MF Working Paper Series

No 02-2010 January 2010

Magdalena Zachłod-Jelec

Interrelations between consumption and wealth in Poland

Ministry of Finance in Poland
Financial Policy, Analyses and Statistics Department

Interrelations between consumption and wealth in Poland

Magdalena Zachłod-Jelec *

January 7, 2010

Abstract

This paper studies the long-run relationship between consumption, labour income and asset wealth in Poland. Within cointegrated VAR model dynamic responses of the variables in the system to shocks are studied. In addition series are decomposed into permanent and transitory components.

Main conclusion of this paper is that deviations of the three variables from their estimated long-run relationship are better explained with fluctuations of labour income than assets. A tentative explanation of this finding is presented. Additionally, the magnitude of the asset wealth effect in Poland is calculated and compared with other studies for European countries and for the U.S.

Keywords: wealth, cointegration, Beveridge-Nelson decomposition, impulse responses

JEL Classification: E21, C32

*Ministry of Finance, Financial Policy, Analysis and Statistics Department, e-mail: magdalena.zachlod@gmail.com; Ph.D. student at the Warsaw School of Economics

1 Introduction

According to economic theory, consumption should react only to permanent changes in wealth (defined as total resources of consumers). To the extent that consumers perceive certain asset prices fluctuations as a temporary phenomenon, they should not adjust their consumption. If temporary fluctuations of wealth leave consumption unaffected then it should be possible to identify them with fluctuations in consumption-wealth ratio. The pioneering papers by Lettau and Ludvigson investigated whether consumption-wealth ratio predicts excess returns (Lettau, Ludvigson, 2001) and reinterpreted commonly understood asset wealth effect on the basis of the permanent-transitory decomposition of assets (Lettau, Ludvigson, 2004). Lettau and Ludvigson (2004) found that only a small fraction of the variation in household net assets in the U. S. is related to variation in aggregate consumer spending. Similar analysis to Lettau and Ludvigson's was conducted by Fernandez-Corugedo et al. (2003) for the UK and Hamburg et al. (2005) for Germany. While the conclusions obtained by Fernandez-Corugedo and his co-authors for the UK do not differ substantially to what Lettau and Ludvigson found, Hamburg and her co-authors presented some contrasting results for German economy. Their conclusion was that deviations of the consumption-wealth ratio from the long-run trend predict changes in labor income rather than changes in asset wealth. In addition, they found that assets in Germany have predominant permanent component.

In the spirit of the above tradition, this paper is an attempt to investigate the consumption-wealth link for Poland. As there are no estimates of the housing wealth or other tangible assets we only account for the financial assets in our analysis. The main goal of the paper is twofold. First, using the cointegrated VAR framework we study the interrelations between consumption and wealth in Poland with particular attention attached to the permanent-transitory decomposition of the series. Second, we make an attempt to calculate asset wealth effect for Poland and compare it with other European countries as well as the U.S.

The paper is organized as follows. Section 1 introduces. In section 2 we present derivation of the consumption-wealth ratio from the theory. Section 3 covers empirical analysis. In section 3.1 we discuss data and investigate the properties of the data generating processes for the series. In section 3.2 we show how the theory can be applied within the cointegrated VAR framework. We present decomposition of the series into permanent and transitory components in section 3.3. In section 3.4 we present the application of the earlier results to interpret asset wealth effect. Section 4 concludes.

2 Theoretical background

Following Lettau and Ludvigson (2001, 2004), we assume an representative agent economy in which all wealth is tradable. Defining W_t as total wealth (comprised of human wealth, H_t, and asset wealth, A_t) at the begining of period t, $R_{w,t+1}$ as the net return on aggregate wealth, the law of motion for total household wealth can be written as:

$$W_{t+1} = (1 + R_{w,t+1})(W_t - C_t). \tag{1}$$

Human wealth is defined as $H_t = E_t \sum_{j=0}^{\infty} \prod_{i=0}^{j} (1 + R_{h,t+i})^{-i} Y_{l,t+j}$, where R_{ht} denotes net return on human wealth.

Campbell and Mankiw (1989) show that for stationary ratio of consumption and aggregate wealth, the consumer's budget constraint can be approximated by taking first-order Taylor expansion of (1). Solving the resulting difference equation for logarithm of wealth forward, imposing that $\lim_{i\to\infty} \rho_w^i (c_{t+i} - w_{t+i}) = 0$ (so-called "transversality condition") and taking expectations it is possible to obtain (lowercase letters denote logarithms of adequate variables in levels and we omit linearization constants):

$$c_t - w_t = E_t \sum_{i=1}^{\infty} \rho_w^i (r_{w,t+1} - \Delta c_{t+i}), \tag{2}$$

where E_t is the expectation operator conditional on the information available at time t, ρ_w is interpreted as an average share of invested wealth in total wealth ($\rho_w \equiv 1 - exp(\overline{c - w})$) and we defined r as equal to $log(1 + R)$. Assuming that logarithm of aggregate wealth can be approximated as the sum of logarithm of assets and logarithm of human wealth it is possible to write:

$$w_t \approx \gamma a_t + (1 - \gamma) h_t, \tag{3}$$

where γ equals the average (steady state) share of asset wealth in total wealth, $\gamma = \frac{A}{W}$. It can also be expressed in terms of a steady-state non-property income and returns as $\gamma = R_h A/(Y_l + R_h A)$ (see Lettau, Ludvigson, 2001). Consequently, return on wealth can be approximated as a weighted average of returns from assets and human wealth (it was showed by Campbell, 1996):

$$r_t \approx \gamma r_{at} + (1 - \gamma) r_{ht}. \tag{4}$$

Using (3) and (4) allows transforming (2) into the following formula:

$$c_t - \gamma a_t + (1 - \gamma) h_t = E_t \sum_{i=1}^{\infty} \rho_w^i \{[\gamma r_{a,t+i} + (1 - \gamma) r_{h,t+i}] - \Delta c_{t+i}\}. \tag{5}$$

