectively. Wald 1 = Wald test on equality of error correction term, Wald 2 = Wald test on equality of  $\Delta Y_{i,t}$ .



The results of the Panel ECM for unified Germany between 1991 – 2007 still show significant coefficients for the error correction term. However, the long run relation in the cointegration equation between the variables is less tight. Interestingly, for the short-run parameter coefficients the fraction of significantly liquidity constrained household is smaller compared to the West German sample. Turning to the post estimation tests, the Hausman m-statistic still shows homogeneous long-run restriction. However, the Wald test for variable  $\Delta y$  strongly rejects slope homogeneity with respect to the short-run reaction of regional consumption patterns due to current income changes. We aim to explore this asymmetry more in depth in the following.

## 5 Are there different long and short run regimes?

In this section we finally relax the strong all-or-nothing decision between the short-run homogeneous DFE and heterogeneous PMG. We therefore adapt a testing routine from the literature about (growth) convergence clubs and apply clustering algorithms. Our routine involves the following steps:

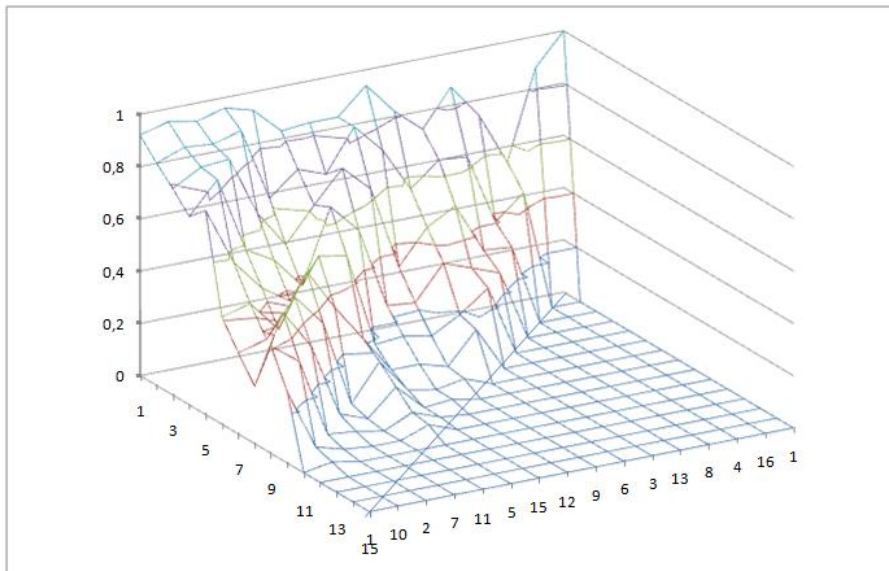
1. Estimate the short-run parameters for each cross-section separately and order them according to coefficient size.
2. Perform for the first  $k$  cross-sections ( $k=2$ ) a Wald test for coefficient equality and add further units until the test for parameter restriction is rejected.
3. Form a second group from all cross-sections outside the first club and test for parameter restrictions.
4. If rejected repeat step 1-3 on remaining cross-sections to search for subgroups that form further clubs.

Additionally we also run pairwise Wald tests to search for slope homogeneity. Figure 3 and 4 plot the pairwise Wald tests for short-run coefficient equality of  $\delta y$  as well as a surface plot of the obtained p-values. The results show two clubs starting from 1991 which exactly match the West and East German macro region respectively. However, different to prior expectations Eastern states are less open to liquidity constraints. One likely explanation for this phenomena is the time smoothing of regional consumption paths in East German states due to massive income transfers from the West.

