ON WEALTH VOLATILITY, ASYMMETRIES AND THE AVERAGE PROPENSITY TO CONSUME IN THE UNITED STATES

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Abstract

We explore the relationship between the volatility of wealth and consumption expenditure in the US. Using a GARCH-in-mean VAR model, we find that increases in the volatility of the wealth to income ratio leads to a fall in the average propensity to consume. Further analysis suggests that the effects of negative shocks to the wealth to income ratio have more persistence than equal but opposite positive shocks.

Keywords: volatility; average propensity to consume; wealth; asymmetries.

JEL Codes: E2, E1, C3.

1. Introduction

The recent years surrounding the global financial crisis have underlined the importance of financial markets in affecting consumer behaviour. In particular, the experiences bring into consideration the impact of housing and stock market volatility on consumption expenditure. Further to this, the large negative shock to personal sector wealth associated with the crisis can be contrasted with large positive wealth shocks in earlier decades. This leads one to consider whether asymmetries might be present in the response of consumption to wealth shocks. An improved understanding of how financial market uncertainty and volatility impacts on consumption is of importance to policymakers with an interest in setting stabilisation policy as well as forecasters in general with an interest in cyclical behaviour of the economy. In this paper, we employ a GARCH-in-mean VAR procedure advocated by Elders and Serletis (2010) to focus on these issues. Moreover, we consider whether the average propensity to consume (APC) is driven by the volatility of the wealth to income ratio, and whether equal and opposite wealth shocks leads to an asymmetric response in the APC.

Being based on the largest component of GDP, the APC assumes a position of considerable importance. A high APC may be unsustainable insofar as being associated with low domestic savings. The events of the most recent years surrounding the global financial and debt crises and fall in personal sector wealth have witnessed a decline in the APC. However, an APC that is too low can be associated with economic growth that is relatively reliant on the more volatile components of GDP such as investment expenditure and exports.

While there is an ongoing debate concerning the time-series properties of the APC against a background of structural breaks and nonlinearities, our investigation offers a new perspective and insight into its behaviour. The literature on modelling and estimating consumption functions is extensive. However, studies that explicitly model

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and estimate the determinants of the APC are very limited [see, for example, Arestis and Hadjimatheou (1982), Holmes (1993), Abeysinghe and Choy (2004), Ibrahim and Habibullah (2010) and Holmes and Shen (2013)]. While a common finding in this literature is that the APC is closely related to personal sector wealth, there has been little focus on asymmetries in the behaviour of the APC. Exceptions to this are Ibrahim and Habibullah (2010), who find that the APC exhibits error-correction behaviour only when it is larger than its long-run value, and Holmes and Shen (2013), who point to the possibility that error correction on the part of the APC could be subject to threshold effects. Nonetheless, no study to date has investigated the relationship between wealth volatilities and APC behaviour.

Further to this, very few studies have assessed the impact of wealth volatility and the asymmetric effects of positive and negative wealth shocks in the broader consumption literature. Following Stevans (2004), it can be argued that consumer expenditure may react in an asymmetric fashion to changes in stock market performance on account of the volatility or uncertainty that exists in equity prices during stock market downturns. Adverse selection effects on the part of lenders may contribute to a decline in lending, borrowing and spending. The rise in asymmetric information will also affect the time path towards the lower target spending level associated with increased hysteresis in consumer spending. In contrast, periods of rising equity prices are associated with less uncertainty leading to a smoother adjustment process thereby eliminating the gap between actual and target consumer expenditure relatively quickly. A further rationalization as to why consumption is relatively ‘sticky’ in the downward direction during periods of falling equity prices is based on habit persistence where consumer utility depends on consumption history [Sundaresan (1989). Alexandre et al. (2007) argue that conventional linear estimation of the wealth effect is indeterminate due to the presence of asset volatility. That is, the estimate of the wealth effect should decease when asset volatility increases. On asymmetries, Stevans (2004) and Apergis and Miller (2005a and 2005b) examine whether stock market wealth affects consumer spending asymmetrically and find that the positive share price shock has a stronger impact than an equal but opposite negative stock price shock. Das et al. (2010) show that positive house price shocks have a significant impact on consumption in South Africa, but this is not the case for negative property price shocks. On the other hand, Peltonen et al. (2009) ascertain that a negative financial and housing wealth shock has a larger effect than an equal but opposite positive shock in emerging economies.

