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Investigating the US consumer credit determinants using linear and non-linear cointegration techniques[☆,☆☆]

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ABSTRACT

This paper has investigated the determinants of total consumer credit for the USA over the period 1968:Q1 to 2011:Q3. Using Breitung's (2001) non-parametric rank tests, we find the existence of linear cointegrating relationships in the consumer credit models. Enders and Siklos' (2001) threshold adjustment tests revealed that non-linearity is present slightly (with a statistical significance of 10% level) in the consumer credit model with a short-term interest rate (federal funds rate), while there exists a linear and symmetric cointegrating relationship in the models with medium (3 years) and long (10 years) term interest rates. Application of the linear cointegrating techniques (fully modified OLS, canonical cointegrating regression and general to specific) show that consumer credit responds more significantly to the medium and long-term interest rates than the short-term interest rate. We use these results to assess the popular belief that abnormality in the consumer credit set the stage for the 2007–08 crisis and severe recession.

1. Introduction

Research on the US consumer credit determinants and in particular evidence that credit is influenced by federal funds rate is very trivial. The monetary policy to have desirable impacts on consumer borrowing, the relationship between consumer credit and federal funds rate is expected to be statistically significant. It is well known that the Federal Reserve Bank (Fed henceforth) cannot control inflation or stimulate output and employment directly; instead, it affects them indirectly, primarily by altering the federal funds rate. This often in the first instance induces investment and consumption spending and then output and employment. In this process, consumer credit does play an important role. Furthermore, investigating the consumer credit demand allows us to assess the popular belief that American consumers 'over-borrowed' during the 1990s and 2000s and that this behavior set the stage for the crisis and severe recession that followed, beginning in 2007–08. To this end, there might be unstable or lack of a well-defined cointegrating relationship for demand for consumer credit.

A number of studies that have modeled consumer credit and examined its determinants for the USA or other countries, made use of the linear cointegration techniques (Hartropp, 1992; Calza et al., 2001, 2003; De Nederlandsche Bank, 2000; Hofmann, 2001; Schadler et al., 2004). There is a risk that theoretical foundations and policy insights that have been formulated based on these studies may be flawed, if indeed, the true cointegration relationship of consumer credit is non-linear. In this paper we explore the total consumer credit — defined as the sum of revolving and non-revolving credit — for the USA considering the demand-side factors viz., real disposable income, real wealth and real interest rates (federal funds rate, 3-year constant maturity rate and 10-year constant maturity rate). Our specification and approach are consistent with the Life Cycle Hypothesis (LCH) of Modigliani and Brumberg (1955). The long-run relationships between consumer credit, income, wealth and interest rates are investigated using alternative specifications and different techniques. In particular, classical linear cointegrating techniques (canonical cointegrating regression, general to specific and fully modified ordinary least squares), Breitung's (2001) non-parametric rank tests, and Enders and Siklos (2001) threshold equilibrium adjustment are applied.

The contribution of this paper is as follows. We examine the consumer credit relationships using two non-linear cointegration techniques (Breitung, 2001; Enders and Siklos, 2001). Breitung's technique is different from testing for non-linear error correction, or testing for non-linear equilibrium correction towards a linear long-run cointegrating relation as suggested by Enders and Siklos (2001). The relationship between the economic variables can be highly non-linear; see Fan et al. (2004). For example, market frictions, heterogeneous agents and official intervention

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could shift the demand for credit; so a constant behavior of consumer credit may not be observed. Our results based on Breitung's (2001) non-parametric rank tests revealed the existence of linear cointegrating relationships in the consumer credit models. Enders and Siklos' (2001) threshold adjustment tests show instead that non-linearity is present slightly (with a statistical significance of 10% level) in the consumer credit model with short-term interest rate (federal funds rate). We use these results to assess the popular belief that abnormality in the consumer credit sets the stage for the 2007–08 crisis and severe recession.

The structure of this paper is as follows. Section 2 presents the theoretical framework and discusses recent empirical studies on consumer credit demand. Empirical results are discussed and presented in Section 3. Finally, Section 4 concludes.

2. Theoretical framework and empirical studies

2.1. Theoretical framework

The LCH offers a classical explanation of why some households borrow to finance consumer spending. According to the LCH, households in the first few years borrow to maintain a desired level of consumption exceeding current income. The gap between consumption and income is financed by borrowing which the households repay with future savings.[1] Our model is a standard two period model and follows the work of Hartropp (1992). Fama (1970) showed that the multi-period problem can be reduced to a two-period problem using dynamic recursive programming. Let the individual maximize utility (Eq. (1)) subject to the constraint (Eq. (2)):

$$U = f\left(C_t, C^e_{t+1}\right) \tag{1}$$

$$C_t + C^e_{t+1} = Y_t + B_t + Y^e_{t+1} + (1+r)B_t \tag{2}$$

where C is the consumer expenditure, Y is the disposable income, B is the increase in net financial liabilities ($B = C - Y$), and r is the real interest rate on borrowing and saving (assumed equal). The superscript e indicates the expected value. The usual first and second order conditions for the maximum are as follows:

$$(\partial U/\partial C_t)>0;\ \left(\partial U/\partial C^e_{t+1}\right)>0;\ \left(\partial^2 U/\partial C_t^2\right)<0;\ \left(\partial^2 U/\partial \left(C^e_{t+1}\right)^2\right)<0. \tag{3}$$

We assume that Y^e_{t+1} depends on income at time t: $Y^e_{t+1} = w_t Y_t$, where w_t is the weight on income in period t. From the first order conditions we know that the ratio of the marginal utility of C_t to the marginal utility of C^e_{t+1} equals $1 + r$. Hence, both C_t and C^e_{t+1} will be determined by Y_t, Y^e_{t+1}, r and the household's relative preference given in Eq. (1) for C_t against C^e_{t+1}. Since $B_t = C_t - Y_t$, we can write:

$$B_t = f(Y_t, w_t \cdot Y_t, r) - Y_t. \tag{4}$$

From the borrowers' perspective the following conditions must be satisfied:

$$f'(Y_t)>0, f'(w_t \cdot Y_t)>0, f'(r)<0. \tag{5}$$

The function f includes the household's preference for consumption today as opposed to consumption tomorrow. While the new borrowing is clearly related negatively to the real interest rate, the overall effect of Y_t on new borrowing is ambiguous. Y_t influences B_t in three ways: (a) one dollar increase in Y_t directly reduces new borrowing by one dollar (assuming no change in C_t); (b) an increase in Y_t of dY_t directly shifts the budget constraint to the right by an amount of dY_t, and therefore tends to increase C_t; and (c) an increase in Y_t shifts the budget constraint up and to the right (by an amount of $w_t \cdot Y_t$) indirectly through its effect on Y^e_{t+1}. Note that (b) and (c) effects tend to offset the effect of (a) and hence the overall effect on B_t is ambiguous; however we leave this data to depict which effect prevails.