Figure 3: Pairwise Wald tests for short-run coefficient equality of  $\delta y$

| BL | 1    | 10   | 2    | 7    | 11   | 5    | 15   | 12   | 9    | 6    | 3    | 13   | 8    | 4    | 16   | 14 |
|----|------|------|------|------|------|------|------|------|------|------|------|------|------|------|------|----|
| 1  |      |      |      |      |      |      |      |      |      |      |      |      |      |      |      |    |
| 10 | 0,92 |      |      |      |      |      |      |      |      |      |      |      |      |      |      |    |
| 2  | 0,85 | 0,95 |      |      |      |      |      |      |      |      |      |      |      |      |      |    |
| 7  | 0,79 | 0,89 | 0,93 |      |      |      |      |      |      |      |      |      |      |      |      |    |
| 11 | 0,72 | 0,85 | 0,88 | 0,96 |      |      |      |      |      |      |      |      |      |      |      |    |
| 5  | 0,78 | 0,84 | 0,86 | 0,91 | 0,93 |      |      |      |      |      |      |      |      |      |      |    |
| 15 | 0,41 | 0,56 | 0,55 | 0,63 | 0,63 | 0,83 |      |      |      |      |      |      |      |      |      |    |
| 12 | 0,31 | 0,44 | 0,44 | 0,49 | 0,49 | 0,72 | 0,84 |      |      |      |      |      |      |      |      |    |
| 9  | 0,22 | 0,34 | 0,32 | 0,37 | 0,37 | 0,62 | 0,68 | 0,84 |      |      |      |      |      |      |      |    |
| 6  | 0,41 | 0,48 | 0,48 | 0,52 | 0,52 | 0,65 | 0,74 | 0,84 | 0,94 |      |      |      |      |      |      |    |
| 3  | 0,27 | 0,33 | 0,33 | 0,36 | 0,36 | 0,5  | 0,54 | 0,64 | 0,73 | 0,82 |      |      |      |      |      |    |
| 13 | 0    | 0,04 | 0,03 | 0,05 | 0,03 | 0,22 | 0,14 | 0,23 | 0,32 | 0,56 | 0,73 |      |      |      |      |    |
| 8  | 0    | 0,02 | 0,01 | 0,02 | 0,01 | 0,17 | 0,08 | 0,15 | 0,22 | 0,45 | 0,65 | 0,87 |      |      |      |    |
| 4  | 0    | 0    | 0    | 0    | 0    | 0,11 | 0,04 | 0,08 | 0,13 | 0,35 | 0,51 | 0,64 | 0,74 |      |      |    |
| 16 | 0    | 0    | 0    | 0    | 0    | 0,05 | 0,01 | 0,02 | 0,05 | 0,19 | 0,31 | 0,32 | 0,37 | 0,54 |      |    |
| 14 | 0    | 0    | 0    | 0    | 0    | 0,04 | 0,01 | 0,02 | 0,04 | 0,17 | 0,27 | 0,27 | 0,31 | 0,46 | 0,88 |    |

Figure 4: Surface plot of p-values for short-run coefficient equality of  $\delta y$



## 6 Robustness Check: Spatial Filtering

As a robustness check for the above results we finally apply Getis (1995) spatial filtering to remove any potential spatial autocorrelation from the data. The idea to spatially filter the variables is similar to idea of filtering out seasonality in time series data. We are thus able to decompose the original variable  $Y$  into a structural component  $Y^*$  and a purely spatial component  $S$  according to  $Y^* = (Y - S)$ . The Getis approach uses the local  $G_i(d)$  statistic by Getis & Ord (1992). The approach needs a binary weighting matrix, which we define in terms of a common border (matrix cell =1) and zero otherwise. In a first step we check for the presence of spatial autocorrelation in the residuals of the aspatial Panel ECM results. We find that spatial autocorrelation is highly present and may thus introduce a potential bias to the estimated coefficients. We then apply the Getis approach to the original variables used for estimation. Table 9 reports the results.

**Table 9: Moran's  $I$  values for income and consumption based on the Getis approach**

|             | Common Border |                 |                |                  |
|-------------|---------------|-----------------|----------------|------------------|
| year        | <i>biprpc</i> | <i>biprpc</i> * | <i>consrpc</i> | <i>consrpc</i> * |
| <b>2007</b> | 0.21**        | -0.12           | 0.35***        | -0.04            |
| <b>2006</b> | 0.22**        | -0.12           | 0.34***        | -0.05            |
| <b>2005</b> | 0.21**        | -0.11           | 0.32***        | -0.05            |
| (...)       |               |                 |                |                  |
| <b>2000</b> | 0.26**        | -0.07           | 0.37***        | 0.03             |
| (...)       |               |                 |                |                  |
| <b>1995</b> | 0.28***       | -0.05           | 0.43***        | 0.04             |
| (...)       |               |                 |                |                  |
| <b>1991</b> | 0.42***       | -0.06           | 0.44***        | 0.09             |

Note: \*\*\*, \*\*, \* = denote significance levels at the 1%, 5% and 10% level respectively.

As the table shows, the Getis approach is very effective in decomposing the variable into a structural component, which is free of any spatial autocorrelation. We then use the filtered variables to reestimate the Panel ECM. The results are shown in table 10 together with post estimation tests for remaining spatial autocorrelation in table 11.

