

# Consumption and Income

## Panel econometric Evidence for West Germany

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**Abstract.** In this paper the permanent income hypothesis (PIH) of consumption is tested by means of panel econometric techniques, which are applied to West German regional data. Panel unit-root tests which are robust to cross section correlation indicate the stationarity of the savings rate, revealing weakly support for the PIH. However, consumers may be subject to liquidity constraints. The relevance of non-optimizing consumers is examined within the error-correction  $\lambda$ -model. Approximately a fraction of 45 percent of disposable income is earned by households that do not behave according to the PIH, and the share is positively related to regional unemployment rates.

**Key words:** Permanent Income Hypothesis, Liquidity Constraints, Panel Unit Root Tests

**JEL classification:** C23, E21

### 1 Introduction

Consumption of private households is by far the biggest aggregate in GDP. In Germany consumption expenditures amount to 55 percent in GDP on average over the last two decades. Other industrialized countries experience quite similar shares. Explaining consumption spending is therefore a prerequisite for accurate forecasts of GDP.

Despite its relevance only a few actual studies analyse aggregate consumption behaviour in Germany. Wolters (1992) e.g. discusses the empirical performance of several specifications, especially error correction models build on consumption and disposable income series. Reimers (1997) and Hassler (2001) extend the cointegration analysis by

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introducing a wealth variable, representing accumulated savings. The former author reports tests for seasonal cointegration, while the latter discusses multicointegration topics between consumption, income and wealth.

This paper chooses the life-cycle permanent income hypothesis (PIH) for consumption as the point of departure. The PIH implies cointegration between consumption and disposable income or a stationary saving series. As a novelty, this hypothesis is tested within a paneleconometric framework by the means of recently developed panel unit root tests. These tests utilize a broader long run information set and are therefore better suited to detect a false null hypothesis than the conventional unit root procedures. The panel is based on data from West German regions. To our knowledge no other study has analysed the regional data set so far.

However, note that cointegration analysis can only provide weak tests of the validity of the PIH. Stationarity of the savings series is also consistent with several specifications, including an ordinary Keynesian type consumption function. Here expenditures depend not on permanent, but on actual income, which can be justified on the basis of myopic or credit rationing behaviour, see Campbell and Mankiw (1991). Thus, the relevance of market imperfections has to be investigated and this is also done within the panel error correction framework. Essentially, regional consumption functions are estimated using the SUR technique, and the restriction of an equal feedback parameter across the panel members is tested by the means of a standard Wald test. The coefficient of actual income growth may be interpreted as the income share earned by liquidity constrained consumers. This view is supported by a correlation analysis between the individual shares and measures of regional economic performance such as the unemployment rate. As a result of the analysis, regional weights of actual and permanent income in explaining consumption expenditures per capita are estimated.

The rest of the paper is organized as follows. The next section briefly reviews the relevant theory of the PIH including the modifications in the presence of liquidity constraints. Second, panel unit root tests are discussed. After describing the regional data set, the empirical analysis is performed. Section 6 concludes.

## 2 Permanent Income and $\lambda$ -model

The PIH in a stochastic environment results from optimizing behaviour of infinitely long living agents, c.f. Hall (1978). Representative consumers have a felicity (instantaneous utility) function  $u(C_t)$ , relating utility only to the level of current consumption  $C_t$  per capita. In fact, the expected life cycle utility

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

is maximized subject to the wealth accumulation equation

$$(2) \quad W_{t+1} = (1+r)W_t + Y_t - C_t$$

where  $W_t$  is financial wealth at the beginning of the period  $t$  and  $r$  the real interest rate.  $rW_t$  is capital income, and  $Y_t$  labor income received in the current period. The time preference rate  $\delta$  and the real interest rate are assumed to be constant.  $E_t$  indicates expectations conditional on the information set available to the representative consumer at time  $t$ . This amounts to consumers having rational expectations. Recursive substitution in equation (2) and taking expectations leads to the infinite horizon budget constraint

$$(3) \quad \sum_{i=0}^{\infty} \left( \frac{1}{1+r} \right)^i E_t C_{t+i} = (1+r)W_t + \sum_{i=0}^{\infty} \left( \frac{1}{1+r} \right)^i E_t Y_{t+i}$$

where no Ponzi games have been ruled out.<sup>1</sup> According to equation (3), the present value of lifetime consumption is equal to the sum of financial and human wealth, and the latter is defined as the present value of expected future labour incomes. Differentiating the Lagrangean of the constrained optimization problem stated in equations (1) and (3) gives the familiar stochastic Euler equation

$$(4) \quad E_t u'(C_{t+i}) = \frac{1+\delta}{1+r} u'(C_t)$$

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<sup>1</sup> The no Ponzi game assumption sets the present value of financial wealth (debt) to 0 as  $i$  goes to infinity. Otherwise agents can borrow indefinitely to finance an increase in consumption.

