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IS THE US CONSUMER CREDIT ASYMMETRIC?

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Abstract

We investigate the existence of any asymmetric effects in the US consumer credit. In doing so, we utilize the asymmetric cointegration methods proposed by Breitung (2001, 2002) and Enders and Siklos (2001). Furthermore, we tend to explore two additional dimensions in this literature. First, whether asymmetries (if any) are persistent over-time in the credit demand model. Second, whether the Great recession contributed to any asymmetric impacts on the credit demand. Our results revealed that the long-run relationship of consumer credit (credit, income, wealth and interest rate on personal loans) is asymmetric. While it is difficult to identify the direct sources of this asymmetric result, our intuition is that it is linked to the structural breaks in the rate of personal loans encountered in the early 1980s. Moreover, we find no strong evidence that much of the asymmetric impacts in consumer credit were experienced in the Great recession. Neither have we attained evidence that asymmetric impacts on credit demand are persistent over-time.

I Introduction

> During the recent financial crisis, the Federal Reserve and other policymakers throughout the government took unprecedented actions to mitigate the fallout from severely distressed market conditions and support the flow of credit to consumers and businesses. Nonetheless, the level of credit outstanding for households has been very slow to rebound and remains lower than it was at the onset of the crisis. The reasons for the slow rebound are, without a doubt, complex and multidimensional. Still, it is worthwhile to examine the data and try to understand why credit growth is not more robust.
>
> Governor Elizabeth A. Duke of Federal Reserve System of the US[1]

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[1] Speech can be found at: http://www.federalreserve.gov/newsevents/speech/duke20101202a.htm

Central banks such as the Federal Reserve (Fed) and European Central Bank (ECB) does consider the role of consumer credit in their policy decisions (see ECB, 2004; Bernanke, 2006). There is a widespread consensus among economists and policy-makers that consumer credit is strongly linked to output and inflation. There are a number of factors that consumers consider when making borrowing decisions. For the Fed to precisely measure and monitor the consumer borrowing, there is a need to identify the factors that influence consumer borrowing and decisions. Consumer spending is the largest share of GDP in the United States. and has been a key driver of economic growth the country has experienced since 1990s. Consumer borrowing has always played a strong part in stimulating consumption spending (Paradiso *et al*., 2014); in addition it also has links to consumer debt accumulation. As a proportion of personal income, consumer credit has more than doubled during the postwar period, with particularly sharp and sustained increases occurring in the 1980s (see figure 1, Ludvigson, 1999). In spite of the attention paid to movements in consumer credit in the press and the Wall Street, the determinants and effects of growth in consumer credit have not been a major focus of researchers; hence there has been limited research on the determinants of consumer credit in an economy.

There are many important strands in this literature, for example, existence of a credit channel in the transmission mechanism of the monetary policy (Angeloni *et al*., 2003), role of credit as a non-linear propagator of shocks (Balke, 2000; Gambacorta and Rossi, 2010), exploring how the dynamics of real and financial variables are affected by financial shocks using the dynamic stochastic general equilibrium (DSGE) model (Urban and Quadrini, 2012). Other studies in this literature have modelled consumer credit using the linear cointegration methods (Hartropp, 1992; De Nederlandsche Bank, 2000; Calza *et al*., 2001, 2003; Hofmann, 2001; Schadler *et al*., 2004). None of the above studies have considered modelling credit demand using alternative measures of interest rate such as rate on personal loans (24 months), short-term rate (3 years), medium-term rate (10 years), short-term consumer loan rate and the federal funds rate. With the exception of Balke (2000) and Gambacorta and Rossi (2010), none have explored the asymmetric long-run adjustments in the demand-determined consumer credit model.

The objective of this article is to investigate the existence of any asymmetric effects in the US consumer credit. In doing so, we utilize the asymmetric cointegration methods proposed by Breitung (2001, 2002) and Enders and Siklos (2001). Existing studies on credit demand asymmetries are scant and we aim to partly fill this gap by investigating the existence of asymmetric effects on credit demand. Furthermore, we tend to explore two additional dimensions in this literature. First, whether the asymmetries (if any) are persistent over time in the credit demand model. Second, whether the Great recession contributed to any asymmetric impacts on the credit demand. The testing of non-linear relationships has gained an increasing attention in the time series literature and it is generally believed that some economic variables may be highly non-linear; see Fan *et al*. (2004) and Paradiso *et al*. (2014). Generally speaking, the sources of non-linearity may be due to market frictions, heterogeneous agents and official interventions. Our results revealed that the long-run relationship of consumer credit (credit, income, wealth and rate on personal loans) is asymmetric. While it is difficult to identify the direct sources of this asymmetry, our intuition is that it is linked to the structural breaks in the rate of personal loans encountered in the early 1980s. Moreover, we find no strong evidence that much of the asymmetric impacts in consumer credit were experienced in the Great recession. Neither have we attained evidence that asymmetric impacts on credit demand are persistent over-time.

The outline of this paper is as follows: Section II discusses the developments in the literature and presents the model specification. Empirical results are discussed in Section III. Finally, Section IV concludes.