Moreover, another important variable in the new borrowing decisions is net wealth (NW) defined as total assets minus total liabilities. An increase in wealth may induce new borrowing. In theory, a positive marginal propensity to consume (MPC) out of wealth, for a given Y_t, induces a higher B_t. The estimated equation therefore becomes:

$$B_t = \alpha_0 + \alpha_1 Y_t + \alpha_2 NW_{t-1} + \alpha_3 r_t. \tag{6}$$

The above model predicts that $r\langle 0, NW\rangle 0$, whereas the sign of Y is empirically determined. We use Eq. (6) for our analysis. We are following a demand-side approach assuming that the demand for consumer credit mainly influences the credit market dynamics.[2] We recognize that the supply factors may play an important role, however the role of demand should not be undermined. For example, consumer finance sector in the U.S. since 1950s was largely driven by the increase in demand for many products and services (Ryan et al., 2011). In the presence of this increasing trend, firms responded with innovations offering consumers more choices and products. The available data on consumer credit demand and supply collected in the Senior Loan Officer Opinion Survey conducted by the Fed seems to confirm this point. Fig. A (see Appendix A) shows that the demand changes were clearly leading the supply changes of credit over the last 20 years. This gives us the intuition that the dynamics of consumer demand are more important with respect to the supply.[3] Nevertheless, credit supply was higher than demand in the Great Recession period. Regrettably we are unable to analyze this aspect of the market due to the data limitations.

2.2. Recent empirical evidence

Tests of the empirical determination of consumer credit are limited. Most studies have utilized the survey data to explore the structure of consumer credits, for instance Jappelli (1990), Cox and Jappelli (1993), Crook (2001), Magri (2002), Crook and Hochguertel (2005), Del-Rio and Yong (2005) and Benito and Mumtaz (2006). Benito and Mumtaz (2006) provide a comprehensive review of this literature. There are a few studies on consumer credit that used aggregated time series data, for example Hartropp (1992), Calza et al. (2001, 2003), De Nederlandsche Bank (2000), Hofmann (2001) and Schadler et al. (2004).

Using the UK data, Hartropp (1992) found that current income and current and past wealth have a positive influence on consumer borrowing, and that the interest rate has a negative effect. Calza et al. (2001) estimated the credit demand for the Euro Area. They found a long-run relationship between credit demand, real weighted short-term and long-term interest rates, and real GDP.[4] Similar analyses on credit demand have been performed in De Nederlandsche Bank (2000) for several EU countries, including Japan and the USA. Using the cointegrating vector autoregression (VAR) model, Hofmann (2001) attained a long-run relationship linking real credit positively to real GDP and real

[1] Other theories may also be relevant. The Permanent Income hypothesis (PIH) theory of consumption suggests that consumer spending depends on permanent income, which gives a low weight in its estimation to current income. In this situation, a rise in income would result in increased saving and not debt. Partly due to this reason, the PIH is unable to explain the facts of consumer credit.

[2] The supply of consumer credit is modeled as being essentially demand determined. To this end, the supply of consumer credit may on the whole adjust directly to meet the demand, with or without the price (interest rate) changes in proportion to excess demand. This theory implies that the quantity of consumer credit traded for a given interest rate is that shown by the demand curve.

[3] A Granger causality test (not reported for brevity but available from the authors upon request) confirms our intuition.

[4] In another study, Calza et al. (2003) considered a new measure of the cost of borrowing, obtained as a weighted average of bank lending rates and extracted information content of the loan overhang/shortfall of the future inflation.

property prices, and negatively to real interest rate for 16 industrialized countries. Schadler et al. (2004) estimated a vector error correction model (VECM) for the Euro Area to find a statistically significant relationship between credit–GDP ratio, real long-term interest rate and real per capita income.

A number of studies examined the asymmetries and non-linearities in credit markets. Stiglitz and Weiss (1981) propose separated models of random credit rationing, with either adverse selection or moral hazard, illustrating why interest rate or collateral is not used to ration credit. Stiglitz–Weiss model is particularly helpful to understand why in some developing countries bank credit is severely rationed with bank lending rates unresponsive to excess demand for credit. Kaufman (1996) used the Stiglitz–Weiss model to explain some of the aspects of Argentina's economic crisis of 1995–96. Using time series data for the US for 1968–1989, Martin and Smyth (1991) find evidence for a backward bending supply curve for mortgages both for a representative loan and for aggregate loan volume. Analogously, Drake and Holmes (1997) find a backward bending supply curve for mortgages using the UK data for 1980s. In an earlier study they also found a backward bending supply curve for non-mortgage consumer credit (Drake and Holmes 1995). Perraudin and Sørensen (1992) use data from surveyed US households and find that the demographic characteristics of borrowers together with their income and job status influence lending decisions of banks. Gambacorta and Rossi (2010) employing the asymmetric vector error correction model addresses possible asymmetries in the transmission mechanism and concludes that the effect of a monetary policy tightening on credit, GDP and prices is larger than the effect of a monetary policy easing.

The existing empirical studies explain the flow of non-mortgage lending to households. However, in the case of the USA, empirical evidence related to the relationship between consumer credit and federal funds rate is limited. Further, the existing time series studies on this topic have exclusively utilized the linear cointegration techniques and gave narrow focus on the non-linearity of the model. If the true relationship is non-linear in nature then the existing findings cannot be relied upon. This study attempts to address these issues.