which implies that on the optimal path marginal utilities of consumption follow a first order Markov process. Moreover, marginal utilities and thus the corresponding levels of consumption per capita are constant, when the time preference rate is equal to the real interest rate. In the following, this restriction is assumed to be met.

Further analysis requires the specification of the instantaneous utility function. Utility may be quadratic in consumption,

$$u(C_t) = -0.5(C_t - C^*)^2$$

where  $C^*$  is some constant bliss level of consumption. Inserting into equation (4) and shifting time back one period yields:

$$(5) \quad C_t = C_{t-1} + v_t$$

where  $v$  is a white noise forecasting error, stemming from the rational expectations assumption. According to equation (5) consumption follows a random walk (without drift).<sup>2</sup> Moreover  $E_t C_{t+i} = C_t$  holds for all  $i$ , implying that the intertemporal budget constraint may be rewritten as

$$(6) \quad C_t = rW_t + \frac{r}{1+r} \sum_{i=0}^{\infty} \left( \frac{1}{1+r} \right)^i E_t Y_{t+i}$$

which is the consumption function under the PIH, c.f. Flavin (1981). Ex ante consumption is equal to permanent income, and the latter is defined as the right hand side of equation (6) which is the sum of capital income and expected future labor income. The terms are obtained as present values at the beginning of the period.

Disposable income is equal to the sum of both income components:  $YD_t = rW_t + Y_t$ . Utilizing the definition of saving,  $S_t = YD_t - C_t$  and equation (6) the PIH may be rewritten as

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<sup>2</sup> See Hall (1978). The random walk hypothesis can be reformulated in the logs of consumption, when a CES utility function is assumed, see Campbell and Mankiw (1991).

$$(7) \quad S_t = -\sum_{i=0}^{\infty} \left( \frac{1}{1+r} \right)^i E_t \Delta Y_{t+i},$$

where  $\Delta$  is the first difference operator. This formulation essentially captures the rainy day feature of the PIH. Saving occurs if a decline in labour income is expected. Moreover, saving serves as an optimal predictor for anticipated changes in labor income, c.f. Campbell (1987). Finally, if labour income has a unit root, saving is stationary since it is the sum of stationary terms with exponentially declining weights, provided the real interest rate is positive. As a result, consumption and disposable income are cointegrated, and the vector corresponding to the long run equilibrium relation is (1,-1).

According to the PIH and the assumptions needed for equation (5), changes in consumption correspond to revisions in expected labor income changes. This can be seen by inserting equation (2) in equation (6). Surprises in the income process are purely white noise implying that changes in consumption are generally unforecastable.

However, empirical evidence suggests that predictable changes of income do have an impact on actual changes in consumption, see Muellbauer (1994) for a survey. According to Campbell and Mankiw (1991), this so-called excessive sensitivity of consumption may be a reflection of liquidity constraints. Credit-rationed agents do not save and spend their whole income, which solely stems from labor. For example, young people with low income, who expect to earn more in the future are faced by such restrictions, when desired credits are denied. In opposite, optimizing households consume without constraints, and their expenditures fluctuate randomly. The shares of the two distinct groups in the income distribution are  $\lambda$  and  $1-\lambda$ , respectively. Therefore, the changes in consumption may be reproduced by the so called  $\lambda$ -model

$$(8) \quad \Delta C_t = \lambda \Delta Y D_t + (1-\lambda)v_t + \gamma(C_{t-1} - Y D_{t-1}),$$

see Campbell and Mankiw (1991). Short-run movements of consumption are explained via a linear combination of the ordinary Keynesian function and the PIH. A test for the

importance of liquidity constraints relies on the null hypothesis  $\lambda=0$ .<sup>3</sup> If the null is rejected, the estimated parameter may be interpreted as the fraction of credit-restricted consumers in the economy; the complement fraction  $1-\lambda$  is earned by optimizing consumers.<sup>4</sup> Unpredictable fluctuations in permanent income are accounted for by the white noise error term  $v$ . In any case, the lagged saving series must be included as an additional regressor if consumption and disposable income are cointegrated ( $\gamma < 0$ ).