II Recent Literature and Model Specification

Credit has been the focus of a large body of literature, particularly after the seminal paper of Stiglitz and Weiss (1981). One of the reasons for this considerable attention is the implication that credit has for the transmission mechanism of monetary policy. In what follows, we review some key studies on this topic to gain insights about the behaviour of credit market as well as the important findings that executed through the application of symmetric and asymmetric models and methods.

Credit market

Much of the earlier literature has focused on the link between credit aggregates and economic activity; see Gertler (1988) for a review of earlier studies. In the business cycle models, credit plays a vital role by propagating and strengthening productivity and monetary policy shocks. In the standard real business cycle model and the standard Keynesian textbook IS-LM model, credit market conditions do not exhibit any significant macroeconomic outcomes; see Hofmann (2001) for more details. Due to this, the credit market is characterized as frictionless. In an early study, Blinder (1987) argued that the performance of the economy depends on whether or not the credit constraint is binding.[2] Since then, several studies have utilized general equilibrium models and incorporated financial market imperfections to examine the links between shocks and fluctuations in output. Bernanke and Blinder (1988) developed an analogue to the simple IS-LM model which embodied an unconventional view of the monetary transmission mechanism. In their model, the central bank policy operates through influencing the bank loans and deposits. Similar framework was used in Bernanke and Blinder (1992).

Bernanke and Gertler (1989) argued that the existence of asymmetric information in credit markets can provide borrowers a role to play in the business cycle through their impact on the cost of external finance.[3] Due to the presence of information asymmetries, agents are borrowing constrained and

[2] See Calza and Sousa (2005) for details.

[3] See Christensen and Dib (2008) for a review.

their capacity to borrow relies upon their net worth. Bernanke *et al.* (1996) also reinforced this mechanism, i.e. the financial accelerator hypothesis. Bernanke *et al.* (1999), Kiyotaki and Moore (1997), Carlstrom and Fuerst (1997) and Azariadis and Smith (1998), among others, argued that financial frictions may intensify the stickiness of oscillations in income. In this regard, Azariadis and Smith (1998) proposed an overlapping generations model in which the economy transits into a traditional Walrasian equilibrium system or in a binding credit constraints type system.

Using a threshold framework, McCallum (1991) examined the response of income to changes in monetary policy in the United States. They found that the impact of growth in money on income is high when critical thresholds do not exceed credit constraints.[4] Following similar line of work, Galbraith (1996) and Balke (2000) allowed for endogenous determination of critical values in the threshold model. In such models, the credit conditions are viewed as expansionary or contractionary. The examination of the asymmetric response to shocks is then examined in the subsequent stage. Analogous investigation was also done by Atanasova (2003) for the United Kingdom.

Gambetti and Musso (2012) examined the role played by loan supply shocks over the business cycle in the Euro Area, United Kingdom and the United States. They find that loan supply shocks have a significant impact on economic activity and credit market variables. Urban and Quadrini (2012) developed a model with debt and equity financing to explore how the dynamics of real and financial variables are affected by financial shocks. They find that financial shocks contributed significantly to the observed dynamics of real and financial variables.

Symmetric findings in the literature

In the empirical literature,[5] many studies perceived that credit is demand-determined (Bernanke and Blinder, 1988; Fase, 1995; Calza *et al.*, 2001).[6] Fase (1995) and Hofmann (2001) well discussed the crucial implications of credit demand. In this framework, banks target the loan demand and interest rates on borrowing or lending to equilibrate loan demand with banks' desired portfolio of loans. In an earlier study, Pollin (1988) examined the role of demand-side factors in determining borrowing in the United States. The regression results revealed two major demand-side influences on net borrowing, i.e. the rise in housing prices concurrent with declines in the real median incomes since the mid-1990s, and the attraction of financial investment when real borrowing costs are falling and bank yields are rising. Downes *et al.* (1997)[7] explored the extent to which the variation in private consumer debt can be

[4] See Calza and Sousa (2005) for details.

[5] Some studies did utilize survey data to examine the credit market dynamics, see Crook and Hochguertel (2005) and Crook and Crook (2001).

[6] Credit is also supply-determined. Changes in firms and households income may influence the financial institutions willingness to lend. Further, the costs of borrowing (interest rates) may also affect the financial institutions supply of credit.

[7] Their study was based on the Barbadian economy.

explained by changes in demand-side factors like income, wealth, nominal and real interest rates, inflation and government policy controls. They found that current wealth (+ve), real disposable income (+ve), inflation (−ve) and interest rates (−ve) have significant impacts on credit demand in the long-run; the expected signs are indicated in parentheses. In addition, they found that government policy controls restricts consumer borrowing in the short-run, an outcome which appears counter to central bank's policy stance.

Hartropp (1992) formulated a credit demand model to examine the flow of non-mortgage lending to consumers in the United Kingdom. His findings suggest that current income and past wealth have significant positive impacts on consumer borrowing. The nominal and real rates of interest and credit controls (past levels of income) had a significant negative (no) impact on consumer credit demand. Calza *et al.* (2001) investigated the borrowing demand for the Euro Area and attained a long-run relationship between credit, real weighted short-term and long-term interest rates and real GDP. In an analogous study, Calza *et al.* (2003) estimated a vector error correction model of euro area loans to the private sector loan deflated by the GDP deflator, real GDP, annualized quarterly inflation and a measure of the average nominal lending interest rate defined as a weighted average of retail bank lending rates to households and firms. Also see De Nederlandsche Bank (2000) for similar analyses on several EU countries including the United States. 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.