3. Empirical results

3.1. Data

In this study we use quarterly data from the USA for the period 1968:Q1 to 2011:Q3. Data include real total consumer credit (B = defined as the sum of revolving and non-revolving credit), real disposable income (Y = total income minus taxes), real net wealth (NW), real federal funds rate (r^{ff}), real 3-year constant maturity rate (r^{3Y}) and real 10-year constant maturity rate (r^{10Y}). Non-revolving credit cannot be used again once they have been repaid (for example, student loans) whereas revolving credit does not have a fixed number of payments (for example, credit cards). All data have been seasonally adjusted and are used in natural log form, except for the three real interest rates. The Data appendix provides more details on the definitions and sources of the data. The key descriptive statistics for all variables are presented in Table 1 and the plots of the series are illustrated in Fig. 1.

Table 1
Descriptive statistics 1968:Q1 to 2011:Q3.

	Mean	Std. error	Max	Min
Total consumer credit	2.408	0.497	3.162	1.621
Disposable income	3.567	0.346	4.140	2.960
Federal funds rate	2.104	2.487	9.029	−4.043
3-year constant maturity rate	2.527	2.395	15.720	−3.897
10-year constant maturity rate	3.113	2.280	9.340	−3.847
Net wealth	5.626	0.476	6.448	4.903

Notes: Std. error = standard error, Min = minimum value and Max = maximum value. All variables are expressed in real terms. Total consumer credit, net wealth and disposable income are in natural log form.

3.2. Unit root tests

The integrated properties of the variables are tested with the Lee and Strazicich's (2003) two break minimum Lagrange multipliers (LM) unit root tests. This test is more powerful than the single break tests such as Perron (1989), Zivot and Andrews (1992), Lumsdaine and Papell (1997), among others. Obviously this test is attractive compared to the conventional unit root tests (for example, Augmented Dickey–Fuller and Phillips–Perron) that do not account for breaks. The break dates in Lee and Strazicich test are endogenously determined and can be explained using two models i.e., model A and model C. These models are based on alternative assumptions about structural breaks, for example model A allows for two shifts in the intercept and model C includes two shifts in the intercept and trend as follows:

$$\begin{aligned}&\text{Model A}: Z_t = [1, t, D_{1t}, D_{2t}]' \\ &\left(D_{jt} = 1 \text{ for } t \geq T_{Bj} + 1, j = 1, 2, \text{ and } 0 \text{ otherwise}\right)\end{aligned} \tag{7}$$

$$\begin{aligned}&\text{Model C}: Z_t = [1, t, D_{1t}, D_{2t}, DT_{1t}, DT_{2t}]' \\ &\left(DT_{jt} = t - T_{Bj} \text{ for } t \geq T_{Bj} + 1, j = 1, 2, \text{ and } 0 \text{ otherwise}\right).\end{aligned} \tag{8}$$

T_{Bj} denotes the break date. Eqs. (9) and (10) state the null and alternative hypothesis of the two models, respectively.

$$\begin{aligned}&H_0: y_t = \mu_0 + d_1 B_{1t} + d_2 B_{2t} + y_{t-1} + \nu_{1t}; \\ &H_1: y_t = \mu_1 + \gamma t + d_1 D_{1t} + d_2 D_{2t} + \nu_{2t};\end{aligned} \tag{9}$$

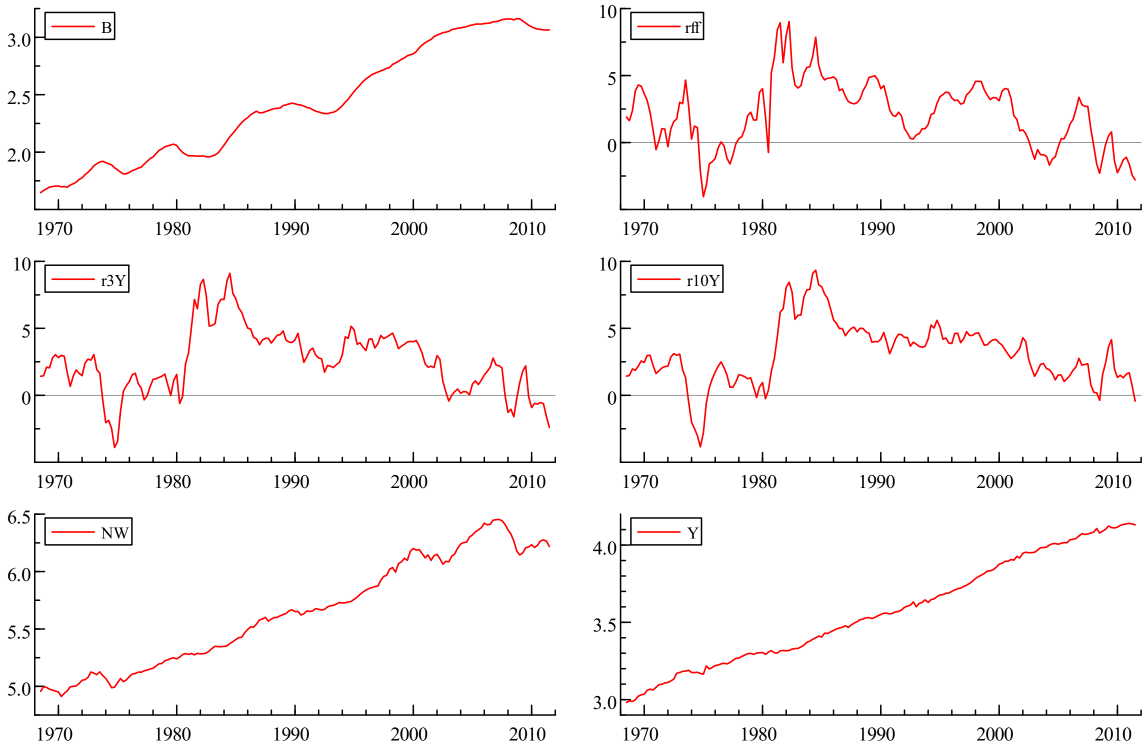
$$\begin{aligned}&H_0: y_t = \mu_0 + d_1 B_{1t} + d_2 B_{2t} + d_3 D_{1t} + d_4 D_{2t} + y_{t-1} + \nu_{1t}; \\ &H_1: y_t = \mu_1 + \gamma t + d_1 D_{1t} + d_2 D_{2t} + d_3 DT_{1t} + d_4 DT_{2t} + \nu_{2t}.\end{aligned} \tag{10}$$