### 3 Panel Unit Root Tests

In the following, testing for cointegration between consumption and disposable income is implemented as a test for a unit root in the saving series. The analysis is carried out within a paneleconometric framework. Panel unit root tests have greater power than the conventional unit root tests. Since the time series dimension is enhanced by the cross section, the tests rely on a broader long run information set. Thus they are better suited to detect a true alternative hypothesis even in the presence of nearly unit root alternatives, see Levin and Lin (1993) and Maddala and Wu (1999). However, the increase in power is not costless. In particular, contemporaneous correlation can arise and the unit root properties may differ across the panel members. These two pitfalls are discussed in turn.

First, the contemporaneous correlation may be attributed to common shocks which affect jointly all the panel individuals. As a consequence, independent long run information is lost. If the correlation structure is not recognized, test statistics will suffer from substantial size distortions, see O'Connell (1998). In fact, the true size of the unit root tests can be far above the nominal level even if the correlation is moderate. Thus, the results of the tests are highly questionable when the dependencies are not modelled appropriately.

Second, panel unit root tests do not reveal any individual specific information in respect to the unit root feature. If the joint null hypothesis of a unit root is rejected, the series

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<sup>3</sup> This also may reflect some kind of myopic behaviour. Furthermore, if future income is more heavily discounted due to income uncertainty, current income is necessarily of higher relevance in determining actual consumption.

may be stationary for all cross sections or only for a subgroup. The null might be rejected even if there is only one stationary individual in an otherwise unit root environment, see the Monte Carlo evidence presented by Taylor and Sarno (1998). Therefore test results must be interpreted with some caution. However, this vagueness seems to be less important in the present study. The panel consists of West German regions, and consumer behaviour should not differ substantial across the individuals.<sup>5</sup>

Several panel unit root tests are available, see Banerjee (1999) and Breuer, McNown and Wallace (1999) for recent surveys. Most popular are the tests of Levin and Lin (1993) and Im, Pesaran and Shin (1997), hereafter IPS. Both rely on ADF principles, and the joint null is always a unit root for all panel members. The tests differ in their alternatives, respectively. In particular, if the null is rejected, Levin and Lin (1993) conclude the stationarity of the series for all regions, while in the IPS (1997) setting, at least one individual is stationary. A different degree of heterogeneity allowed for the cross sections is responsible for this outcome. In the Levin and Lin (1993) procedure, heterogeneity is restricted to the deterministic components and short run dynamics, but the unit root properties are restricted to be the same. Instead, IPS (1997) relax the assumption of an equal order of integration across the panel members.

In the presence of contemporaneous correlation the distributions of the test statistics are no longer valid. IPS (1997) consider the special case where cross-correlation between the panel members is caused by time specific effects. The dependencies can be removed, when testing relies on mean-adjusted data. Time effects are identical for all individuals and so the mean is computed over the cross sections. Unfortunately, this approach is useful only for a specific correlation structure and hence, not applicable in general.

In order to control for any type of contemporaneous correlation the multivariate extension of the ADF test recently proposed by Taylor and Sarno (1998) is preferred. The multivariate ADF test (hereafter MADF test) investigates the joint null that the series

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<sup>4</sup> If  $\lambda \neq 0$ , saving is a multiple  $(1-\lambda)$  of (7), see Jin (1995). Thus, saving will be smoother than under pure PIH conditions.

<sup>5</sup> Breuer, McNown and Wallace (1999) suggest complementary ADF tests in a SUR setting in order to test for individual unit roots. Taylor and Sarno (1998) present a Johansen style likelihood ratio test in order to determine whether all individual series are stationary. Due to the degree-of-freedom problem, the latter test is only applicable in panels where the cross section dimension is small.

has a unit root for all regions. Robust testing is done on the basis of a SUR analysis, which is essentially GLS using an estimate of the contemporaneous covariance matrix of the errors obtained from individual OLS estimation in the previous step. The  $i$ th equation of the system is given by

$$(9) \quad \Delta y_{it} = \alpha_i + \delta_i y_{it-1} + \sum_{j=1}^{p_i} \phi_{ij} \Delta y_{it-j} + v_{it}$$

where  $v_{it}$  denotes the individual White Noise error and  $i$  and  $t$  indicate the cross section and the time series dimension of the panel, respectively ( $i=1, \dots, N$ ;  $t=1, \dots, T$ ). The errors are allowed to be contemporary correlated. Member specific effects, short run dynamics and unit root properties are permitted to vary across the individuals, as in the procedure suggested by IPS (1997). The joint null of a unit root  $\delta_1 = \dots = \delta_N = 0$  is judged by the means of a Wald test. If the null is rejected, at least one individual is stationary. As far as similarity of West German states may be justified, stationarity can be concluded for all panel members.