Several studies (Borio *et al.*, 1994; IMF, 2000; Hofmann, 2001) have stressed the importance of credit markets and property prices. Using cointegrating vector autoregression (VAR) model, Hofmann (2001) analysed the determinants of credit to the private non-bank sector for 16 industrialized countries. The cointegration tests suggest that the long-run development of credit cannot be explained by standard credit demand factors. However, when real property prices are incorporated into the model, then a stable long-run relationship is identified linking real credit positively to real GDP and real property prices and negatively to the real interest rate. Goodhart (1995) examined the factors influencing credit growth in the United States and United Kingdom. For the United Kingdom, change in house prices has a significant positive impact on credit growth; this result is not achieved for the United States. Considering a large sample of advanced countries, Borio *et al.* (1994) explored the links between credit-GDP ratios and aggregate asset prices. Their findings suggest that credit is statistically significant factor in driving asset prices. Analogously, Goodhart and Hofmann (2001) also attained a long-run relationship between credit and house prices for a number of countries.

Asymmetric findings in the literature

Numerous studies have explained asymmetries associated to the credit market. In their classic paper, Stiglitz and Weiss (1981) propose separated models of random credit rationing, with either adverse selection or moral hazard. Their model implies that interest rate or collateral are not appropriate means to ration credit. Using Stiglitz and Weiss model of credit rationing, Kaufman (1996) found that the credit crunch in Argentina (1995-96) resulted from an increase in the share of illiquid borrowers induced by the rise in interest rates, and increased incidence of adverse selection problems.[8] Using time series data for the United States 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. Their findings imply that the concave supply function seems to result in the temporal shifting of housing demand. As interest rates rise above the bank-optimal rate, housing demand increases in the subsequent periods. This may prevent the smoothing of housing production, as it may ultimately raise the cost of housing.

Drake and Holmes (1995, 1997) find presence of adverse selection, in the form of backward-bending credit supply curve, in the US and the UK mortgage markets, as well as in the UK market for consumer credit. Gambacorta and Rossi (2010) investigated the possible non-linearities in the response of bank lending to monetary policy shocks in the euro area. The credit market is modelled using the asymmetric vector error correction model. Their findings suggest that the effect on credit, GDP and prices of a monetary policy tightening is larger than the effect of a monetary policy easing. They support the existence of an asymmetric broad credit channel in the Euro Area. 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. Grant and Padula (2013) examined the impact of informal credit and judicial costs on the repayment behaviour for Italy. They find support for both adverse selection and moral hazard in the credit market.

Why study asymmetries in credit markets? A credit supply explanation

The market for bank debt is imperfect, for example see Kashyap and Stein (1995, 2000), Stein (1998), Kishan and Opiela (2000) and Ehrmann *et al.* (2003). The credit literature has shown that the monetary policy shocks may have asymmetric impacts on output and inflation (Gambacorta and Rossi, 2010). The theoretical perspective that supports this conjecture elaborates about the characteristics of the loan supply curve. One of the key assumptions is that the loan supply curve has some form of rigidity. For example, Stiglitz and Weiss (1981) show that an adverse selection problem leads to a backward bending supply of credit and a consequent credit rationing on the upside. Asymmetric behaviour may also be perceived even if loan supply always matches loan demand. For instance, reflecting on the 'bank lending channel', one may argue that during the times of monetary easing it is not viable for a bank to raise new loan supply due to capital regulations. In such situation, it is mandatory to maintain the proportions of bank's capital and lending. How-

[8] See Agenor and Aizenman (1998) for a review.

ever, in the case of a monetary contraction, this rigidity does not exist. To this end, the loan demand and income both show a decreasing trend. De Long and Summers (1988) provide an alternative explanation for the asymmetric effects on credit; see Gambacorta and Rossi (2010) for details.

Our paper fills an important gap in the literature by investigating the existence of asymmetric impacts on the long-run credit demand relationship in the United States. Numerous studies have examined the existence of asymmetric effects in the context of credit supply (see for example, Gambetti and Musso, 2012). Our paper tends to shed light on two other issues that received no attention yet. First, we assess whether asymmetries (if any) are persistent over-time in the credit demand model. Second, we investigate whether the Great recession contributed to any asymmetric impacts on the credit demand.

Theoretical framework and specification

The theory of consumer debt and credit constraints is linked to the life-cycle model (Modigliani, 1986) and the permanent income hypothesis (Friedman, 1957). A representative household maximizes the utility function subject to an intertemporal budget constraint[9]:

$$\max E\left[\sum_{t=0}^{T}(1+\theta)^{-t}u(c_t)\right] \tag{1}$$

$$A_{t+1} = (1+r)(A_t + y_t - c_t) \tag{2}$$

where c = consumption, y = labour income, A = household net assets, r = rate of return on assets and θ = discount rate. The standard Euler equation is attained through performing the first order condition of the problem

$$E_t u'(c_{t+1}) = \frac{(1+\theta)}{(1+r)}u'(c_t) \tag{3}$$

The implication of (3) is that consumers maximize their utility by smoothing marginal utility over the life cycle. During the times of low income, consumers borrow to smooth their path of consumption and they repay during the times of high income. Assuming perfect capital markets, consumers are able to borrow to smooth their path of consumption.