ν_{1t} and ν_{2t} are stationary error terms and $B_{jt} = 1$ for $t = T_{Bj} + 1$, $j = 1, 2,$ and 0 otherwise. To attain the LM test statistic, the following regression is estimated:

$$\Delta y_t = \delta' \Delta Z_t + \phi \overline{S}_{t-1} + \mu_t \tag{11}$$

where $\overline{S}_t = y_t - \overline{\psi}_x - Z_t \overline{\delta}$, $t = 2, \ldots, T$; the regression of Δy_t provides estimates of $\overline{\delta}$; $\overline{\psi}_x = y_1 - Z_1 \overline{\delta}$ and the first observations of y_t and Z_t are y_1 and Z_1, respectively. The LM statistic tests for the unit root null hypothesis against otherwise. The optimal lag lengths (from a maximum of 8 lags) are selected using the t-sig method of Ng and Perron (1995).

The results are reported in Table 2. The test statistics of the LM unit root tests for all variables do not exceed the critical values in absolute terms and therefore the unit root null cannot be rejected at the 5% level. For the first differences of these variables the unit root null is rejected at the 5% level. In majority of the cases, the t-statistics corresponding to the break dates are statistically significant at the conventional levels (not reported for brevity). The break dates are consistent with the timings of macroeconomic events that were experienced by the USA economy, for instance, the collapse of the Bretton Woods system and first oil crisis in 1973, stock market crash during the 1973–74, the second oil crisis in 1979, deregulation policies were employed during the period 1974–1992, bubble in stock valuations and recession in the early 2000s and global financial crisis in the late 2000s. Many of these events (break dates) change after differencing the series, and this could be a signal that break dates detected in reality are not very robust.



Notes: All variables are expressed in real terms. B = Total consumer credit; NW = Net wealth ; Y = Disposable income; rff = Fed funds rate; $r3Y$ = 3-year constant maturity rate; $r10Y$ = 10-year constant maturity rate. B, Y, and NW are in natural log form.

Fig. 1. Plot of the series.

Table 2
Two-break minimum *LM* unit root test 1968:Q1–2011:Q3.

Variables	Level				First difference			
	Model A		Model C		Model A		Model C	
	Test statistic	Break dates	Test statistic	Break dates	Test statistic	Break dates	Test statistic	Break dates
B	−2.906 [4]	1985Q1; 1994Q1	−4.738 [1]	1990Q4; 2001Q3	−5.169 [6]	1978Q1; 1994Q3	−5.980 [6]	1989Q1; 1994Q2
Y	−2.333 [4]	1975Q2; 2001Q4	−3.886 [4]	1980Q2; 1999Q2	−8.431 [1]	1983Q1; 1992Q3	−17.447 [0]	1974Q3; 2002Q3
NW	−2.610 [3]	1975Q3; 2001Q4	−4.667 [4]	1975Q2; 2004Q3	−10.313 [0]	1974Q3; 2007Q3	−11.355 [0]	1974Q4; 2007Q3
r^{ff}	−3.579 [6]	1980Q3; 2007Q3	−4.676 [2]	1974Q1; 1979Q4	−6.466 [6]	1977Q1; 1992Q1	−7.632 [6]	1979Q4; 2004Q4
r^{3Y}	−3.155 [3]	1983Q2; 2007Q3	−4.641 [6]	1973Q3; 1981Q1	−7.061 [3]	1985Q3; 2006Q1	−9.044 [6]	1980Q1; 1984Q4
r^{10Y}	−3.296 [1]	1981Q2; 2007Q3	−5.001 [7]	1977Q4; 1982Q3	−7.171 [3]	1981Q2; 1990Q1	−8.648 [3]	1981Q4; 1986Q4

Notes: The 5% critical values for models A and C are −3.842 and −5.286, respectively. The number in square brackets indicates the optimal number of lagged first-differenced terms included in the unit root test to correct for serial correlation. Critical values are taken from Lee and Strazicich (2003, 2004). Those of Kumar and Webber (2013) contain more details on this test. RATS 7.2 was used to perform this test.

3.3. Breitung rank tests

Breitung's (2001) technique tests the null hypothesis of no cointegration against the alternative hypothesis of cointegration of either linear or non-linear form. Consider the following multivariate rank statistic to test for cointegration among $k+1$ variables y_t, x_{1t}, …, x_{kt}:[5]

$$\Xi_T^* = \frac{T^{-3}\sum_{t=1}^{T}\left(\tilde{u}_t^R\right)^2}{\hat{\sigma}^2_{\Delta\tilde{u}}} \tag{12}$$

where $\tilde{u}_t^R = R(y_t) - \sum_{j=1}^{k}\tilde{b}_j R\left(x_{jt}\right)$, in which $\tilde{b}_1, \ldots, \tilde{b}_k$ are the least squares estimated from a regression of $R(y_t)$ on $R(x_{1t})$, …, $R(x_{kt})$ and $\tilde{u}_t^R$ are the estimated residuals. $\hat{\sigma}^2_{\Delta\tilde{u}}$ is included to avoid possible correlation among the series. The null of linear cointegration between the variables are rejected if the test statistics are smaller than their respective critical values; the critical values are available in Breitung (2001).

Breitung (2001) also suggested a score test statistic to identify the linearity nature of the cointegrating relationship. The score test statistic is given by TR^2 from the following least squares regression:

$$\tilde{u}_t = c_0 + c_1 x_t + c_2 R(x_t) + e_t \tag{13}$$

where T is the sample size and R^2 is the estimate of the determination in Eq. (13). The $\tilde{u}_t$ are the errors possibly corrected for serial correlation and endogeneity using for example the Stock and Watson's (1993) dynamic ordinary least squares method (DOLS).

We tested for the non-linear cointegration for total credit using different interest rates, viz. The fed funds rate, 3-year constant maturity rate and 10-year constant maturity rate. Each model includes real

[5] A good description of this technique can be found in Haug and Basher (2011) and Liew et al. (2009).