Note that the distribution of the MADF is nonstandard mainly because of the presence of nuisance parameters which represent the cross section correlations under the null hypothesis.<sup>6</sup> Thus, the empirical finite sample distribution has to be derived through simulation and details for the present study are given in the appendix. The critical values are specific to the estimated covariance matrix, the sample size and the number of panel members.<sup>7</sup> Due to this requirement panel testing might be less attractive, but nevertheless it is a reliable way to proceed.

After testing for cointegration between consumption and disposable income, an error correction model in the style of equation (8) is implemented to address the liquidity constraint feature. This model is also estimated along the lines of the SUR framework. At this step, the usual asymptotics apply, since all variables are stationary.

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A normal limit theory can be obtained if the errors are cross-sectionally independent, see Im, Pesaran and Shin (1997).

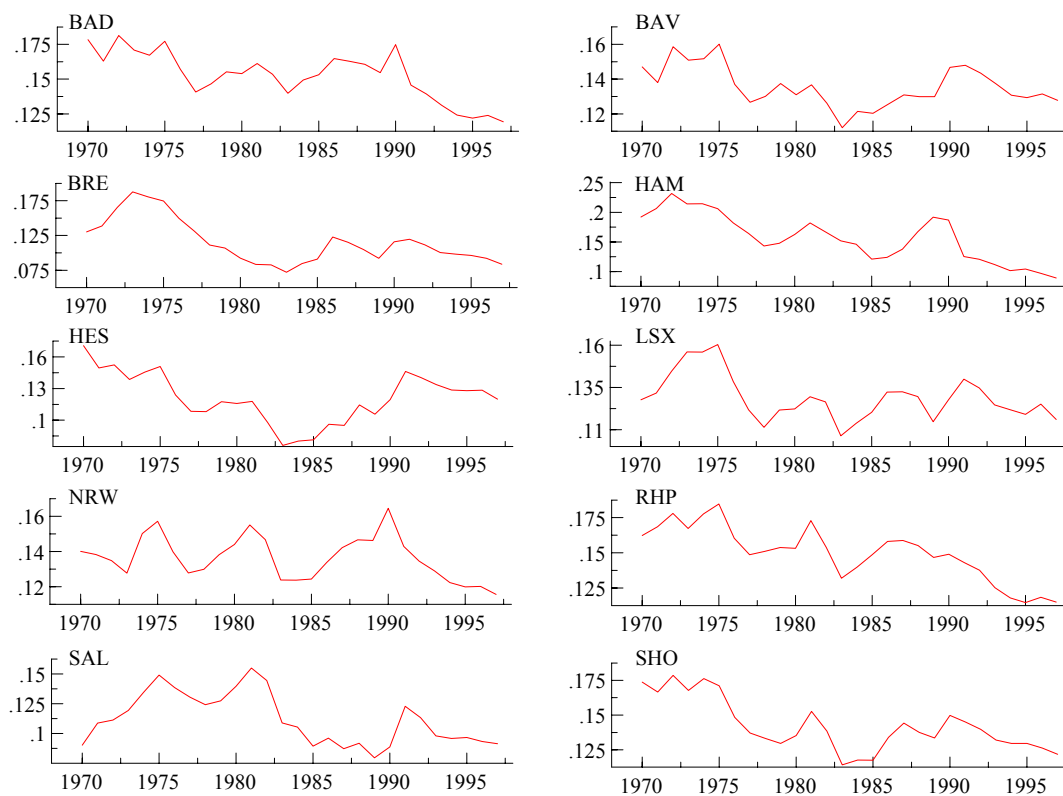
<sup>7</sup> When the tests of Levin and Lin (1993) and Im, Pesaran and Shin (1997) are applied, a non-parametric bootstrap may also be appropriate in order to deal with the problem of contemporaneous correlation, see Berkowitz and Kilian (2000).