Following Hartropp (1992) and Park (1993), the empirical specification of the demand for consumer credit may appear as follows:

$$C_t = \alpha_0 + \alpha_1 Y_t + \alpha_2 W_t + \alpha_3 i_t \tag{4}$$

where C = real consumer credit, Y = real disposable income, W = real net wealth and i = measures of real interest rate. It is well known that the distribution of income affects the aggregate borrowing behaviour of consumers. Consumers, who do not have enormous current income but expects stable future income, may on average actively borrow to finance current consumption. Alternatively, consumers with high incomes borrow less and consumers

[9] See Chen and Chivakul (2008) for more details.

with unstable jobs may be quite reluctant to borrow. Park (1993), Hendricks *et al.* (1973) and Kennickell and Shack-Marquez (1992) present detailed insights about income distribution and borrowing. Hartropp (1992) argued that rising income (current and expected) is a necessary condition for extra debt. In other words, rising income may, for a particular borrower-type household, create new borrowing but it does not automatically cause new borrowing. Therefore, income is an important determinant of credit, although there is no fully deterministic link.

On the demand side, an increase in net wealth implies that an individual can afford more desired consumption and may not need to borrow. On the other hand, a direct relationship between wealth and borrowing is also possible. Hartropp (1992) argued that present and past increases in wealth may, for borrower types, result in new borrowing. This conjecture is also supported by Green and Hadjimatheou (1990). Green and Hadjimatheou (1990) find that housing wealth is an important determinant of consumption in the United Kingdom. They argued that the borrowing-wealth relationship is much more straightforward than that between borrowing and income.

Consumer borrowing decisions are influenced by interest rates (Hartropp, 1992; Park, 1993). When interest rates are high on borrowing, this implies a larger sacrifice of future income for a given level of current consumption financed by future income. Park (1993) showed that a negative relationship exists between the interest rate on 48-month new car loans and the automobile credit ratio. Moreover, interest rates also have implications on debt repayments. A high cost of carrying debt tempts borrowers to repay prevailing debt rapidly. Therefore, in addition to decelerating consumption, increases in interest rates diminish the share of consumption financed with debt and raises the repayment rate. Thus, this causes sluggish growth in consumer credit relative to consumption.

III Empirical Results

Data

We use US quarterly data over the period 1972:Q2 to 2011:Q3. The variables considered are real consumer credit (C = defined as the sum of revolving and nonrevolving credit), real disposable income (Y = total income minus taxes), real net wealth (W), real interest rate on personal loans 24 months (i), real federal funds rate (i^{ff}), real 3-year constant maturity rate (i^{3Y}) and real 10-year constant maturity rate (i^{10Y}) C, I, I^{ff}, i^{3Y} and $i^{10Y}$and are attained from the Federal Reserve Economic Data (FRED). Disposable income is constructed from National Income and Product Account (NIPA); see Ludvigson and Steindel (1999). Total net wealth is attained by flow-of-funds accounts of the US Bureau of Economic Analysis (BEA). All variables are deflated by personal consumption expenditure chained type price index. All data have been seasonally adjusted and are used in natural log form, except for the three real interest rates. Appendix A provides details on the definitions and sources of the data. The key descriptive statistics for all variables are presented in Table 1.

Table 1
Summary statistics 1972:Q2–2011:Q3

	C	Y	W	i	i^{ff}	i^{3Y}	i^{10Y}
Mean	2.485	3.624	5.693	3.371	2.121	2.580	3.220
Standard error	0.037	0.025	0.036	0.143	0.205	0.199	0.189
Median	2.395	3.588	5.655	3.420	2.266	2.543	3.611
Standard deviation	0.462	0.314	0.447	1.797	2.573	2.506	2.371
Sample variance	0.214	0.099	0.200	3.228	6.621	6.279	5.622
Kurtosis	−1.435	−1.319	−1.323	0.687	−0.239	−0.015	0.527
Skewness	0.114	0.159	0.052	0.837	0.101	0.109	−0.027
Range	1.359	1.020	1.485	1.500	13.072	13.007	13.187
Minimum	1.803	3.120	4.963	3.294	−4.043	−3.897	−3.847
Maximum	3.162	4.140	6.448	9.201	9.029	9.110	9.340

Notes: All series are in real terms. Total consumer credit, net wealth and disposable income are in natural log form.

Unit root tests

The integrated properties of the series are tested using Lee and Strazicich (2003) (LS henceforth) and Carrion-i-Silvestre *et al*. (2009) (CKP henceforth) tests. Methodological details of both tests are available in Appendix A. The results are reported in Table 2. The test statistics of LS 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 (*t*-statistics are not reported for brevity; * in Table 2 denotes statistical significance at 5% level). Unsurprisingly, we attain consistent results using the CKP test. To this end, we used their feasible point optimal statistic ($P_T^{gls}(\lambda^0)$) to derive results. The test statistics are more negative than the critical values implying that the unit root null cannot be rejected at the 5% level. Since both the tests point to non-stationarity in the series, we argue that the series are I (1). The endogenous break dates yielded by both tests are plausible.[10] They are consistent with the timings of macroeconomic events that were experienced by the United States, for instance, the second oil crisis in 1979, deregulation policies were employed during the period 1974–1992, recessions in the early 1980s, 1990s and 2000s, bubble in stock valuations in the early 2000s and recent global financial crisis.