Our paper is structured as follows. The following section outlines the GARCH-in-mean VAR procedure that is central to the analysis. The third section reports and discusses our key findings. We find that wealth volatility is a significant driver of the APC and that positive wealth shocks are associated with an overshooting consumption response. The final section concludes.

2. Methodology

Our starting point is to consider a basic empirical consumption function such as

\[ C_t = \sigma + \lambda Y_t + \beta Z_t + w_t \]  

(1)
where \( C \) denotes non-durable consumption expenditure, \( Y \) denotes labour income, \( Z \) denotes personal sector financial and housing wealth, and \( w_t \) is a residual with all series in natural logarithms. Subtracting \( Y \) from both sides and imposing the restriction \( \lambda + \beta = 1 \) enables us to write

\[
APC_t = \sigma + \beta WY_t + w_t
\]

(2)

where the \( APC \) is measured as \((C-Y)\) and explained by the wealth income ratio \((WY)\) which is measured as \((Z-Y)\). Following the approach of Elder and Serletis (2010) in their study of oil price uncertainty, suppose we define \( y \) as the vector \([APC, WY]\) based on \( N=2 \) series, then a GARCH-in-mean VAR model may be written as

\[
By_t = C + \Gamma_1 y_{t-1} + \Gamma_2 y_{t-2} + \cdots + \Gamma_p y_{t-p} + \Lambda(L)H_t^{1/2} + \epsilon_t
\]

(3)

where \( \dim(B) = \dim(\Gamma_1) = (N \times N) \), \( \epsilon_t \mid \psi_{t-1} \sim \text{iid } N(0, H_t) \), \( H_t^{1/2} \) is diagonal, \( \Lambda(L) \) is a matrix polynomial in the lag operator, and \( \psi_{t-1} \) denotes the information set at time \( t-1 \), which includes variables dated \( t-1 \) and earlier. The system is identified by imposing a sufficient number of exclusion restrictions on the matrix \( B \), and assuming that the structural disturbances, \( \epsilon_t \), are uncorrelated. The conditional variance \( h_t \) is modelled as bivariate GARCH, a general version of which is presented in Engle and Kroner (1995) as

\[
h_t = C_v + \sum_{j=1}^{1} F_j vec(\epsilon_{t-j} \epsilon'_{t-j}) + \sum_{i=1}^{1} G_i h_{t-i}
\]

\[z_t \sim \text{iid } N(0, I)\]

\[\epsilon_t = H_t^{1/2} z_t\]

where \( C_v \) is \( N^2 \times 1 \), \( F \) and \( G \) are \( N^2 \times N^2 \), and \( h_t = vec(H_t) \). This specification is too general for most applications, however, with \( 1/2 N(N+1)(N^2 + N + 1) \) distinct variance function parameters for \( J = I = 1 \). This specification also does not ensure that \( H_t \) is positive definite. Re-dimensioning the variance function parameter matrices \( C_v, F, \) and \( G \), the variance function reduces to

\[
\text{diag}(H_t) = C_v + \sum_{j=1}^{J} F_j \text{diag}(\epsilon_{t-j} \epsilon'_{t-j}) + \sum_{i=1}^{J} G_i \text{diag}(H_{t-i})
\]

(4)

where diag is the operator that extracts the diagonal from a square matrix.