Table 3
Breitung's rank tests of non-linear cointegration 1968:Q1–2011:Q3.

Specification	Total credit model	
	Ξ_T^*	TR^2
$B = f(Y, NW, r^{ff})$	0.004	0.403
$B = f(Y, NW, r^{3Y})$	0.001	1.164
$B = f(Y, NW, r^{10Y})$	0.001	0.092

Notes: The 5% critical values for Ξ_T^* and TR^2 test statistics are 0.019 and 5.990, respectively. The null hypothesis of no cointegration is rejected for a test statistic value smaller than the critical value. For TR^2 the null hypothesis of linear relationship exists against the alternative of existence of non-linear relationship. Reject the null hypothesis if computed TR^2 value exceeds the critical value. The non-linear-score test follows a χ^2 distribution with one degree of freedom.

disposable income and real net wealth. Table 3 presents the results. The results strongly reveal that we can reject the null of no cointegration in favor of cointegration of either linear or non-linear form in all models at the 5% level. To this end, all the three measures of real interest rate seem to perform well. Having determined the credit cointegrating relationships, in the next stage we investigate the linearity nature of the existing cointegrations. In this respect, the non-linear score test statistics are statistically significant at the 5% level and this clearly indicates that the existing cointegrating relationships are linear in nature. To this end, the existing literature on consumer credit demand[6] is not flawed; indeed most studies have utilized the linear cointegration techniques.

3.4. Threshold cointegration

The threshold cointegration approach is introduced by Enders and Siklos (2001), who developed the unit root test of Enders and Granger (1998) taking asymmetry into account to test for threshold cointegration. Let $\{x_{it}\}_1^T$ denote observable random variables integrated of order one i.e. $I(1)$. The long-run equilibrium relationship is given by:

$$x_{1t} = \beta_0 + \beta_2 x_{2t} + \ldots + \beta_n x_{nt} + \mu_t \tag{14}$$

where β_0 is the constant, $\beta_2, \ldots, \beta_n$ are the estimated parameters, and μ_t is the disturbance term. The existence of the long-run relationship requires μ_t to be stationary. The stationarity of μ_t has to be investigated in the second step, after having estimated the long-run relationship using the OLS method.[7] The second step procedure is given by:

$$\Delta\mu_t = \rho\mu_{t-1} + \varepsilon_t \tag{15}$$

where ε_t is the white noise disturbance. If $-2 < \rho < 0$, the long-run equilibrium (Eq. (14)) with symmetric adjustment is accepted. However, this procedure is misspecified if the adjustment process is asymmetric and therefore, Enders and Siklos (2001) proposed the following asymmetric adjustment model so called the threshold autoregressive (TAR) model:

$$\Delta\mu_t = I_t\rho_1\mu_{t-1} + (1-I_t)\rho_2\mu_{t-1} + \varepsilon_t \tag{16}$$

$$I_t = \begin{cases} 1 & if \quad \mu_{t-1} \geq \tau \\ 0 & if \quad \mu_{t-1} < \tau \end{cases} \tag{17}$$

where I_t is the Heaviside indicator and τ is the value of the threshold. Since the exact nature of non-linearity is not known, it is also possible to allow the adjustment to depend on the change in μ_{t-1} (i.e. $\Delta\mu_{t-1}$) instead of the level of μ_{t-1}. In this case, the Heaviside indicator in Eq. (17) becomes:

$$I_t = \begin{cases} 1 & \text{if} \quad \Delta\mu_{t-1} \geq \tau \\ 0 & \text{if} \quad \Delta\mu_{t-1} < \tau \end{cases}. \tag{18}$$

This variant model is used by Enders and Granger (1998) and Caner and Hansen (1998) and allows a variable to display differing amounts of autoregressive decay depending on whether it is increasing or decreasing. This model is so called the momentum-threshold autoregressive (M-TAR) model. To satisfy the necessary and sufficient conditions of the stationarity of μ_t, $\rho_1 < 0$, $\rho_2 < 0$, $(1+\rho_1)(1+\rho_2) < 1$ is required.

The threshold value τ, which is unknown, is estimated according to Chan's (1993) method as suggested by Enders and Siklos (2001). Moreover, Enders and Siklos (2001) have proposed tests when τ is known ($\tau = 0$). When the adjustment process (Eq. (16)) is serially correlated, Eq. (16) is rewritten as:

$$\Delta\mu_t = I_t\rho_1\mu_{t-1} + (1-I_t)\rho_2\mu_{t-1} + \sum_{i=1}^{p}\gamma_i\Delta\mu_{t-p} + \varepsilon_t. \tag{19}$$

To test for threshold cointegration, Enders and Siklos (2001) proposed the test statistic Φ. The Φ statistic is computed using an F-statistic which tests for the null hypothesis $\rho_1 = \rho_2 = 0$. The F-statistic for the null hypothesis $\rho_1 = \rho_2 = 0$ using the TAR specification of Eq. (17) and M-TAR specification of Eq. (18) are called Φ_μ and Φ_μ^*, respectively. The critical values to test the null hypothesis are tabulated in Enders and Siklos (2001) and successively re-tabulated by Wane et al. (2004) for more specific cases (that is, the case of more than two variables). If the null hypothesis of no cointegration is rejected, the null hypothesis $\rho_1 = \rho_2$ can be tested with a standard F-statistic. The equilibrium relationship with symmetric adjustment is accepted when the null hypothesis with no cointegration is rejected and the null hypothesis $\rho_1 = \rho_2$ is not rejected. In this case, the Engle and Granger model characterized by symmetric adjustment is supported.