Is the credit demand relationship asymmetric?

We first utilize Breitung's (2001) rank tests for non-linear cointegration to test the null hypothesis of no cointegration against the alternative of cointegration in either linear or non-linear type. For details about this test procedure, see

[10] In the CKP test we allowed for only two breaks to examine whether we achieve similar dates as given in LS test. Our results show that the break dates yield by LS and CKP are very consistent.

Table 2
Unit root test results, 1972:Q2–2011:Q3

	LS test								CKP test	
	Level				First difference				Level	
	Model A		Model C		Model A		Model C		$P_T^{gls}(\lambda^0)$	
Variables	Test statistic	Break dates	Test statistic	Break dates	Test statistic	Break dates	Test statistic	Break dates	Test statistic	Break dates
C	−1.823 [3]	1984Q3*; 1990Q4*	−2.003 [2]	1990Q4*; 2002Q1	−6.192 [5]	1990Q4*; 1994Q3	−5.735 [5]	1989Q4*; 1990Q4*	−5.835 (−15.12)	1990Q4; 1984Q1
Y	−3.620 [4]	1978Q2; 2003Q1	−2.184 [3]	1984Q2; 1990Q2*	−7.855 [1]	1985Q2; 1999Q4*	−14.930 [1]	1979Q1*; 1990Q2*	−21.200 (−35.60)	1984Q2; 1990Q2
W	−2.521 [5]	1979Q4*; 2000Q3*	−3.957 [4]	1991Q2*; 2001Q4*	−12.566 [0]	1992Q3; 2000Q1*	−18.120 [0]	1991Q4*; 2007Q1*	−8.028 (−12.71)	2001Q3; 2001Q4
i	−1.029 [2]	1981Q1*; 2000Q4	−3.377 [2]	1980Q4*; 1981Q3*	−5.962 [5]	1981Q2; 2000Q4	−5.432 [6]	1980Q1*; 2001Q2	−10.911 (−26.84)	1981Q2; 1981Q3
i^{ff}	−2.026 [5]	1980Q3*; 2007Q2*	−3.991 [3]	1988Q2; 1979Q4	−6.021 [5]	1991Q3*; 2007Q1	−5.901 [6]	1979Q4; 2004Q3	−4.508 (−17.43)	2006Q4; 2007Q1
i^{3Y}	−1.890 [3]	1983Q2; 2006Q4*	−2.001 [5]	1980Q3*; 1981Q1*	−7.133 [3]	1983Q2; 2006Q4*	−8.328 [5]	1980Q1*; 1982Q4	−42.657 (−47.51)	1980Q3; 2006Q4
i^{10Y}	−3.155 [1]	1981Q2*; 2007Q4*	−5.245 [5]	1992Q3; 1982Q3*	−6.904 [3]	1981Q2*; 1990Q1*	−9.020 [4]	1982Q2*; 1982Q3	−14.028 (−20.76)	1982Q2; 2007Q3

Notes: LS test: 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). Break dates that are statistically significant at the 5% level are indicated by *. Kumar and Webber (2013) contains more details on this test. RATS 7.2 was used to perform this test. CKP test: considers breaks in constant and time trend. The 5% critical values are given in parentheses. We allowed for maximum of two breaks. Gauss was used to run this test.

Table 3
Breitung Tests, 1972:Q2–2011:Q3

	Breitung (2001) test		Breitung (2002) test	
Specification	Ξ_T^*	TR^2	H0: rank ≤	Test statistic
$C = f(Y, W, i)$	0.005 (0.020)	9.720 (6.250)	0	124.280 (59.95)
			1	16.735 (32.10)
$C = f(Y, W, i^{ff})$	0.008 (0.019)	2.019 (5.990)	0	109.040 (261.70)
			1	7.911 (56.54)
$C = f(Y, W, i^{3Y})$	0.012 (0.020)	0.083 (4.895)	0	74.381 (125.87)
			1	14.025 (45.90)
$C = f(Y, W, i^{10Y})$	0.009 (0.020)	1.726 (5.990)	0	155.63 (197.24)
			1	10.620 (72.65)

Notes: Breitung (2001) test: The 95% critical values for Ξ_* and TR^2 test statistics are reported in parentheses. The null hypothesis of no cointegration is rejected for a test statistic value smaller than the critical value. For TR^2, the null hypothesis is that a 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. Breitung (2002) test: the 95% critical values are reported in parentheses. The null of no cointegration is not rejected when the test statistic is lower than the critical value.

Appendix A. We test for non-linear cointegration between real credit, real disposable income, real wealth and various measures of interest rate (rate on personal loans, Fed funds rate, 3-year constant maturity rate and 10-year constant maturity rate). The cointegration test results of Breitung (2001) are reported in Table 3. The results strongly indicate that we can reject the null of no cointegration in favour of cointegration of either linear or non-linear type in all models at the 5% level; see Ξ_T^* results. In the next stage, we examine whether the credit cointegrating relationships are linear or non-linear. To this end, the non-linear score test statistics (TR^2) does not exceed the critical values in all cases except in the model that incorporates the interest rate on personal loans. These results imply that the long-run relationship between credit, income, wealth and measures of interest rate such as Fed funds rate, 3-year constant maturity rate and 10-year constant maturity rate (rate on personal loans) are linear (nonlinear) in nature.