## 4 Data

The analysis is based on real per capita data for private consumption expenditure and disposable income. The price index of consumption (1991=100) is used to deflate the consumption and income series, and per capita variables are constructed by dividing through the population figures. Annual data is employed and the sample period ranges from 1970 through 1997. All series are obtained for 10 West German regions, which constitute the panel: Baden-Württemberg (BAD), Bavaria (BAV), Bremen (BRE), Hamburg (HAM), Hesse (HES), Lower Saxony (LSX), North Rhine Westphalia (NRW), Rhineland Palatinate (RHP), Saarland (SAL) and Schleswig Holstein (SHO). Berlin is excluded from the sample because data are available only up to 1994. In total 280 observations are included.

**Figure 1:** Per capita savings Rates of West German states



The series are taken from the *Arbeitskreis Volkswirtschaftliche Gesamtrechnungen der Länder* that conducts the regional accounting for Germany. The variables are measured in natural logarithms and a cointegration relation between the logs of consumption and income requires the stationarity of the consumption income ratio or the savings rate, respectively.<sup>8</sup> Per capita saving series are plotted in Figure 1.

Despite some dissimilarities, per capita saving in percent is roughly comparable across the regions. Relatively high saving rates can be observed in particular at the beginning of the sample period. Since the first oil crisis in 1973/74, the rates have declined for some panel members, namely Baden-Württemberg, Rhineland Palatinate, Schleswig Holstein and the city states Bremen and Hamburg, whereas for the rest of the regions, the savings rate tends to be mean reversible. Starting with German unification in 1990, a fall in the savings rate seems to be more significant. According to the PIH, this may be attributed to the low growth rates of the economy at the end of the sample period. For example between 1995 and 1997 the growth rate of German GDP was only 1.3%, which is significantly lower than the average growth rate in the whole sample period. Currently the saving rate partially recovers from the historically low levels in the second half of the 1990s. In the first quarter of 2001, it has risen up to 11% for the whole economy, compared to 9% in the corresponding quarter of 2000.

## 5 Results

As a preliminary for setting up the  $\lambda$ -model, the MADF test for a unit root in the savings rate is conducted. In a first step the standard ADF tests are in order to motivate carried out the panel data approach. Here the lag length  $k$  in the individual ADF-regressions is determined by the procedure suggested by Campbell and Perron (1991). Specifically, an upper bound of  $k=3$  is set for all regions. Then  $k$  is reduced sequentially by 1 until the last lag becomes significant. If no lags are significant, the straightforward choice is  $k=0$ . To assess the significance of the lags, the 10% value of the normal distribution is used. This procedure suggests  $k=0$  or  $k=1$  for exactly one half of the panel

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<sup>8</sup> The behaviour of saving and the change in consumption under the PIH can be restated in terms of the log consumption-income ratio and the growth rate of consumption, see Campbell and Mankiw (1991).

members. In addition the ADF regression includes a constant, but no time trend. Results are reported in Table 1.

Table 1: ADF tests for the savings rate

	k	ADF		K	ADF
BAD	0	-1.580	LSX	1	-2.658 <sup>(*)</sup>
BAV	0	-2.246	NRW	1	-2.811 <sup>(*)</sup>
BRE	1	-1.686	RHP	0	-0.910
HAM	1	-1.615	SAL	0	-1.715
HES	0	-2.344	SHO	0	-1.846

Notes: \* Significance at a level of 5 %; (\*) significance at a level of 10 %

The 5% critical value is -2.980. Hence the null of a unit root in the savings rate cannot be rejected in any case. At least at the 10% level the rate seems to be stationary for Lower Saxony and North Rhine Westphalia. These results might be substantial i.e. they may indicate no integration of the logs of per capita consumption and income with the vector (1,-1) but they could also reflect the low power of standard unit root tests.

In order to overcome the problem the MADF approach is employed and this is done for two panel sizes. Panel A includes the entire regions, while Panel B excludes Lower Saxony and North Rhine Westphalia. The subpanel is chosen to ensure that any problems arising from potentially different unit root features of the individuals are avoided. The MADF procedure is based on a SUR analysis, where each series is modelled according to the specifications found in the conventional ADF settings.<sup>9</sup> The finite sample distributions have to be derived by Monte Carlo simulation, and details of the methods applied are given in the appendix. Table 2 reports the results of the MADF test for the different panel sizes.