The general result from the estimation output in table 10 is that the coefficient estimates are broadly the same for the spatially filtered Panel ECM and even more in line with our apriori expectations. The only notable difference is the result for the Wald2 test, which now does not reject the restriction of slope homogeneity with respect to  $\Delta Y$ . If we look at the estimated coefficients for each region we now observe a big difference for Brandenburg and its strong spatial interrelations with Berlin. Here it would be desirable for future research to additionally apply spatial panel data techniques which allow spatial

**Table 10: Estimation results of the spatially filtered Panel ECM for  $\Delta C_{i,t}$**

|                               | G 91-07            | G 91-07            | G 91-07            | G 91-07            | G 91-07            | G 91-07            |
|-------------------------------|--------------------|--------------------|--------------------|--------------------|--------------------|--------------------|
|                               | FMG                | FPMG               | FDFE               | FMG                | FPMG               | FDFE               |
| <b>Long run coefficient</b>   |                    |                    |                    |                    |                    |                    |
| $Y_{i,t}$                     | 0.60***<br>(0.00)  | 0.71***<br>(0.00)  | 0.58***<br>(0.00)  | 0.60***<br>(0.00)  | 0.47***<br>(0.00)  | 0.59***<br>(0.00)  |
| <b>Short run coefficients</b> |                    |                    |                    |                    |                    |                    |
| $\phi$                        | -0.46***<br>(0.00) | -0.31***<br>(0.00) | -0.27***<br>(0.00) | -0.34***<br>(0.00) | -0.25***<br>(0.00) | -0.23***<br>(0.00) |
| $\Delta C_{i,t-1}$            | 0.01**<br>(0.05)   | 0.02***<br>(0.00)  | 0.02***<br>(0.00)  | 0.02***<br>(0.00)  | 0.01***<br>(0.00)  | 0.01***<br>(0.00)  |
| $\Delta Y_{i,t}$              |                    |                    |                    | 0.08<br>(0.11)     | 0.14***<br>(0.00)  | 0.11**<br>(0.00)   |
| $\Delta Y_{i,t-1}$            | -0.10<br>(0.20)    | -0.07<br>(0.26)    | -0.01<br>(0.80)    |                    |                    |                    |
| $ m  - stat.$                 |                    | 2.75<br>(0.09)     | 0.01<br>(0.99)     |                    | 10.38<br>(0.01)    | 0.06<br>(0.80)     |
| Wald 1                        |                    |                    | 20.58<br>(0.15)    |                    |                    | 21.35<br>(0.13)    |
| Wald 2                        |                    |                    | 23.27<br>(0.08)    |                    |                    | 20.25<br>(0.16)    |

Note: \*\*\*, \*\*, \* = denote significance levels at the 1%, 5% and 10% level respectively. Wald 1 = Wald test on equality of error correction term, Wald 2 = Wald test on equality of  $\Delta Y_{i,t}$ .

**Table 11: Moran's  $I$  values for Panel ECM residuals**

| year        | Border               |                    |
|-------------|----------------------|--------------------|
|             | $resid_{unfiltered}$ | $resid_{filtered}$ |
| <b>2007</b> | 0.29***              | -0.16              |
| <b>2006</b> | 0.28***              | -0.15              |
| <b>2005</b> | 0.30***              | -0.16              |
| (...)       |                      |                    |
| <b>2000</b> | 0.30***              | -0.15              |
| (...)       |                      |                    |
| <b>1995</b> | 0.40***              | -0.02              |
| (...)       |                      |                    |

Note: \*\*\*, \*\*, \* = denote significance levels at the 1%, 5% and 10% level respectively.

variables to be included in the estimation approach and thus be able to interpret their results (in the spirit of spatial lag or spatial Durbin models). Finally, the Moran's  $I$  based post estimation test reported in table 11 shows that from a statistical perspective the latter filtered Panel ECM does not show any uncaptured spatial dependence in the model's residuals.

## 7 Conclusion

We have tested the PIH approach with German regional data for the period 1970 (or 1991) up to 2007. In the short-run PIH model with habit persistence we do not obtain evidence for excess sensitivity of consumption changes to current income shocks and thus a significant fraction of liquidity constrained households. Combining the long- and short-run perspective income and consumption are found to be cointegrated (supporting PIH). In the short run adjustment of our Panel ECM we find a significant but lower fraction of liquidity constrained agents (around 30-35 % compared to 45 % in Dreger & Kosfeld, 2003). Our findings match earlier results reported in Wolters (1992) who estimates  $\rho = 0.29$  for West Germany. Thus, both short- and long-run estimation results indicate the limited effectiveness of fiscal policies to strengthen the demand side. Especially for the total German sample since re-unification the degree of excess sensitivity to income shocks is small. The empirical findings for East German states having even lower fractions of liquidity constrained households compared to the West may be best explained by massive East-West transfers smoothing regional consumption paths. The explicit account of spatial dependence seems to be important. To give an interpretation of the role of spatial spillovers future research should more carefully take into account the concept of global time-space cointegration (see e.g. Beenstock & Felsenstein, 2009).

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