The existence of non-linear cointegration among the series could also be tested using the procedure in Breitung (2002); see Appendix A for details on this test. This test allows for a non-linear process where a lag structure or deterministic term need not be estimated. In addition to this, there are a number of advantages over the Bierens (1997) non-parametric procedure.[11] Our results show that the null of no cointegration cannot be rejected for credit models that include interest rates such as Fed funds rate, 3-year constant maturity rate and 10-year constant maturity rate. Interestingly, we find one cointegrating vector when credit model accommodates interest rate on personal loans. These results support our earlier findings from Breitung (2001) test. Based on these results, we infer that the long-run relationship of credit

[11] See Holmes and Panagiotidis (2009) for details.

demand is asymmetric when rate on personal loans is used as a measure of cost of borrowing.

Moreover, we run an additional robustness test to confirm whether the credit model augmented with the rate on personal loans exhibit non-linear long-run relationship. In doing so, we employ the Enders and Siklos (2001) threshold cointegration test. Details of this test are available in Appendix A. Table 4 presents the results of this test. In the first stage, credit models are estimated using the OLS method (see panel A in Table 4). In the second stage, the residuals are specified as the asymmetric Dickey–Fuller equation (see Table 4 notes), which is then used to test for threshold cointegration. This procedure tests the null hypothesis of symmetry of the coefficients against the alternative of asymmetry. In other words, it tests the null hypothesis of no cointegration against the alternative of cointegration with TAR or M-TAR adjustment. Our results (see panel B in Table 4) show that only in the version with rate on personal loans, we can reject the null hypothesis for both TAR and M-TAR models at the 5 percent level of significance. Models with other interest rates (Fed funds rate, 3-year constant maturity rate and 10-year constant maturity rate) do not yield sufficient evidence of threshold cointegration. These results support our earlier results of Breitung (2001, 2002).

Sources of asymmetries in credit demand: our intuition

In the light of our empirical results, it is difficult to identify the direct sources of asymmetries in the credit demand model. Our results encourage us to assume that asymmetries in our data are linked to the borrowing cost i.e. interest rate on personal loans. On the other hand, we find strong evidence of symmetric long-run relationship between credit and other determinants (income, wealth, Fed funds rate, 3-year constant maturity rate and 10-year constant maturity rate). We achieve non-linearity only when credit model is augmented with the rate on personal loans. While it is difficult to explain the direct sources of this asymmetric result, our intuition is that it is linked to the structural breaks in rate of personal loans encountered in the early 1980s (1980Q4 and 1981Q1–Q3, see Table 2).

In what follows, we try to exclude the structural breaks associated with the rate on personal loans and re-assess the existence of asymmetric effects in the credit demand model. Table 2 indicates the break dates (statistically significant) linked to i (i.e. 1981Q1, 1980Q4 and 1981Q3 in LS test and 1981Q2 and 1981Q3 in CKP test). Excluding these break dates would mean that our sample period becomes 1982Q1–2011Q3.[12] To test for asymmetric effects in the sub-sample data, we utilize Enders and Siklos test. The credit demand specification used is $C = f(Y, W, i)$, where i is the rate on personal loans. The F-statistic $\rho_1 = \rho_2 = 0(\Phi_\varepsilon \text{ or } \Phi_\varepsilon^*)$ for the null hypothesis of no threshold cointegration in TAR and M-TAR models are 7.948 and 6.011, respectively. The 5% critical values for Φ_ε (TAR) and Φ_ε^* (M-TAR) are 10.160 and 12.158, respectively. Results suggest that we cannot reject the null hypothesis of no

[12] Sample period prior to 1980 is too short and cannot be used for testing in this case.

Table 4
Enders and Siklos asymmetric cointegration test, 1972:Q2–2011:Q3

Panel A: Estimates of long-run model (OLS method)

	$C = f\,(Y,\ W,\ i)$	$C = f\,(Y,\ W,\ i^{ff})$	$C = f(Y,\ W,\ i^{3Y})$	$C = f\,(Y,\ W,\ i^{10Y})$
Intercept	−13.290 (9.12)***	−5.205 (10.27)***	−2.892 (2.91)**	−9.054 (14.25)***
Y	0.828 (7.50)***	0.619 (4.63)***	0.728 (16.35)***	0.921 (4.52)***
W	0.512 (4.16)***	0.404 (2.71)**	0.621 (4.56)***	0.503 (4.38)***
Interest rate measure	−0.012 (6.07)***	−0.009 (1.74)*	−0.005 (4.31)***	−0.008 (2.41)**