On the left-hand side of equation (5) an unobservable human wealth appears what makes it impossible to apply this formula directly in an empirical analysis. Interpreting aggregate non-property income (i.e. broadly defined labor income) as the dividend on human wealth and using adequate approximation for logarithms, it is possible to write (see Lettau, Ludvigson, 2001):

$$h_t = \delta + y_{lt} + \upsilon_t, \tag{6}$$

where δ is a constant, y_{lt} – aggregate non-property income and υ_t is a mean-zero stationary random variable given by $\upsilon_t = E_t \sum_{i=1}^{\infty} \rho_h^i [\Delta y_{l,t+i} - r_{h,t+i}]$. In formula (6) nonstationary component of human wealth is captured with non-property income.
Assuming for simplicity $\rho_w = \rho_h$, using (3) and (6), we yield an expression that can be applied in further analysis as it uses only observable variables on the left-hand side:

$$c_t - \gamma a_t + (1-\gamma) y_{lt} \approx E_t \sum_{i=1}^{\infty} \rho_w^i (\gamma r_{a,t+i} + (1-\gamma)\Delta y_{l,t+1+i} - \Delta c_{t+i}). \tag{7}$$

Interpretation of the above formula is crucial for the analysis in this paper. First, left-hand side of (7) describes the long-run relationship between consumption, asset wealth and non-property income. This formula may be interpreted as an logarithmic approximation to the "stochastic" version of the permanent income model as proposed by Hayashi (1982):

$$C_t = \varphi(A_t + H_t) + \xi_t = \varphi A_t + \varphi \omega Y_{lt} + \xi_t. \tag{8}$$

It states that permanent consumption, C_t, is proportional to aggregate wealth expressed as a sum of asset wealth and human wealth, $(A_t + H_t)$ while taking into account the transitory consumption or measurement error, ξ_t. Second, if non-property income follows a random walk and the expected return to human wealth is either constant or proportional to the expected return to asset wealth, left-hand side of formula (7) is proportional to the logarithm of consumption-wealth ratio, $c_t - w_t$, so it is possible to refer to the left-hand side of the equation (7) as consumption-wealth ratio. Third, the logarithm of aggregate non-property income in (7) captures the non-stationary component of human wealth. Forth, if right-hand side variables are stationary, the left-hand side variables are cointegrated and the right-hand side of (7) is equal to the cointegrating residual. Fifth, if the cointegrating residual is not constant, it must forecast either one of the three variables from the right-hand side of expression (7), or some combination of them.

Before we proceed with cointegration analysis of the three-variable system, we first examine the properties of the data generating processes (DGPs) for them.

3 Empirical implementation

3.1 Data

Our data vector will be referred as $x_t' = [c_t \quad y_{lt} \quad a_t]'$ where c_t is aggregate households' consumption expenditure, y_{lt} is aggregate non-property income and a_t stands for financial assets. All variables are expressed in logarithms, per capita, in constant 2000 prices. Our data are quarterly, seasonally unadjusted and span the first quarter of 1995 through the second quarter of 2009. For detailed description and graphs of the data, see Appendix at the end of the paper.

As it is visible from the graphs on Figure 1, both consumption expenditure and non-property income are trending and presenting seasonal pattern, albeit much more strong in the data on income than consumption. The value of net financial assets series is growing steadily in time, however at the end of the sample it diminished quite significantly mirroring the correction at the Warsaw Stock Exchange starting from the mid-2007.

As a prerequisite for cointegration analysis, it is necessary to check whether the data generating processes (DGPs) for all series are integrated of the same order. To do so, we conduct two types of unit root test. We run Augmented Dickey–Fuller test (ADF test) and Phillips–Perron test (PP test). As a sensitivity check, we also run Kwiatkowski–Phillips–Schmidt–Shin test (KPSS test) which tests the null hypothesis of stationarity. The results for the unit root and stationarity tests for the series are presented in Table 1. As is visible from the table, all series may be treated as integrated of order one (I(1)). Some conflicting results were obtained for consumption and financial assets. According to the ADF test (at 5% critical level) consumption might be generated by the trend-stationary process, however both KPSS as well as PP test do not confirm this finding.Test results also show that the DGP for financial assets might be integrated of order between one and two: both ADF and PP tests strongly reject the null hypothesis that first differences of the process contain unit root, but KPSS test is unable to reject the null of stationarity of the series only at 1% critical level.

3.2 VECM analysis

Our further analysis will require a correctly specified vector error correction model (VECM). Accordingly, we first run a VAR model with two lags, here written in a form of an error correction model:

$$\Delta x_t = \mu_0 + \Gamma_1 \Delta x_{t-1} + \Pi x_{t-1} + \Phi D_t + \epsilon_t, \tag{9}$$

$$\epsilon_t \sim IN_p(0, \Omega),$$

where Δx_t is the vector of first differences of the series in logarithms, $[\Delta c_t \quad \Delta y_{lt} \quad \Delta a_t]'$, μ_0 is (3 x 1) vector of constants, Γ_1 is first-order distributed lag operator, D_t is (6 x 1) vector of dummy variables (three centered seasonal dummies and three additional zero-one dummies to correct some instability in the financial assets series at the end of the sample), and Φ is (3 x 6) matrix of dummy coefficients.
We use model with two lags in levels on the basis of the Hannan-Quinn criterion. Schwartz criterion pointed at just one lag, but the difference between the criterion value for one and two lags was negligible, so we decided to incorporate two lags in the final model specification.

Then, we apply the Johansen procedure to our trivariate data vector. As consumption and income reveal apparent seasonal pattern, we include deterministic seasonal components in our model. Based on graphical inspection of the data in log-levels and in differences as well as discussion above, we estimate the model with a trend restricted to the cointegration relations and an unrestricted constant. In such a specification we allow for linear trends in the data and in the cointegrating relations, so we do not *a priori* assume that they "cancel out" in the cointegration space (see Juselius, 2006, p. 100). To correct some residual autocorrelation in our VAR model, we incorporate three dummy variables at the ending part of our sample (see Appendix for description of the dummies). The dummies are needed because of the apparent correction at the Warsaw Stock Exchange connected with global financial turmoil. We might either shorten the sample span for the analysis to the forth quarter of 2007, or use dummies. Choosing the latter possibility we are able to conduct analysis on the whole sample what will prove valuable for our conclusions later on. To check the stability of the model, on Figure 2 in the Appendix we present results for the constancy of the log-likelihood test (scaled by the 95% quantile of the appropriate asymptotic distribution). The graph on Figure 2 confirms that the model is reasonably stable.