The bivariate GARCH-in-mean VAR, equations (3) and (4), can be estimated by full information maximum likelihood (FIML). The procedure is to maximize the log
likelihood \( \sum_{t=1}^{T} l_t \) with respect to the structural parameters \( B, C, \Gamma_1, \Gamma_2, \ldots, \Gamma_p, A, C_v, F, \) and \( G \), where

\[
l_t = -\left( N/2 \right) \ln(2\pi) + 1/2 \ln|B|^2 - 1/2 \ln|H_t| - 1/2 (\varepsilon_t' H_t^{-1} \varepsilon_t)
\]

Following Elder and Serletis, we set the pre-sample values of the conditional variance matrix \( H_0 \) to their unconditional expectation and condition on the pre-sample values \( y_{0t}, y_{1t}, \ldots, y_{(p+1)t} \). To ensure that \( H_t \) is positive definite and \( \varepsilon_t \) is covariance stationary, the following restrictions are imposed: \( C_v \) is element-wise positive, \( F \) and \( G \) are element-wise non-negative, and the eigenvalues of \((F+G)\) are less than one in modulus. Provided that the standard regularity conditions are satisfied, FIML estimates are asymptotically normal and efficient, with the asymptotic covariance matrix given by the inverse of Fisher’s information matrix.

3. Data and Results

All data are obtained from the Bureau of Economic Analysis where quarterly data for the nondurable consumption \( (C) \), labor income \( (Y) \), financial wealth \( (FW) \) and housing wealth \( (HW) \) are expressed in real per capita terms for the study period 1952Q1-2012Q3. Total wealth \( Z \) is based on the addition of \( FW \) and \( HW \). The \( APC \) and \( WY \) series are plotted in Figure 1. While the two series appear move together over time in a positive fashion, it is evident that there have been episodes of sharp relative swings that point towards the possibility of nonlinearities or structural breaks in how they are related.

Figure 1. The APC and Wealth to Income Ratio (W/Y)

We first consider the time series properties of the \( APC \) and \( WY \). Investigations that include King et al. (1991), Harvey et al. (2003) and others point towards non-stationarity of the \( APC \). In contrast, Romero-Avila (2009) and Cerrato et al. (2012), for example, employ panel data unit root testing and find that some form of stationarity is
present, but only after allowing for structural breaks and nonlinearities. Holmes and Shen (2012) find that APC could be stationary, but around a deterministic non-linear trend. Gabriel et al. (2008) argue that nonlinear adjustment may provide a better explanation of fluctuations in the consumption-wealth ratio. Non-stationarity of both these series is initially confirmed in Table 1 which reports ADF tests that are unable to reject non-stationarity at the 10% significance level throughout. Table 2 reports results based on the Perron (1997) unit root tests that allow for a single (unknown) structural break. Again, non-stationarity of the both the APC and WY is accepted throughout.

Table 1. ADF Unit root tests

<table>
<thead>
<tr>
<th></th>
<th>no trend</th>
<th>trend</th>
</tr>
</thead>
<tbody>
<tr>
<td>APC</td>
<td>-0.353</td>
<td>-1.977</td>
</tr>
<tr>
<td>WY</td>
<td>-1.750</td>
<td>-2.246</td>
</tr>
</tbody>
</table>

Notes: in all cases, the lag length is selected according to the Schwarz Information Criterion (SIC).

Table 2. Perron (1997) unit root tests

<table>
<thead>
<tr>
<th>Model:</th>
<th>IO1</th>
<th>IO2</th>
<th>AO</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>$T_b$ $t_{\hat{a}}$</td>
<td>$T_b$ $t_{\hat{a}}$</td>
<td>$T_b$ $t_{\hat{a}}$</td>
</tr>
<tr>
<td>WY</td>
<td>1968Q4 -4.047</td>
<td>1972Q3 -4.384</td>
<td>1977Q1 -3.919</td>
</tr>
</tbody>
</table>

Notes: the models are the Innovational Outlier model (IO1) incorporating a change in the intercept, the Innovational Outlier model (IO2) incorporating a change in the intercept and the slope, and the Additive Outlier (AO) model incorporating a change in the slope only, but both segments of the trend function are joined at the time break. $T_b$ denotes the time of the break and $t_{\hat{a}}$ denotes the test statistic for a unit root. Non-rejection of the null at the 10% significance level is based on a critical value of -4.82 (IO1), -5.25 (IO2) or -4.38 (A0).