Panel B: Threshold cointegration

	TAR	M-TAR	TAR	M-TAR	TAR	M-TAR	TAR	M-TAR
ρ_1	−0.201 (6.18)***	−0.372 (3.47)***	−0.098 (1.71)*	−0.120 (0.63)	−0.364 (3.92)***	−0.298 (0.64)	−0.501 (1.40)	−0.490 (1.82)*
ρ_2	−0.348 (4.52)***	−0.279 (2.74)**	−1.738 (0.92)	−0.391 (1.58)	−0.189 (0.76)	−0.017 (1.68)*	−0.319 (1.25)	−0.092 (1.31)
α_1	−0.301 (3.95)***	−0.359 (4.17)***	−0.480 (2.36)***	−0.407 (1.69)*	−0.120 (1.48)	−0.545 (1.36)	−0.158 (1.40)	−0.328 (1.99)**
th	0.013	0.004	0.009	−0.100	0.006	0.001	0.039	0.002
BG(1) LM test	0.245 [0.99]	2.162 [0.25]	0.128 [0.82]	3.280 [0.17]	0.725 [0.41]	0.216 [0.94]	2.631 [0.22]	0.116 [0.99]
BG(5) LM test	0.172 [0.96]	0.864 [0.39]	0.210 [0.94]	0.239 [0.79]	0.124 [0.99]	0.790 [0.45]	0.168 [0.87]	0.824 [0.40]
$\rho_1 = \rho_2 = 0(\Phi_\varepsilon \text{ or } \Phi^*_\varepsilon)$	19.158***	22.041***	10.827	8.348	11.511**	6.833	9.820	8.300
$\rho_1 = \rho_2$ (*F-test*)	9.261***	7.377***	1.260	0.930	1.500	0.841	1.233	0.962

Notes: Absolute t-statistics are reported in the parentheses and p-values are in square brackets. ***, ** and * signify significance at 1%, 5%, and 10% levels, respectively. *th* is the threshold level endogenously determined according to Chan's (1993) method. BG (p) = Bresuch–Godfrey test for serial correlation of order p. $\rho_1 = \rho_2 = 0$ is the F-statistic for the null hypothesis of no threshold cointegration; 5% critical values for Π_ε(TAR) and Φ^*_ε (M-TAR) is 12.836 and 13.729, respectively. 10% critical values for TAR and M-TAR are 10.925 and 11.238, respectively. The critical values for Π_ε and Φ^*_ε are simulated according to Wane *et al*. (2004) approach. $\rho_1 = \rho_2$ is the F-statistic that the two coefficients are equal.
Asymmetric Dickey-Fuller equation: $\Delta\varepsilon_t = I_t\ \rho_1\ (\varepsilon_{t-1}) + (1-I_t)\ \rho_2\ (\varepsilon_{t-1}) + \alpha_1\Delta\varepsilon_{t-1} + \upsilon_t$.

threshold cointegration in both TAR and M-TAR models at the 5 percent level. This implies that the asymmetric effects on credit demand are largely observed only in the early 1980s (reflecting on the full-sample and sub-sample results). However, the channels through which these breaks exert asymmetric behaviour are not yet known. We hope that future research may focus on this issue. On the basis of these results, we argue that asymmetric impacts on credit demand are not persistent over-time.

The existence of asymmetries on credit demand in the early eighties suggests that they may be linked to the Volker's experiment. Volcker did engage in severe tightening during 1979–1981 to reduce inflation and thereafter, pursued steady and low-inflation policies. The disinflation in Volcker's regime was due to targeting non-borrowed reserves and probably not through an explicit interest rate targeting. The nominal interest rate was characterized by high levels of volatility.

Asymmetric effects of interest rates on other macroeconomic variables have been well documented in the literature. Sensier *et al*. (2002) used non-linear models to examine the issues in the context of interest rate effects on quarterly UK GDP growth. They found strong evidence of non-linearity, with asymmetry relating to the business cycle through lagged GDP regimes and interest rate changes. Non-linear borrowing cost is also supported by Arias *et al.* (2000). They developed a model, in the context of agricultural sector, where optimal hedging is evaluated under non-linear borrowing costs. Demirtas (2006) evaluated the non-linear asymmetric models of the short-term interest rate. Their empirical results suggest that the non-linear asymmetric models are better than the existing (symmetric) models in forecasting the future level and volatility of interest rate changes.

Did the Great Recession contribute to credit demand asymmetries?

We investigate whether the asymmetries in credit demand are any way associated to the Great recession. In doing so, we test for the presence of asymmetric impacts in the credit model ($C = f\ (Y,\ W,\ i)$) by splitting the sample as follows: (1) sample prior to the Great recession (1972Q2–2006Q4), (2) sample including *some periods* of Great recession (1972Q2–2008Q4) and (3) sample including *main perio*ds of Great recession (1972Q2–2009Q4). We utilize Enders and Siklos test to explore the existence of asymmetries in the credit demand model for the above sample periods. Table 5 reports these results from the M-TAR model.[13] The results reveal that the null of no cointegration is rejected against the alternative of asymmetric cointegration in all cases. The estimates of error correction are not very different across sub-samples (or overtime). These results imply that asymmetries in the credit demand model is not strongly linked to the Great recession. In fact, asymmetries are present in the sample prior to the Great recession period, implying that there are other

[13] TAR model yield consistent results, these are not reported to conserve space. For the M-TAR model, we report only the key result for the purpose of brevity. Additional results are available from the authors upon request.