Having properly specified model, we test for the number of the cointegrating vectors using the model of the form:

$$\Delta x_t = \mu_0 + \Gamma_1 \Delta x_{t-1} + \alpha \beta' x_{t-1} + \Phi D_t + \epsilon_t, \tag{10}$$

where α is (3 x 1) vector and β is the (3 x 1) vector of cointegrating coefficients. The term $\beta' x_{t-1}$ gives last period's equilibrium error, or cointegrating residual and α is the vector of adjustment coefficients that tells us which variable(s) subsequently adjusts to restore the common trend when a deviation from an equilibrium occurs. Both the eigenvalue as well as the trace test clearly indicate two unit roots meaning that the rank would be equal to one (see Table 3, upper panel). Trace test strongly indicates one cointegration relation with p-value 0.645 for the model without the Bartlett correction and 0.705 for the Bartlett corrected model. The Bartlett correction helps to control the size of the test. Still the power of the test, i.e. the ability to reject the false null hypothesis, might be very low.

We run tests of stationarity in a cointegrated system as a sensitivity check for our earlier results. All variables may be treated as generated by I(1) processes and neither of them seem to be trend-stationary (the results are not presented here but are available upon request). We then test whether deterministic trend is actually needed in cointegrating relations. On the basis of the likelihood ratio (LR) test (Chi-square distributed with one degree of freedom) with p-value of 0.436 we exclude trend from the cointegration relation. Residuals from the final model do not exhibit autocorrelation nor heteroscedasticity and they are close to normal (see Table 2). Again, we test for a number of the cointegration relations. Tests indicate presence of one cointegrating vector in the system. The p-value for the trace test is 0.537 for the model without the Bartlett correction and 0.581 for the Bartlett corrected model (see Table 3, lower panel). Our estimated cointegrating vector for $x_t' = [c_t \quad y_{lt} \quad a_t]'$ is $\beta' = (1 \quad -0.831 \quad -0.133)'$.
As sum of the coefficients of non-property income and assets in the cointegrating vector is close to one, we may test for homogeneity of consumption with respect to income and assets implied by the theory. Indeed, the LR test is unable to reject the null hypothesis of homogeneity with p-value 0.501. Long-run elasticity of consumption with respect to income is equal to 0.881 and with respect to assets: 0.119.

From Granger representation theorem we know that cointegration between variables implies that at least one of them should restore the equilibrium what means that in the error correction representation the respective alpha coefficients are significant. In Table 4 we present estimates from the VECM model (constant and dummy variables coefficients are not shown). None of the variables in the system is predictable by consumption. Growth rates of lagged consumption, non-property income and assets are significant in all equations. An adjustment parameter for consumption is economically small and insignificantly different from zero, which means that consumption does not take part in restoring equilibrium. Weak exogeneity of consumption was tested in the system and not rejected. As it appears from the VECM results, both non-property income and assets are adjusting in the system, albeit income more strongly than assets.

Our conclusions are in line with permanent income hypothesis. First, consumption is unpredictable in a sense that it is generated by a random walk process. This is a major implication of the rational expectations permanent income hypothesis pointed by Hall (1978). Second, consumption does not exhibit error correction and therefore predictability over long horizon. This suggests that consumption has no or only a small transitory component. According to the theory, equilibration should take place via wealth and this reflects the forward looking behavior of households — consumers save in response to expected future changes in income and asset returns (see Fernandez-Corugedo, 2003). Our results, reported in Table 4, indicate that it is indeed the case here. Consumption-wealth ratio, embodied by the cointegrating residual (which we denote "cay", following Lettau and Ludvigson, 2001) predicts non-property income growth, implying that deviations in income from the common trend uncover transitory fluctuations in non-property income. *Cay* is also significant in asset growth equation. However, comparing cointegrating residual (*cay*) with (detrended) consumption to income ratio (*detcy*), we can see that cointegration residual is highly correlated with consumption-income ratio, suggesting that fluctuations in financial assets contribute little to fluctuations in consumption-wealth ratio (see Figure 3). It will be investigated later in more details.

The above analysis may be complemented as follows. Formula (7) implies that cointegrating residual (left-hand side variable) should forecast at least one of the variables that appear on the right hand side. We can test it running long-run regressions as Lettau and Ludvigson (2004) did. In Table 5 we presents the results from regressions of long-horizon consumption growth, Δc_{t+h} (defined as $(c_{t+h} - c_t)$), long-horizon non-property income growth, $\Delta y_{l,t+h}$ and long-horizon assets growth, Δa_{t+h}, on the estimated cointegrating residual and growth of consumption, non-property income and assets over horizons h, ranging from 1 to 16 quarters.
Panel A of the table shows that *cay* is not a significant predictor of the consumption growth at any horizon reflecting the fact that the transitory component for consumption is insignificant. The results are similar with respect to asset wealth (see Panel C of Table 5). The only variable for which cointegration residual has any forecasting power is non-property income, consistent with earlier conclusion that it has an important transitory component. In terms of R^2 statistics, *cay* has the greatest predictive power at 1 to 8 quarters, peaking at 4 quarters (1 year horizon). Non-property income is mean-reverting and adapts over long horizons, however as horizon increases, it is also forecasted by growth of consumption, assets and itself (see Panel B of Table 5).

Contrary to Lettau and Ludvigson (2004) and similar to Hamburg et al. (2005), we found that cointegrating residual does not predict asset returns and it only predicts non-property income. Cointegrating residual contains almost the same information as consumption–non-property income ratio in that it does not predict changes in asset prices but changes in income. This will be further explained with a differences of the households' net financial assets composition between the U.S. (and more general, Anglo-Saxon economies) and European continental countries. For the latter countries fluctuations in labor income are relatively more important in explaining fluctuations in consumption-wealth ratio (see Hamburg et al., 2005).

3.3 Permanent and transitory components of consumption, income and assets

Cointegration may be used to decompose the series into innovations that are distinguished by their degree of persistence. As there is one cointegration vector in the system we consider, there are two common trends and two permanent shocks and, consequently, one transitory shock. Identification of the permanent shocks is straightforward as cointegration restricts the matrix of long-run multipliers of shocks in the system, which identifies the permanent components. The transitory shock is identified as a residual (see e.g. King et al., 1991). The permanent shocks to x_t are defined as (see e.g. Juselius, 2006, p. 278):

$$u_{lt} = \alpha'_{\perp}\epsilon_t \qquad (11)$$

and the transitory shocks are given by:

$$u_{st} = \alpha'\Omega^{-1}\epsilon_t \qquad (12)$$

where $\alpha'_{\perp}$ is transposed matrix of orthogonal complements for α, ϵ_t is a vector of errors from the reduced form VECM model and Ω is their covariance matrix. Such a decomposition of "structural" shocks (u_t) ensures orthogonality between permanent and transitory shocks.