The estimated long-run relationships (obtained with OLS method) are:

$$B_t = -\underset{(47.6)}{3.039} + \underset{(7.8)}{0.720}\,Y_t + \underset{(7.8)}{0.513}NW_{t-1} - \underset{(1.5)}{0.003}\,r_t^{ff} \tag{21}$$

$$B_t = -\underset{(48.9)}{3.041} + \underset{(8.1)}{0.687}\,Y_t + \underset{(8.8)}{0.536}NW_{t-1} - \underset{(3.4)}{0.006}\,r_t^{3Y} \tag{22}$$

$$B_t = -\underset{(49.4)}{3.041} + \underset{(9.1)}{0.731}\,Y_t + \underset{(8.8)}{0.510}NW_{t-1} - \underset{(3.8)}{0.007}\,r_t^{10Y} \tag{23}$$

where t-statistics are in parentheses. In correspondence of these long-run estimates, Table 4 reports the results of threshold cointegration tests. The results show that only in the version with the fed funds rate, we can reject the null hypothesis of no cointegration for both TAR-C and M-TAR-C models at the 10% level of significance. In the version with r^{3Y} the null hypothesis of no cointegration is rejected, but the ρ_2 coefficient is not statistically significant at the conventional levels.

With these results, we infer that TAR-C and M-TAR-C models are appropriate with the fed funds rate variable, although it is statistically significant at only 10% level. This implies the existence of slight threshold cointegration between credit and fed funds rate and, consequently, an asymmetric effect of Fed policy on consumer credit dynamics. The point estimates of ρ_1 and ρ_2 in the fed funds rate version suggest faster convergence for positive than for negative discrepancies from long-run equilibrium once a certain minimum deviation is exceeded. These results seem to indicate that divergence from long-run equilibrium resulting from increases in the federal funds rate with respect to consumer credit (and such that $\mu_{t-1} \geq 0.025$ and $\Delta\mu_{t-1} \geq 0.001$ in TAR-C and M-TAR-C, respectively) are eliminated relatively quickly, whereas opposite changes (i.e. decreases in fed funds) displays more

[6] Particularly, the studies that have utilized the USA data.

[7] Enders and Siklos (2001) use this method in their procedure. For maintaining coherence with their procedure, we also use the OLS method. The long-run estimation results with OLS is very similar, in our case, to the results obtained via FMOLS and CCR.

Table 4
Threshold adjustment cointegration tests, 1968Q1–2011Q3.

Model	Lag	ρ_1	ρ_2	Φ or Φ^*	$\rho_1 = \rho_2$	BG(1,4)
r^{ff} version						
TAR	2	−0.114 (0.03)***	−0.070 (0.03)**	7.689	0.966	[0.28, 0.22]
M-TAR	2	−0.113 (0.03)***	−0.065 (0.03)**	7.781	1.136	[0.31, 0.07]
TAR-C ($\tau = 0.02507$)	2	−0.186 (0.05)***	−0.055 (0.02)**	8.227*	5.039**	[0.32, 0.22]
M-TAR-C ($\tau = 0.00114$)	2	−0.134 (0.03)***	−0.048 (0.03)#	9.178*	3.747*	[0.36, 0.08]
r^{3y} version						
TAR	2	−0.110 (0.03)***	−0.076 (0.03)**	7.629	0.573	[0.40, 0.52]
M-TAR	2	−0.150 (0.03)***	−0.040 (0.03)	10.609**	6.054**	[0.52, 0.32]
TAR-C ($\tau = 0.02881$)	2	−0.196 (0.06)***	−0.053 (0.02)**	7.968	5.232**	[0.73, 029]
M-TAR-C ($\tau = 7.053e - 05$)	2	−0.150 (0.03)***	−0.040 (0.03)	10.610**	6.056**	[0.52, 0.32]
r^{10y} version						
TAR	2	−0.102 (0.03)***	−0.077 (0.03)**	7.214	0.304	[0.37, 0.49]
M-TAR	2	−0.132 (0.03)***	−0.049 (0.03)	8.911	3.434	[0.39, 0.50]
TAR-C ($\tau = 0.00781$)	2	−0.118 (0.04)***	−0.070 (0.03)**	7.323	1.051	[0.43, 0.45]
M-TAR-C ($\tau = 2.184e - 04$)	2	−0.133 (0.03)***	−0.049 (0.03)#	8.910	3.440*	[0.39, 0.50]

Notes: Standard errors are below the coefficients in the parentheses and p-values are in square brackets. ***, **, *, # signify significance at 1%, 5%, 10%, 15% levels, respectively. τ is the threshold level endogenously determined according to Chan's (1993) method. BG(p) = Bresuch–Godfrey test for serial correlation of order p. Φ is the F-statistic for the null hypothesis of no threshold cointegration; critical values for Φ_μ (TAR) are 8.17, 9.41, and 11.97 at 10%, 5%, and 1% significance levels, respectively; critical values for Φ_μ^* (M-TAR) are 9.01, 10.28, and 12.99 at 10%, 5%, and 1% significance levels, respectively. The critical values for Φ_μ and Φ_μ^* are obtained from Wane et al. (2004). $\rho_1 = \rho_2$ is the F-statistic that the two coefficients are equal.

persistence. However, these findings are valid only at the 10% statistical significance level.

3.5. Alternative estimates

The existence of linear cointegrating relationships implies that the cointegrating vectors can be estimated using conventional linear cointegration techniques. We utilized three techniques viz. canonical cointegrating regression (CCR), general to specific (GETS) and fully modified ordinary least squares (FMOLS) to estimate the cointegrating vectors of models in Table 4. We used three techniques to check robustness of the results. These techniques are classified as single-equation estimators and they deal with the problem of second-order asymptotic bias arising from serial correlation and endogeneity and are asymptotically equivalent and efficient. Park (1992) proposed the CCR technique which is quite similar to Phillips and Hansen (1990) FMOLS, and as efficient as methods based on system maximum likelihood estimation. While the CCR technique selects a canonical regression among the

Table 5
Alternative estimates of total consumer credit demand 1968:Q1–2011:Q3.