<sup>9</sup> The average contemporaneous correlation between the residuals obtained from separate OLS regressions is about 0.5. The absolute value drops to 0.3 if the analysis relies on mean-adjusted data. Hence, the correlation remains substantial and SUR is recommended.

Table 2: MADF unit root tests for the savings rate

	MADF	5% critical value	10% critical value
Panel A	72.16 <sup>(*)</sup>	72.50	63.47
Panel B	53.23 <sup>(*)</sup>	59.88	53.02

Notes: \* Significance at a level of 5 %; (\*) significance at a level of 10 %

Almost at the 5% level at significance, the null of a joint unit root is rejected in Panel A. The test statistic declines substantially in the smaller panel but it is still significant at the 10% level. The critical values will also be lower in absolute value as correlation is removed from the system in the smaller Panel B. Although the empirical evidence is not overwhelming, the results seems to be in favor of a stationary saving rate per capita. Equivalently the logs of consumption and disposable income are cointegrated, and the cointegrating vector is (1,-1).<sup>10</sup> However a cointegration analysis can not discriminate between the PIH and a Keynesian type consumption function.

Therefore the  $\lambda$ -model which explains the growth rate of consumption by income growth and the lagged consumption-income ratio as the error correction term is analysed. Note that OLS applied to equation (8) will inevitable result in inconsistent estimators. The reason is that changes in actual and permanent income are likely to be positively correlated, where the latter constitutes the error term. The correlation biases the parameter upwards, implying that the relevance of market imperfections is likely to be overestimated.

A solution to this problem is to use the instrumental variable approach. In particular, lagged growth rates of consumption and disposable income, a constant, a linear time trend and impulse dummies are treated as instruments, and the concrete set of variables is allowed to differ across the panel members.<sup>11</sup> The impulse dummies capture shocks

<sup>10</sup> Wolters (1992) and Hassler (2001) also report a cointegrating relationship between consumption and disposable income, while in Reimers (1997) the long run relationship can only be established after a wealth variable is included in the analysis.

<sup>11</sup> Details are available from the authors upon request. The lagged saving rate is insignificant in these regressions and therefore it does not serve as an instrument. This implies that disposable income is weakly exogeneous with respect to the cointegration structure, and thus a single equation error correction model can be estimated.

traced to the oil crises and to the German unification, respectively.<sup>12</sup> On average, the adjusted  $R^2$  of a regression of the growth rate of income on its instruments is approximately 0.6, where the lowest figures 0.3 and 0.5 are computed for the two city states Bremen and Hamburg, respectively. The fitted values replace the original regressor i.e. the growth rate of income, which renders the estimation consistent.

However, IV estimation is not efficient. Efficiency gains can be realized by exploiting the contemporaneous correlation structure across the regions. This is the rationale of a SUR analysis. Moreover, in this framework cross equation restrictions can be evaluated for example by the means of standard Wald tests. In particular, the hypothesis of an equal feedback parameter, i.e. the coefficient of the error correction term, can be broadly accepted. The test statistic is distributed as chi-square with  $N-1=9$  degrees of freedom. It has a value of 4.74 while the 5% critical value is 16.9.<sup>13</sup> The results of the restricted error correction model are presented in Table 3.

Table 3: IV-SUR estimation of the  $\lambda$ -model (8)

	$\Delta y_t$	$ec_{t-1}$		$\Delta y_t$	$ec_{t-1}$
BAD	0.439 (6.956)		LSX	0.507 (7.872)	
BAV	0.497 (9.326)		NRW	0.368 (6.297)	
BRE	0.519 (4.853)	-0.075 (6.099)	RHP	0.453 (6.777)	-0.075 (6.099)
HAM	0.374 (4.276)		SAL	0.531 (5.884)	
HES	0.467 (5.636)		SHO	0.445 (6.716)	

Notes: The terms in brackets are the  $t$ -values,  $ec_t=c_t-y_t$  is the consumption-income ratio with  $c_t=\log(C_t)$  and  $y_t=\log(Y_t)$ . As explained above, instruments are used instead of the original  $\Delta y_t$  series.

All regressors are significant at the 1% level. Both the PIH and an ordinary keynesian approach seem to explain the growth rate of consumption per capita. The estimator of  $\lambda$

<sup>12</sup> Dummies are not needed in the case of the panel unit root tests because of co-breaking, see Hassler (2001). The shocks affect jointly consumption and income and their net impact, which is included in the savings rate, is negligible.