Table 6 presents the variance share of transitory shocks in the forecast error for consumption, assets and non-property income. It is apparent that non-property income is the variable for which transitory shocks matter the most among the variables in the system. However, transitory shocks dominate permanent shocks in the variance of the forecast error of non-property income only in the very short-run, up to two quarters. Since then, the contribution of transitory shocks to the forecast error in non-property income decays quite considerably, reaching nearly 6% at 20-year horizon. The variable to which transitory shocks contribute the least is consumption. Transitory shocks constitute only around 2% of its forecast error variance at horizons starting from 2 years on, being slightly more before. Permanent shocks are also significant for asset wealth. They contribute to the forecast error variance in assets in around 85% at all horizons considered.

Formal decomposition of the series into permanent and transitory components may be done by applying multivariate Beveridge-Nelson decomposition in a way proposed by Garratt et al. (2006). The method takes into account the cointegratedness of the variables in the system, so the decomposition of the series into permanent (or trend) and transitory (or cyclical) components is based on the fundamental underlying stationary economic processes. Following Garratt and his co-authors, Beveridge-Nelson trends are defined as limiting forecasts as the forecast horizon goes to infinity, corrected for deterministic growth, being the value to which all permanent components converge in expectation with the forecast horizon increase:

$$x_t^{BNT} = \lim_{h \to \infty} (E_t x_{t+h} - gh) = \lim_{h \to \infty} (E_t x_{t+h}^P), \tag{13}$$

where x_t^{BNT} is Beveridge-Nelson trend, g – vector of the trend growth rates of the variables, h – forecast horizon, and x_{t+h}^P are permanent components evaluated at horizon h.
As Garratt and others note, Beveridge-Nelson trends may also be expressed in a from they call "infinite horizon error correction" representation:

$$x_t^{BNT} = x_t + \alpha_\infty(\beta' x_t - \kappa) + \Phi_\infty(\Delta x_t - g), \tag{14}$$

where $\beta' x_t$ is the estimated cointegrating vector, κ is the steady state value of a cointegrating relation, α_∞ and Φ_∞ are (3 x 1) and (3 x 3) matrices of expected responses of the variables at infinite horizon in the current disequilibria, respectively in the cointegrating relations and growth rates. In an infinite horizon representiation all disequilibria must in expectation be fully eliminated.

Accordingly, the Beveridge-Nelson trends for consumption, income and assets are calculated using formula (15). The Beveridge-Nelson trends together with actual series are presented on Figure 4. As it is visible from the graphs, consumption is almost indistinguishable from its Beveridge-Nelson trend what is not surprising taking into account that it does not adjust in the system. Similarly, asset wealth series is close to its Beveridge-Nelson trend. Uncontroversially again, the only variable in the system that deviates from its trend significantly is non-property income what is not surprising taking into account that this variable adjusts in the system.

3.4 Asset wealth effect

According to the permanent income theory, consumption should not react to transitory shocks at all. From the variance decomposition of shocks as well as multivariate Beveridge-Nelson decomposition of the series we have already seen that consumption is indeed driven mainly by permanent shocks.

If shocks to financial assets are permanent, financial asset wealth effect can be calculated from the parameters of the cointegrating vector given the knowledge of the consumption to assets ratio value. Long-run elasticity of consumption with respect to financial wealth in the cointegration relation is equal to the product of the marginal propensity to consume out of financial assets and a share of asset wealth in total wealth, denoted by:

$$\beta_a = \frac{\partial C_t}{\partial A_t} \frac{A_t}{C_t} \tag{15}$$

implying that marginal propensity to consume (MPC) out of financial assets can be calculated with the formula: $\frac{\partial C_t}{\partial A_t} = \beta_a \frac{C_t}{A_t}$.
Mean of consumption to financial assets ratio over the sample considered equals 0.35, implying that the MPC out of financial wealth achieves 0.042 in the sample period. Interpretation of this static, average financial wealth effect is as follows: a one-zloty increase in financial assets leads to a 4 groszy increase in consumption expenditure on average per quarter. Due to both assets and consumption have prevailing permanent components, this estimate may capture the marginal propensity to consume out of asset wealth quite well.

Lettau and Ludvigson (2004) found similar magnitude of the long-run MPC out of total (i. e. financial and tangible) assets for the United States, but according to their analysis this result applies to only around 12 percent of the total variation in wealth, as most of wealth fluctuations are transitory and thus dissociated from mostly permanent fluctuations in consumption. Almost identical MPC out of total assets was found by Fernandez-Corugedo et al. (2003) for the UK. In contrast to Lettau and Ludvigson (2004), Fernandez-Corugedo and his co-authors concluded that only up to 30 percent variations in assets are transitory meaning that permanent shocks dominate assets behavior in UK. They explain it with a greater importance of the housing wealth in total wealth in UK in comparison to the USA and they argue that shocks to housing wealth are disproportionately permanent compared to other wealth. For Germany, Hamburg et al. (2005) calculated long-run MPC out of total assets to be around 0.044. Very recently, using various wealth measures, Sousa (2009) analyzed wealth effects for the euro area and found long-run MPC out of financial assets to be in the range of 0.0143 and 0.0175. He also found that expanding asset measure with housing wealth reduces total asset wealth effect significantly (to 0.0042) as negative effect of house price increases for renters outweigh the positive effect for current homeowners.
Discussing quite high MPC out of assets for Poland obtained here, it should be stressed first, that it is mainly a result of the relatively high in-sample consumption to financial assets ratio, and second, that it does not include tangible assets as no reliable estimates for Poland exist (especially for hous-

ing).