Variables	CCR			GETS			FMOLS		
Intercept	−3.095 (0.15)***	−3.070 (0.13)***	−3.063 (0.12)***	−2.929 (0.12)***	−2.885 (0.13)***	−2.872 (0.14)***	−3.097 (0.15)***	−3.072 (0.13)***	−3.065 (0.12)***
$\alpha_1 Y_t$	0.556 (0.22)**	0.596 (0.17)***	0.655 (0.16)***	0.690 (0.18)***	0.912 (0.19)***	0.947 (0.20)***	0.555 (0.22)**	0.594 (0.18)***	0.654 (0.16)***
$\alpha_2 NW_t$	0.628 (0.16)***	0.599 (0.12)***	0.561 (0.11)***	0.511 (0.13)***	0.360 (0.14)***	0.335 (0.14)**	0.629 (0.16)***	0.601 (0.13)***	0.562 (0.11)***
$\alpha_3 r_t^{ff}$	−0.005 (0.00)	–	–	−0.009 (0.00)**	–	–	−0.005 (0.00)	–	–
$\alpha_3 r_t^{3Y}$	–	−0.007 (0.00)*	–	–	−0.001 (0.00)	–	–	−0.007 (0.00)*	–
$\alpha_3 r_t^{10Y}$	–	–	−0.007 (0.00)*	–	–	0.002 (0.01)	–	–	−0.007 (0.00)*
λ	−0.050***	−0.045***	−0.043***	−0.058***	−0.054***	−0.051***	−0.050***	−0.045***	−0.043***
$\bar{R}^2$	0.820	0.806	0.804	0.818	0.803	0.802	0.820	0.805	0.804
EG test	−4.535**	−4.337**	−4.280**	−3.995*	−4.223**	−4.243**	−4.536**	−4.339**	−4.278**
LM(1) test (p-value)	0.320	0.150	0.100	0.654	0.372	0.250	0.319	0.149	0.099
LM(2) test (p-value)	0.250	0.104	0.054	0.369	0.191	0.123	0.249	0.102	0.053
LM(4) test (p-value)	0.468	0.321	0.160	0.645	0.317	0.212	0.467	0.318	0.159
JB test (p-value)	0.083	0.088	0.101	0.251	0.131	0.084	0.082	0.089	0.101
BPG test (p-value)	0.465	0.724	0.783	0.623	0.644	0.677	0.467	0.722	0.783

Notes: All variables (excluding interest rate) are expressed in natural log. The standard errors are reported in ( ) brackets. *, **, *** denote significance at 10%, 5%, and 1%, respectively. BPG, Breusch–Pagan–Godfrey heteroskedasticiy test; JB, Jarque–Bera normality test, LM, Breusch–Godfrey serial correlation LM test. FMOLS uses Newey–West automatic bandwidth selection in computing the long-run variance matrix. The p-values are reported for the diagnostic tests. In the ECM formulations we used the following spike dummies: Du80q2 (US recession), Du87q1 (Federal Reserve (Fed) started to react to variations in inflation rates and unemployment; see Curtis (2005) on this point), Du89q1 (slowdown of economy and consumption due to a restrictive monetary policy enacted by the Fed), and Du98q2 (low inflation). EViews 7 was used to estimate the above equations.

Table 6
Quandt–Andrews structural break tests, 1968:Q1–2011:Q3.

Statistics	$B_t = \alpha_0 + \alpha_1 Y_t + \alpha_2 NW_{t-1} + \alpha_3 r_t^{3Y}$			$B_t = \alpha_0 + \alpha_1 Y_t + \alpha_2 NW_{t-1} + \alpha_3 r_t^{10Y}$		
	Value	Break	Prob.	Value	Break	Prob.
Max LR F-stat	2.171	1976Q3	1.00	2.321	1984Q3	1.00
Max Wald F-stat	21.710	1976Q3	0.21	23.537	–	0.12
Exp LR F-stat	0.642	–	1.00	0.796	–	1.00
Exp Wald F-stat	8.220	–	0.17	10.252	–	0.07
Ave LR F-stat	1.228	–	1.00	1.548	–	1.00
Ave Wald F-stat	12.282	–	0.20	17.029	–	0.05

Note: Probabilities calculated using Hansen's (1997) method. EViews 7.0 was used to perform this test.

class of models representing the same cointegrating relationship, the FMOLS modifies series and estimates directly to eliminate the existing nuisance parameters. Operationally, the CCR method concentrates on the data transformations, but FMOLS use the transformations of both the data and estimates. The GETS technique was proposed by the London School of Economics Professor David Hendry and it utilizes the general dynamic specification similar to the autoregressive distributed lag model. The variable deletion tests are applied to attain the parsimonious estimated model; for more details on the GETS technique, see Rao et al. (2010).

The CCR, GETS and FMOLS cointegrating equations are reported in Table 5. The three estimation techniques provided consistent estimates and reveal that real disposable income and wealth are positively linked with total consumer credit over this time period. The estimates of real disposable income and wealth are statistically significant at the 5% level in all models. The coefficient of federal funds rate is statistically insignificant at the conventional levels in FMOLS and CCR methods. In contrast, the federal funds rate is highly significant in GETS specification but residual tests of cointegration (EG) rejects the null of no cointegration only at the 10% significance level. The estimates of 3-year constant maturity rate and 10-year constant maturity rate are statistically significant at the conventional levels in FMOLS and CCR; in addition EG test rejects the null of no cointegration at the 5% level. The factor loading parameter is very low in all formulations and this suggests that error correction mechanism is very slow, i.e. consumer credit reverts towards the equilibrium level very slowly. The diagnostic test results are satisfactory in all cases. On the basis of these results, we argue that the best formulations are FMOLS and CCR estimations with medium and long-term interest rates.

3.6. Structural breaks and stability

We investigated the stability of our estimated equations in Table 5. In doing so, we subjected the FMOLS estimates to Quandt (1960) and Andrews (1993) structural breakpoint tests. Using insights from Quandt (1960), Andrews (1993) modified the Chow test to allow for endogenous breakpoints in the sample for an estimated model. This test is performed at every observation over the interval $[\xi T, (1-\xi)T]$ and computes the supremum of the F_k statistics ($\sup F = \sup_{k \in [\xi T,(1-\xi)T]} F_k$) where ξ is a trimming parameter. Andrews and Ploberger (1994) developed two additional test statistics i.e. the average (ave F) and the exponential (exp F). The null hypothesis of no break is rejected if these test statistics are large, however Hansen (1997) derives an algorithm to compute approximate asymptotic p-values of these tests.