Given that the evidence is pointing towards both the APC and WY being non-stationary, the possibility of a cointegrating relationship being present is initially explored using Engle and Granger (1987) testing. Table 3 reports that we are unable to reject the null of non-cointegration. Further exploration based on Gregory-Hansen non-cointegration testing that allows for a variety of structural breaks is presented in Table 4. With test statistics that are well below the 10% critical values, these results point to the absence of a cointegrating relationship between APC and WY.

Table 3. Linear non-cointegration test on the APC and WY

<table>
<thead>
<tr>
<th>$\tau$ (Engle-Granger)</th>
</tr>
</thead>
<tbody>
<tr>
<td>-2.214 (0.418)</td>
</tr>
</tbody>
</table>

Notes: $\tau$ (Engle-Granger) refers to the non-cointegration tests advocated by Engle and Granger (1987), p-values are reported in parentheses.
Table 4. Gregory and Hansen (1996) non-cointegration tests on the APC and WY

<table>
<thead>
<tr>
<th></th>
<th>Level break, no trend</th>
<th>Level break, trend</th>
<th>Full structural break</th>
</tr>
</thead>
<tbody>
<tr>
<td>$T_b$</td>
<td>$t_{\hat{\alpha}}$</td>
<td>$T_b$</td>
<td>$t_{\hat{\alpha}}$</td>
</tr>
<tr>
<td>1982Q4</td>
<td>-3.892</td>
<td>1965Q1</td>
<td>-4.619</td>
</tr>
<tr>
<td></td>
<td></td>
<td>1982Q4</td>
<td>-3.876</td>
</tr>
</tbody>
</table>

Notes: the 10% critical values are respectively -4.34, -4.72 and -4.68 for the level break model with no trend, the level break model with trend, and the full structural break model. $T_b$ denotes the time of the break and $t_{\hat{\alpha}}$ denotes the minimum test statistic for a unit root. In each case, the lag length is determined by the SIC.

Bearing these initial test results in mind, we therefore focus on the short-run relationship between $\Delta APC$ and $\Delta WY$ that excludes an error correction term. Given our interest in assessing the impact of wealth volatility on the average propensity to consume as well as the impact of positive and negative wealth shocks, we estimate the GARCH-in-VAR model described above. We measure uncertainty about wealth as the standard deviation of the one-step-ahead forecast error, conditional on the contemporaneous information set. The standard deviation of this forecast error is a measure of dispersion in the forecast, and as such, is a measure of uncertainty about the impending realization of personal sector wealth. An initial calculation of the respective AIC values for both a standard VAR and the GARCH-in-mean VAR were -2773.5 and -2802.3 which indicates a significantly stronger performance by the GARCH-in-mean VAR model. In our main result reported in Table 5, the estimated coefficient on $H_{1,1}(t)^{1/2}$ which denotes the coefficient on wealth uncertainty (conditional volatility) in the $\Delta APC$ equation is both negative and significant at the 1% level. This confirms that an increased volatility of wealth is negatively related to the $APC$. Our result is consistent with the finding from Alexandre et al. (2007) that the increase in asset wealth volatility will lead to the decline of wealth effect on consumption.