Similarly to Hamburg et al. (2005) and contrary to Lettau and Ludvigson (2004), we found that shocks to assets are predominantly permanent. It is closely linked to the composition of households financial assets. In continental European countries direct ownership of stocks is very limited comparing with the UK and the USA. In Poland stocks held directly by households constitute around 5% of households' net financial assets (on average between the forth quarter of 2003 and the second quarter of 2009). Stocks ownership of individuals peaked in mid-2007 reaching 9%, but it was falling down since then due to the global financial turmoil coming back to 5% at the end of the second quarter of 2009. If shocks to the other financial assets components are disproportionately permanent compared to stocks, this may explain why in Poland financial assets are dominated by permanent shocks.

To investigate dynamic interactions between consumption, assets and income, in Figure 6 we graph impulse responses together with 95% Hall percentile confidence intervals based on 2000 bootstrap replications. Impulse responses are based on a decomposition of structural shocks into orthogonal permanent and transitory shocks as well as orthogonalization of the two permanent shocks to one another. To achieve identification of the structural VAR, we also assume that consumption does not have instant impact on assets.
In the first column of Figure 5 we show the responses of the variables to the permanent shock to consumption. Middle column presents responses to the transitory shock, which may be treated as a shock to income and in the last column there are presented responses for the second permanent shock which we interpret as a shock to asset wealth. In accordance with our earlier conclusions, transitory shock has virtually no effect on consumption and quite small effect on assets, while the response of non-property income is considerable. It takes about 3 years for all the variables to adjust completely to this transitory shock. After a permanent shock to consumption, consumption reaches its new level immediately and non-property income reaches its new level gradually, after about 4 to 8 quarters. In line with economic theory, consumption 'overshoots' both asset wealth and income in the short run to adjust to its new permanent level immediately. After the permanent shock to consumption, the value of assets drops down. The permanent shock to assets affects all the variables in the system positively. Asset wealth reacts most strongly to this shock and the effect on consumption and non-property income is similar in magnitude to one another.

4 Conclusions

This paper has investigated the consumption-wealth link in Poland during the first quarter of 1995 and the second quarter of 2009. Using the cointegrated VAR framework we estimated long-run relationship between consumption, non-property income and financial assets derived from a theory. In line with permanent income theory, we found that consumption reacts mainly to permanent innovations in assets and income. Contrary to Lettau and Ludvigson (2004), we found that consumption-wealth ratio of an average Polish household does not help to identify the transitory components in Polish stock market. Instead, consumption-wealth link is important in explaining transitory fluctuations in non-property income. Our results are similar to these obtained by Hamburg et al. (2005) and are therefore another confirmation of the differences between Anglo-Saxon and European continental economies in terms of the financial assets structure of an average household.

As consumption and financial assets are affected primarily by permanent shocks, it is possible to calculate a static asset wealth effect measured by the long-run marginal propensity to consume out of assets. The magnitude of this effect for Poland turned out to be in range of estimates for countries with more developed financial markets, in spite of obvious differences in the portfolio decomposition of households between Poland and more developed countries. But the results obtained here should be interpreted with caution due to at least two reasons. First, they stem from quite high in-sample consumption to assets ratio and second, what may also overestimate the asset wealth effect, as recent study by Sousa (2009) showed, we only use data on financial assets of households.
Complementary to this static asset wealth effect on consumption, we also study dynamic interactions between consumption and assets by means of impulse responses from the structural VAR model.

The analysis conducted here may be confronted with respect to at least two dimensions. First, it will be useful to investigate the interrelations between consumption and a wider measure of assets including also tangible, especially housing wealth. Second, as the sample is quite short and at the end part affected by the financial crisis, it would prove valuable to repeat the analysis using longer sample and check the stability of the results, particularly with respect to asset returns predictability.

References

[1] Campbell J. Y., (1996), *Understanding Risk and Return*, "Journal of Political Economy", 104, 298-345.

[2] Fernandez-Corugedo E., Price S., Blake A., (2003), *The Dynamics of Consumers' Expenditure: the UK Consumption ECM Redux*, Working Paper no. 204, Bank of England.

[3] Garratt A., Robertson D., Wright S., (2006), *Permanent vs. Transitory Components and Economic Fundamentals*, "Journal of Applied Econometrics", 21, 521-542.

[4] Hall R. E., (1978), *Stochastic Implications of the Life Cycle-Permanent Income Hypothesis: Theory and Evidence*, "Journal of Political Economy", 96, 971-987.

[5] Hamburg B., Hoffmann M., Keller J., (2005), *Consumption, Wealth and Business Cycles: Why is Germany Different?*, Discussion Paper Series 1: Economic Studies, no. 16, Deutsche Bundesbank.

[6] Hayashi F., (1982), *The Permanent Income Hypothesis: Estimation and Testing by Instrumental Variables*, "Journal of Political Economy", 90, 895-916.

[7] Juselius K., (2006), *The cointegrated VAR model. Methodology and applications*, Oxford University Press, Oxford.

[8] King R. G., Plosser C. I., Stock J. H., Watson M. W., (1991), *Stochastic Trends and Economic Fluctuations*, "American Economic Review", 81, 819-840.

[9] Lettau M., Ludvigson S., (2001), *Consumption, Aggregate Wealth and Expected Stock Returns*, "Journal of Finance", 56, 815-849.

[10] Lettau M., Ludvigson S., (2004), *Understanding Trend and Cycle in Asset Values : Reevaluating the Wealth Effect on Consumption*, "American Economic Review", 94, 276-299.

[11] Sousa R. M., (2009), *Wealth Effects on Consumption. Evidence from the Euro Area*, Working Paper Series, no. 1050, European Central Bank.