<sup>13</sup> In addition, the restriction of an equal  $\lambda$ -parameter is not rejected either, but it is not applied in the following. Without this restriction the interpretation of is easier.

suggests that roughly 45 percent of disposable income is earned by non-optimizing consumers.<sup>14</sup> This figure is broadly comparable across the panel regions, where North Rhine Westphalia (37), Hamburg (37) and Saarland (53) mark the edges of the scale. The complement is received by optimizing agents and therefore, the PIH is still important for explaining consumption behaviour in Germany. Furthermore the results indicate a similar consumption behaviour across the regions.

Finally, the correlation between the individual  $\lambda$  parameters and measures of regional economic performance is considered. This may shed some light upon the sort of the market imperfections. For example, the correlation between the  $\lambda$  coefficients and the mean regional unemployment rates is 0.29 and this number is significant positive at the 5% level. If unemployment serves as a proxy for credit rationing, the findings of the  $\lambda$ -model can be partly traced to the liquidity constraints issue.<sup>15</sup>

## 6 Conclusion

In this paper the permanent income hypothesis (PIH) of consumption is examined on the basis of West German regional data. Standard unit root tests fail to reject the joint null of a unit root in the saving rate, mainly due to the power of these procedures. In fact panel unit-root tests which are robust to cross section correlation indicate the stationarity of the saving rate. This finding supports the PIH, but also other types of the consumption function.

In particular, the relevance of non-optimizing consumers is examined within the  $\lambda$ -model which is extended by an error correction term. Here approximately a fraction of 45 percent of disposable income is earned by households that do not to behave according to the PIH. The share is positively related to measures of regional economic performance i.e. the unemployment rate and may therefore reflect the relevance of liquidity

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<sup>14</sup> The only previous result for Germany is reported by Wolters (1992). He estimates the model without the error correction term and finds a fraction of 0.29 for the whole West German economy.

<sup>15</sup> The assumption that an individual that is unemployed is also faced by credit restrictions merely serves as an approximation. Specifically, a distinction should be drawn between sustained and temporary periods of unemployment. Individuals who expect to be permanently unemployed might behave according to the PIH, and this potentially explains the relative low correlation. However, the data are only available at a yearly frequency and therefore, the argument can not be stressed further.

constraints. All in all the results are broadly comparable for the panel members, thus indicating similar consumption behaviour across the regions.

## Appendix

Due to the dependence between the cross section units, the finite sample distribution of the MADF statistic has to be derived by simulation. The critical values will depend on both the cross section and time series dimension of the panel and on the contemporaneous correlation structure as well. The Monte Carlo experiments are based on 10 autoregressive models describing the fluctuations in the regional saving rates, and the lag structure is reported in Table 1. After performing OLS for the cross sections separately, the contemporaneous covariance matrix of the residuals is computed. This matrix is employed in executing the Monte Carlo experiments.

In particular, error series of length  $100+T$  are drawn from the distribution. To reduce the sensitivity of the results on initial conditions the first 100 observations are discarded, leaving time series of length  $T=30, 50, 100$  and  $250$  for the analysis. Using the estimated values of the coefficients of the lagged variables, pseudo saving rates are generated. Then the system corresponding to equation (9) is estimated by the SUR technique and the Wald test for the joint null of a unit root for all panel members is performed. In

Table 4: Critical values for the MADF test

		N=8				N=10			
$\alpha$	T				T				
	30	50	100	250	30	50	100	250	
10%	53.02	39.66	30.61	24.32	63.47	47.99	37.42	29.75	
5%	59.88	44.78	34.53	27.50	72.50	53.41	41.38	33.26	
1%	76.78	55.94	42.10	33.69	94.08	63.04	49.23	40.73	

Notes: N = number of cross-section units; T = number of time periods;  $\alpha$  = level of significance.

order to find the empirical distribution of the test statistic, the process is repeated 5000 times. The results for the different panel sizes examined in this paper are reported in Table 4. For example at the 5% level of significance and 30 time periods unit the critical value is 72.50, when all cross sections are included ( $N=10$ ). If the MADF test exceeds this level, the joint null is rejected. Note that the critical values depend on a particular contemporaneous correlation structure. Hence they are specific to the analysis and cannot be applied in other studies. The TSP386 program which generates the critical values is available from the author upon request.

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