Table 5. GARCH-in-mean VAR estimation

<table>
<thead>
<tr>
<th>Variable</th>
<th>Coefficient</th>
<th>Standard Error</th>
<th>T-statistics</th>
<th>P-value</th>
</tr>
</thead>
<tbody>
<tr>
<td>Coefficient estimate on wealth volatility in $\Delta APC$ equation</td>
<td>$H_{1,1}(t)^{1/2}$</td>
<td>-0.170</td>
<td>0.031</td>
<td>-5.579</td>
</tr>
<tr>
<td>$\Delta(W/Y)$ variance equation for $H_{1,1}(t)$</td>
<td>Constant</td>
<td>0.000</td>
<td>0.000</td>
<td>5.091</td>
</tr>
<tr>
<td></td>
<td>$\varepsilon_{1}(t-1)^2$</td>
<td>0.238</td>
<td>0.066</td>
<td>3.594</td>
</tr>
<tr>
<td></td>
<td>$H_{1,1}(t-1)$</td>
<td>0.628</td>
<td>0.057</td>
<td>10.961</td>
</tr>
<tr>
<td>$\Delta APC$ variance equation for $H_{2,2}(t)$</td>
<td>Constant</td>
<td>0.000</td>
<td>0.000</td>
<td>10.523</td>
</tr>
<tr>
<td></td>
<td>$\varepsilon_{2}(t-1)^2$</td>
<td>0.089</td>
<td>0.037</td>
<td>2.383</td>
</tr>
<tr>
<td></td>
<td>$H_{2,2}(t-1)$</td>
<td>-0.765</td>
<td>0.108</td>
<td>-7.111</td>
</tr>
</tbody>
</table>

Notes: this table is based on the estimation of the GARCH-in-mean VAR model expressed in equations (3) and (4) with the lag length p=3 based on the AIC.
We now consider impulse-response functions based on positive and negative shocks to $\Delta WY$. We simulate the response of $\Delta APC$ for both a positive and negative $\Delta WY$ shock equal to the unconditional standard deviation of $\Delta WY$ to examine whether the responses to equal but opposite positive and negative shocks are symmetric or asymmetric. Monte Carlo methods are used to construct the confidence bands. These are simulated from the maximum likelihood estimates (MLEs) of the model’s parameters. Confidence intervals are generated by simulating 10,000 impulse responses, based on parameter values drawn randomly from the sampling distribution of the MLEs, where the covariance matrix of the MLEs is derived from an estimate of Fisher’s information matrix. Figure 2 displays the impulse-responses based on positive and negative shocks (where DY1 denotes $\Delta WY$ and DY2 denotes $\Delta APC$).

Figure 2.
These responses differ from general literature on the impact of wealth or asset price shocks on consumption which exhibit hump-shaped responses using standard or structural VARs. In absolute terms, our results are characterised by initial magnitudes of the respective responses which are very similar.

However, the bands around the impulse-responses tell us that the persistence associated with a negative shock to $\Delta WY$ is slightly greater than is the case with a positive shock. That is, the significant effects from a negative shock to $\Delta WY$ last for ten quarters whereas the effects from a positive shock to $\Delta WY$ are gone after eight quarters.

We may reflect further on the implications resulting from Stevans (2004), Sundaresan (1989) and others that the change in consumption in response to a fall in share prices should have a slower speed of adjustment than an equal, but opposite, increase in share prices. However, the key difference in the two sets of responses is that positive shocks to $\Delta WY$ stimulate a negative $\Delta APC$ response after two quarters.

This constitutes an overshooting response of $\Delta APC$ to a positive $\Delta WY$ shock. In terms of explaining this, there are some interesting insights from De Veirman and Dunstan (2011 & 2012) who argue that an increase in wealth will encourage households to bring consumption forward in anticipation of low returns on saving. Households initially expect the net worth shock to be permanent, but gradually realize that it is in fact transitory, which means that households will then reduce consumption.

4. Summary and Conclusion

Although the average propensity to consume is a valuable measure to macroeconomists, it is rarely subject to econometric analysis aimed at understanding what drives its behaviour. In this paper we have contributed to filling this gap in the literature. The employment of a GARCH-in-mean VAR model suggests that the average propensity to consume responds negatively to an increase in wealth volatility or uncertainty. In contrast to existing studies of wealth effects on consumption, we find that positive wealth shocks may actually bring about an overshooting response on the part of the average propensity to consume. These findings point towards the value of stabilisation policy that can also calm financial markets and in doing so, prevent the average propensity to consume from being regarded as too high or low.

References