Appendix

Data description

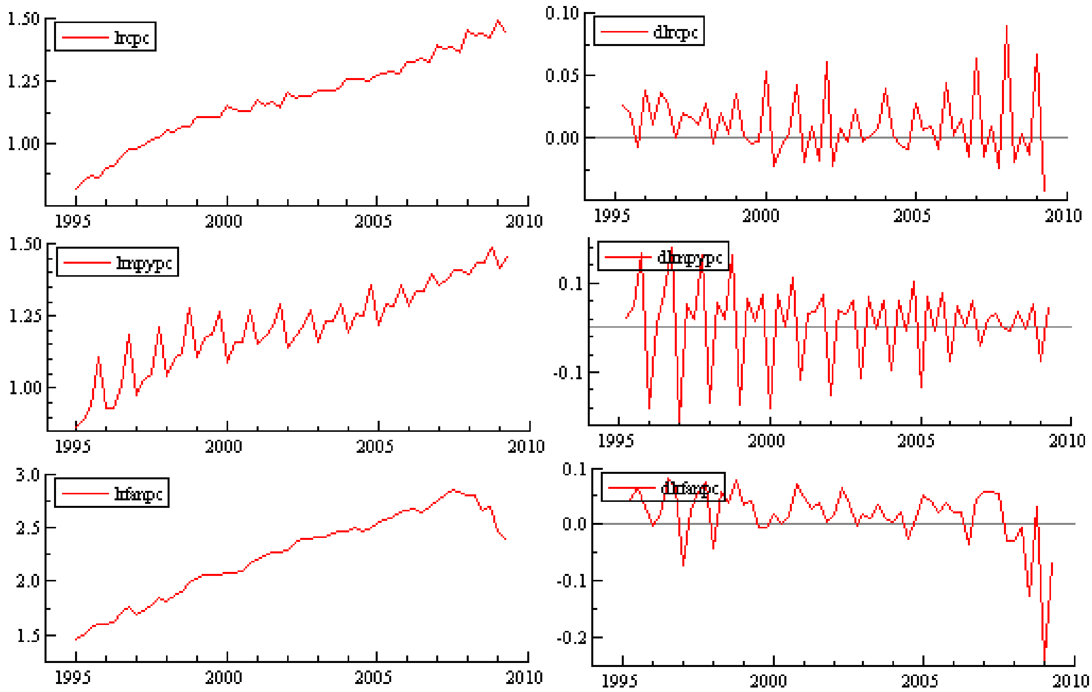
Consumption
Consumption is defined as households' consumption expenditure, in constant 2000 prices, on per capita basis. The source is Central Statistical Office.

Non-property income
Non-property income is defined as households' gross disposable income minus households' property income, deflated by the consumption deflator, on per capita basis. The source for disposable income and property income is Central Statistical Office.

Financial assets
Net financial assets are defined as households' financial assets minus households' liabilities, deflated by the consumption deflator, on per capita basis. Both assets and liabilities for households are taken from the balance sheet accounts of National Bank of Poland. Data prior to 2003 was estimated by the author on the basis of available data on assets and liabilities components.

Dummy variables
Centered seasonal dummies (D_t^{qi}) are defined in a way that they equal 0.75 for the relevant quarter i and -0.25 in quarters $i+1$, $i+2$ and $i+3$.
D_t^{08q1}, D_t^{08q3} and D_t^{09q1} are defined in a way that they equal 1 in the relevant quarter and 0 elsewhere.

Figure 1. Data graphs



Left panel: data in log-levels, right panel: first differences of the series; lrcpc - logarithm of real households consumption expenditure, lrnpypc - logarithm of real non-property income of households, lrfanpc - logarithm of real financial assets of households.

Table 1. Unit root and stationarity tests

series	ADF test	PP test	KPSS test
Deterministic components in the test regression:: intercept and trend			
c_t	-3.76873**	-3.08635	0.175**
y_{lt}	-2.03153	-2.26919	0.147**
a_t	1.34020	2.44564	0.223**
Deterministic components in the test regression:: intercept			
c_t	-1.36160	-2.44986	1.233***
y_{lt}	-0.82990	-1.23716	1.205***
a_t	-1.61040	-2.24416	1.191***
Δc_t	-2.63581*	-7.73724***	0.462*
Δy_{lt}	-3.97713***	-7.48286***	0.192
Δa_t	-5.56104***	-5.85889***	0.535**

Notes: Lag length for ADF test was chosen with Akaike Information Criterion; Lag length for PP and KPSS tests was set to 4 for log-levels and 3 for differences; *** means rejection of the null at 1% crit. level; ** means rejection of the null at 5% crit. level; * means rejection of the null at 10% crit. level.

Source: Own calculations

Table 2. Residual analysis from the cointegrated VAR model

Residual S.E. and Cross-Correlations

	DLRCPC	DLRNPYPC	DLRFANPC
	0.01141333	0.02440789	0.02671985
DLRCPC	1.000		
DLRNPYPC	0.535	1.000	
DLRFANPC	0.481	0.377	1.000

LOG(|Sigma|)		=	-24.243
Information Criteria:	SC	=	-21.439
	H-Q	=	-22.303
Trace Correlation		=	0.767

Tests for Autocorrelation

Ljung-Box(14):	ChiSqr(108)	=	128.137	[0.090]
LM(1):	ChiSqr(9)	=	8.579	[0.477]
LM(2):	ChiSqr(9)	=	11.178	[0.264]

Test for Normality:	ChiSqr(6)	=	3.289	[0.772]

Test for ARCH:

LM(1):	ChiSqr(36)	=	33.138	[0.605]
LM(2):	ChiSqr(72)	=	83.456	[0.168]

Univariate Statistics

	Mean	Std.Dev	Skewness	Kurtosis	Maximum	Minimum
DLRCPC	0.000	0.011	-0.039	3.235	0.030	-0.030
DLRNPYPC	0.000	0.024	0.158	2.553	0.054	-0.052
DLRFANPC	0.000	0.027	-0.553	2.870	0.052	-0.067

	ARCH(2)		Normality		R-Squared
DLRCPC	9.721	[0.008]	1.320	[0.517]	0.799
DLRNPYPC	0.202	[0.904]	0.393	[0.822]	0.929
DLRFANPC	2.432	[0.296]	3.971	[0.137]	0.740

Source: Output from the CATS package

Table 3. Trace test and modulus of the three largest roots of the companion form matrix

r	p-r	Eigenvalues	Trace	Trace Bartlett corrected	p-value	p-value Bartlett corrected	Modulus of the three largest roots r=3	r=2	r=1
				Model with trend restricted to the cointegration space					
0	3	0.561	60.282	56.035	0.000	0.001	0.851	1	1
1	2	0.204	14.210	13.456	0.645	0.705	0.851	0.846	1
2	1	0.025	1.443	1.219	0.983	0.990	0.674	0.673	0.676
				Model with unrestricted constant					
0	3	0.556	52.875	49.150	0.000	0.000	0.988	1	1
1	2	0.115	7.409	7.024	0.537	0.581	0.892	0.919	1
2	1	0.010	0.576	0.441	0.448	0.507	0.676	0.677	0.677