Table 6 displays the Quandt–Andrews test results.[8] The results show that majority of the test statistics do not reject the null of no structural breaks at the 1% level. The detected break dates, although statistically insignificant, are 1976Q3 (in r^{3Y}) and 1984Q3 (in r^{10Y}). The two potential break dates are not totally unexpected. The USA economy during the 1970s and 1980s experienced some major events. For example: oil crisis in the 1970s with the consequent huge increase in inflation and wages; the end of Vietnam War; the end of Gold Standard system; and the recession of 1974–1975. All these events (with some delays) may have potentially affected the consumer credit dynamics. The 1984Q3 break may refer to the beginning of the Great Moderation era (Kim and Nelson, 1999; McConnell and Perez-Quiros, 2000), a period of decreasing business cycle fluctuations. During this period, the standard deviation of quarterly real GDP declined by half, and the standard deviation of inflation declined by two-thirds (Bernanke, 2004). However, based on Quandt–Andrews test results, the two potential breaks detected are not statistically significant, and we infer that our estimated equations are stable and robust.

4. Conclusions

This paper has investigated the determinants of consumer credit — defined as the sum of revolving and non-revolving credit — for the USA for the period 1968:Q1 to 2011:Q3. Our specification and approach are consistent with the Life Cycle Hypothesis (LCH) of Modigliani and Brumberg (1955). The endogenous two break minimum LM unit root tests developed by Lee and Strazicich (2003) are employed to ascertain the time series properties of the variables. The classical linear cointegrating techniques (canonical cointegrating regression, general to specific and fully modified ordinary least squares), Breitung's (2001) non-parametric rank tests and Enders and Siklos (2001) threshold equilibrium adjustment tests are applied to investigate the consumer credit relationships.

The two break minimum LM unit root tests revealed that the level variables are non-stationary and provided break dates that are consistent with the timings of macroeconomic events that were experienced by the USA economy. Breitung's cointegration tests revealed that we can reject the null of no cointegration in favor of cointegration of either linear or non-linear form in all models at the 5% level. The non-linear score test statistics indicate that the existing cointegrating relationships are linear in nature. Application of the Enders and Siklos's threshold tests showed that non-linearity is present slightly only for the version with short-term interest rate (federal funds rate). In the case of consumer credit models with medium (3 years) and long (10 years) term interest rates, we find the presence of a linear and symmetric cointegrating relationship. Our cointegrating estimates based on CCR, GETS and FMOLS techniques revealed that consumer credit responds more significantly to the medium and long-term interest rates than the short-term interest rate. Since consumer credit outstanding is composed of revolving and non-revolving credit (which reflects most consumer short- and intermediate-term credit, excluding loans secured by real estate) this finding is not so expected. In all cases, the estimates of real wealth and real disposable income are statistically significant at the conventional

[8] We report the results for FMOLS specification only. CCR specification gives very similar results and these are available upon request.

levels. Based on the Quandt (1960) and Andrews (1993) structural break tests, we infer that our estimated equations are stable and robust.

Our finding that consumer credit is linked to real wealth, real disposable income, and real interest rates via cointegrating relations indicates that the observed long-run movements in consumer credit as movements up and down a well-defined demand curve reflects optimizing and forward-looking behavior; that finding suggests that myopic or otherwise irrational behavior might not be as important as commonly believed. To this end, the popular belief that American consumers 'over-borrowed' during the 1990s and 2000s and that this behavior set the stage for the crisis and severe recession that followed, beginning in 2007–08 may not be supported. While we do find some evidence of structural breaks in the data and cointegrating relation, these did not create any instability in the consumer credit demand. Moreover, our findings imply that the medium and long-term interest rates are significant determinants of consumer credit, and surprisingly, the consumer credit responds less significantly to the federal funds rate. During the current global downturn, the Fed and registered commercial banks focus on 3-year and 10 constant maturity rates could stimulate the demand for consumer credit. This finding is valuable to the Federal Reserve Bank to pursue an effective monetary policy.

Appendix A

Data appendix

Total consumer credit outstanding (revolving credit outstanding plus non-revolving credit outstanding), federal funds rate, 3-year and 10-year Treasury constant maturity rates, personal consumption expenditure (PCE) price index, and consumer price index (CPI) are obtained from Federal Reserve Economic Data (FRED).

Consumer credit outstanding is composed of revolving and non-revolving credit, which reflect short- and intermediate-term consumer credit, excluding loans secured by real estate. Revolving credit is composed mostly of credit card loans (roughly 95%); the remainders are "lines of credit" extensions, which are used for checking account overdraft facilities. Non-revolving credit includes automobile loans and all other loans not included in revolving credit, such as loans for mobile homes, education, boats, trailers, or vacations. These loans may be secured or unsecured. Disposable income is constructed from National Income and Product Account (NIPA) as did by Luvigson and Steindel (1999). Total net wealth (i.e. total assets (financial and non-financial assets) minus total liabilities) is obtained by flow-of-funds accounts of the U.S. Bureau of Economic Analysis (BEA). All variables are deflated by PCE chained type price index.

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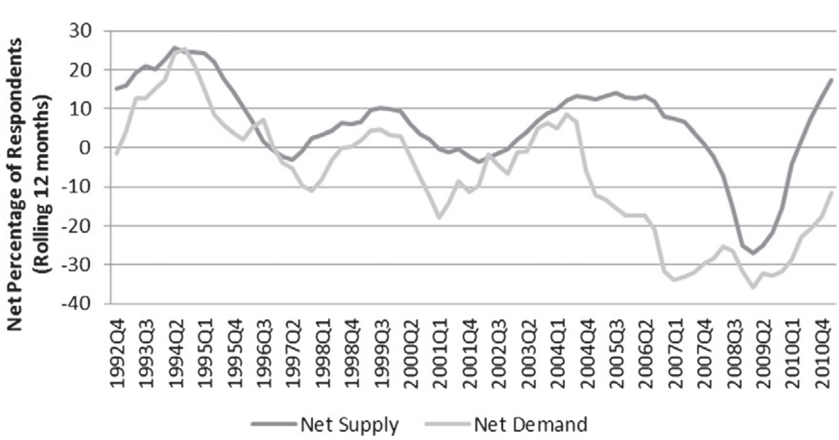

Source: Senior Loan Officer Opinion Survey on Bank Lending Practices (The Federal Reserve Board).

Fig. A. Net supply and demand for consumer credit.

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