Source: Own calculations

Figure 2. Test for constancy of the log-likelihood

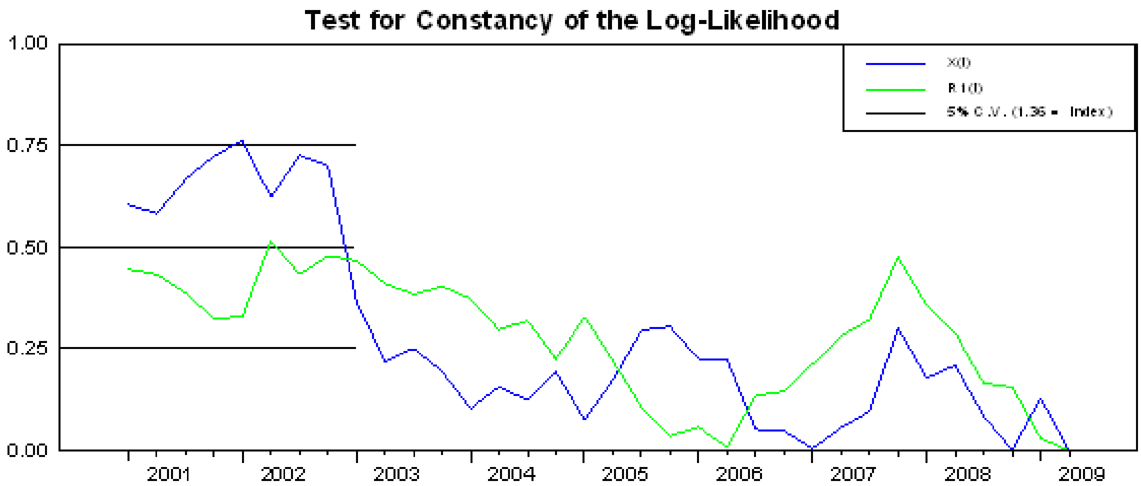


Figure 3. Cointegrating residual vs. detrended consumption to income ratio

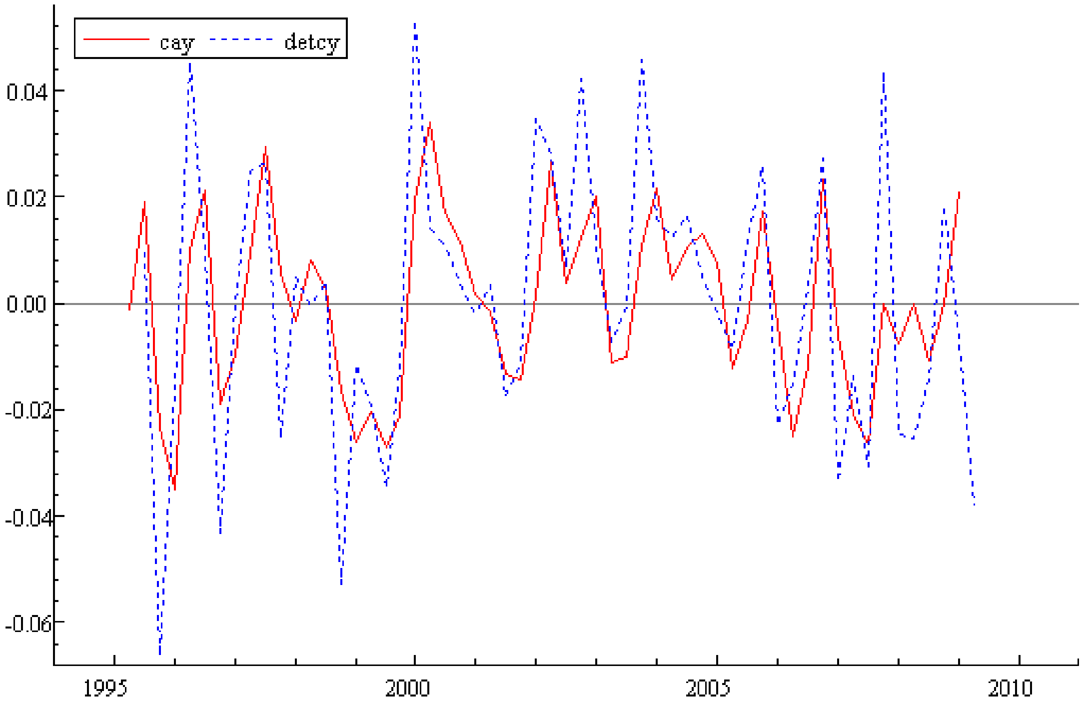


Table 4. VECM estimates

Dependent variable	Equation		
	Δc_t	Δy_t	Δa_t
Δc_{t-1}	-0.118	0.493	0.251
	(-0.83)	(1.68)	(0.78)
Δy_{t-1}	**-0.136**	**-0.394**	**-0.404**
	(-2.38)	(-3.36)	(-3.15)
Δa_{t-1}	**0.09**	**0.158**	**0.251**
	(2.35)	(2.02)	(2.95)
cay_{t-1}	0.051	**1.445**	**0.519**
	(0.47)	(6.45)	(2.12)
$\bar{R}^2$	0.73	0.91	0.68

Notes: Coefficients significant at 5% critical level are in bold face.

Source: Own calculations

Table 5. Long-run regressions

Panel A: $\sum_{h=1}^{H} \Delta c_{t+h}$ regressed on

Horizon H	Δc_t	Δy_{lt}	Δa_t	cay_{t-1}	$\bar{R}^2$
1	**-0.35**	0.005	0.08	0.12	-0.11
	(2.32)	(0.10)	(0.85)	(0.50)	
2	-0.41	0.04	0.35	0.21	-0.61
	(-1.32)	(0.42)	(1.90)	(0.45)	
4	0.02	-0.29	**2.06**	0.66	-1.26
	(0.03)	(-1.51)	(4.79)	(0.66)	
8	3.21	-1.11	**6.64**	3.46	-2.69
	(1.34)	(-1.87)	(4.80)	(1.12)	
12	**10.99**	**-2.80**	**14.13**	5.18	-3.05
	(2.21)	(-2.43)	(5.25)	(0.86)	
16	**20.26**	**-4.47**	**22.31**	9.70	-3.91
	(2.34)	(-2.28)	(4.89)	(0.92)	

Panel B: $\sum_{h=1}^{H} \Delta y_{l,t+h}$ regressed on

Horizon H	Δc_t	Δy_{lt}	Δa_t	cay_{t-1}	$\bar{R}^2$
1	0.24	**-0.49**	0.24	**1.62**	0.32
	(0.56)	(-3.99)	(0.92)	(2.49)	
2	0.62	**-0.89**	0.09	**2.73**	0.36
	(0.84)	(-4.46)	(0.21)	(2.55)	
4	**2.01**	**-1.61**	1.04	**4.52**	0.51
	(2.02)	(-5.98)	(1.71)	(3.10)	
8	**6.93**	**-2.80**	**3.48**	**10.96**	0.34
	(2.22)	(-3.99)	(2.07)	(2.94)	
12	**16.13**	**-3.49**	**6.89**	**18.95**	0.19
	(2.63)	(-2.76)	(2.24)	(2.77)	
16	**24.02**	**-4.40**	**11.60**	**32.39**	0.06
	(2.49)	(-2.27)	(2.44)	(2.99)	

Panel C: $\sum_{h=1}^{H} \Delta a_{t+h}$ regressed on					
Horizon H	Δc_t	Δy_t	Δa_t	cay_{t-1}	$\overline{R}^2$
1	0.33	-0.11	**0.34**	0.52	0.05
	(1.18)	(-1.27)	(2.42)	(1.24)	
2	0.38	-0.33	**1.11**	1.15	0.001
	(0.47)	(-1.40)	(2.34)	(1.09)	
4	0.39	-0.77	**3.90**	2.41	-0.26
	(0.19)	(-1.26)	(2.93)	(0.94)	
8	13.71	0.56	7.20	6.31	-2.78
	(1.82)	(0.29)	(1.68)	(0.88)	
12	24.11	-0.65	**23.50**	4.38	-6.09
	(1.38)	(-0.16)	(2.41)	(0.30)	
16	49.00	0.24	**37.28**	2.62	-11.11
	(1.57)	(0.03)	(2.18)	(0.10)	

Notes: t-statistics in parentheses; significant coefficients at 5% critical level are in bold face; $\overline{R}^2$ is adjusted R^2

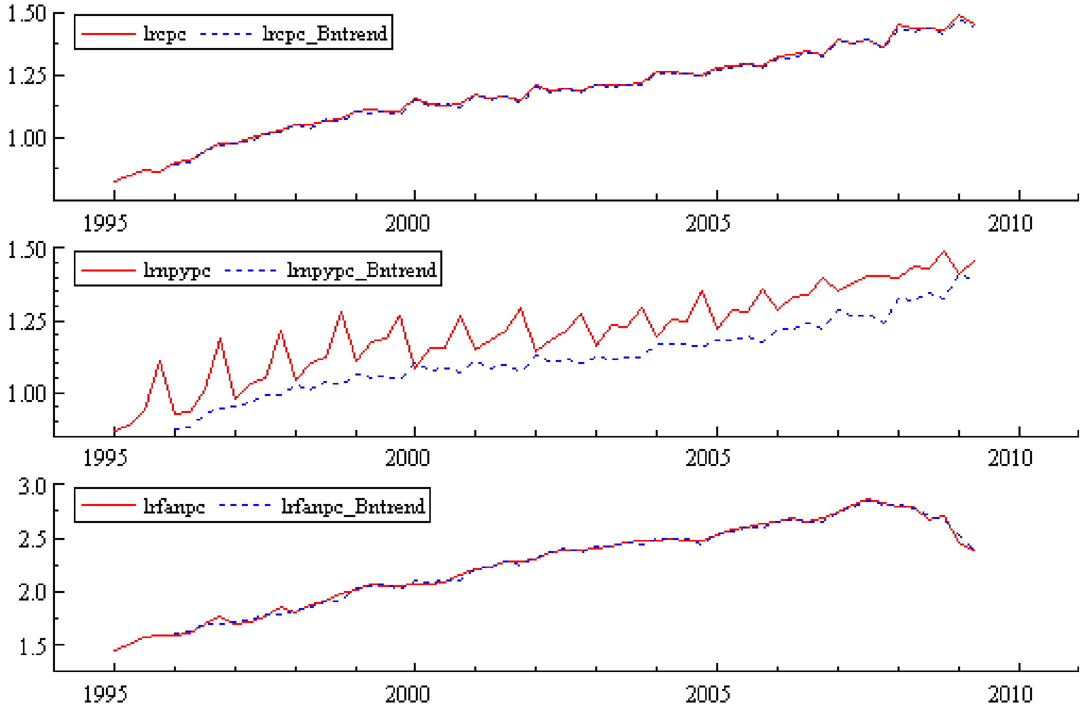
Source: Own calculations

Table 6. Variance share of
a transitory component in forecast errors for consumption, income and assets

	Variance share of transitory component (in percent)								
	Horizon k in quarters								
	1	2	4	8	12	16	20	24	80
$c_{t+k} - E_t(c_{t+k})$	**0.00**	**3.57**	**2.57**	**2.14**	**1.96**	**1.87**	**1.81**	**1.78**	**1.64**
	(0.012)	(0.016)	(0.021)	(0.029)	(0.035)	(0.040)	(0.045)	(0.049)	(0.09)
$y_{l,t+k} - E_t(y_{l,t+k})$	**65.15**	**52.11**	**45.44**	**33.54**	**26.41**	**21.78**	**18.53**	**16.13**	**5.75**
	(0.025)	(0.028)	(0.033)	(0.039)	(0.044)	(0.048)	(0.052)	(0.056)	(0.09)
$a_{t+k} - E_t(a_{t+k})$	**0.00**	**10.28**	**15.01**	**14.25**	**14.13**	**14.09**	**14.08**	**14.07**	**14.03**
	(0.026)	(0.038)	(0.051)	(0.071)	(0.086)	(0.099)	(0.111)	(0.121)	(0.22)

Source: Own calculations

Figure 4. Data series and their Beveridge-Nelson trends



Notes: lrcpc - logarithm of real households consumption expenditure, lrnpypc - logarithm of real non-property income of households, lrfanpc - logarithm of real financial assets of households, BNtrend stands for trends of the respective series from the Beveridge-Nelson decomposition

Figure 5. Impulse responses from the structural VECM model