Table 5
Enders and Siklos tests on sub-sample periods

	Sub-sample 1972Q2–2006Q4	Sub-sample 1972Q2–2008Q4	Sub-sample 1972Q2–2009Q4
ρ_1	−0.231 (4.30)***	−0.185 (−3.07)***	−0.260 (−2.16)**
ρ_2	−0.295 (2.38)**	−0.299 (4.25)***	−0.341 (2.68)**
α_1	−0.179 (1.87)*	−0.202 (2.04)**	−0.331 (1.75)*
$\rho = \rho = 0(\Phi_\varepsilon or \Phi_\varepsilon^*)$	24.301	18.264	27.025
95% CV	13.729	13.729	13.729

Notes: Results are based on M-TAR model. Model specification used: $C = f(Y, W, i)$. CV = critical value. See notes of Table 4 for more details.

factors (for example, non-linearity in the rate on personal loans) associated with this behaviour. This seems to imply that in fact there has been not so much irrationality in the credit market behaviour, in spite of all the volatility surrounding the 'Great recession'.

IV Conclusion and Implications

This paper has investigated the existence of asymmetries in the demand for consumer credit in the United States. The application of Lee and Strazicich (2003) and Carrion-i-Silvestre *et al*. (2009) unit root tests indicate that the variables (real consumer credit, real disposable income, real net wealth, real rate on personal loans, real federal funds rate, real 3-year constant maturity rate and real 10-year constant maturity rate) are I (1) in levels. The break dates yielded by both tests are fairly consistent and match with the timings of macroeconomic events that were experienced by the United States.

We tested for the presence of asymmetries in the credit demand using Breitung (2001, 2002) and Enders and Siklos (2001) methods. We find overwhelming evidence of symmetric credit demand relationship in the presence of exogenous factors such as income, wealth, federal funds rate, 3-year constant maturity rate and 10-year constant maturity rate. However, when the credit demand model is augmented with the rate on personal loans, all tests point towards the existence of asymmetric cointegration. While it is difficult to explain the direct sources of this asymmetric result, our intuition is that it is linked to the structural breaks in the rate of personal loans encountered in the early 1980s. We find that excluding the breaks in the data, we achieve robust symmetric credit demand relationship (i.e. credit, income, wealth and rate on personal loan). Moreover, we find no strong evidence that much of the asymmetric impacts in consumer credit were experienced in the Great recession. Neither do we attain evidence that asymmetric impacts on credit demand are persistent over-time.

Our findings imply that stabilizing or targeting the rate on personal loans is vital to achieve smoothness in the credit demand. The presence of asymmetric associations between credit and rate on personal loans implies some form of myopic or otherwise irrational behaviour in the credit market. It is therefore important to investigate the possible sources, outcomes or results when credit market is characterized by such behaviour. The lower growth rate in credit was observed in the United States. recently.[14] We shed some light on this issue, arguing that credit growth could be stimulated through focusing on 3-year and 10 constant maturity rates. The Fed funds rate is close to the zero lower bound and hence offers limited opportunity to create an impact on the credit market. Moreover, we do not have evidence to support that the associated asymmetries in the credit demand set the stage for the crisis and severe recession that followed from the beginning of 2007–2008 (Great recession).

Appendix A: Unit Root Tests

Lee and Strazicich test

To test for unit root of the series, we first employ the Lee and Strazicich's (2003) two break minimum Lagrange multipliers (*LM*) test. Other unit root tests (e.g. Lumsdaine and Papell (1997)) suffer from bias and spurious rejections in the presence of structural breaks under the null. The two-break LM test does not suffer from bias and spurious rejections and is mostly invariant to the size, location and misspecification of the breaks. This test determines the break dates endogenously. There exist two models (A and C) that have different assumptions about structural breaks. Model A accommodates two shifts in the intercept. Model C contains two shifts in the intercept and trend. Model specifications are as follows:

Model A:

$$\begin{aligned} &Z_t = [1, t, D_{1t}, D_{2t}]' \\ &(D_{jt} = 1 \text{ for } t \geq T_{Bj} + 1, j = 1, 2, \text{ and } 0 \text{ otherwise}) \end{aligned} \tag{A1}$$

Model C:

$$\begin{aligned} &Z_t = [1, t, D_{1t}, D_{2t}, DT_{1t}, DT_{2t}]' \\ &(DT_{jt} = t - T_{Bj} \text{ for } t \geq T_{Bj} + 1, j = 1, 2, \text{ and } 0 \text{ otherwise}) \end{aligned} \tag{A2}$$

T_{Bj} denotes the break date. Equation (A3) and (A4) state the null and alternative hypothesis of the two models respectively.

$$\begin{aligned} &H_0 : y_t = \mu_0 + d_1 B_{1t} + d_2 B_{2t} + y_{t-1} + v_{1t}; \\ &H_1 : y_t = \mu_1 + \gamma t + d_1 D_{1t} + d_2 D_{2t} + v_{2t}; \end{aligned} \tag{A3}$$

$$\begin{aligned} &H_0 : y_t = \mu_0 + d_1 B_{1t} + d_2 B_{2t} + d_3 D_{1t} + d_4 D_{2t} + y_{t-1} + v_{1t}; \\ &H_1 : y_t = \mu_1 + \gamma t + d_1 D_{1t} + d_2 D_{2t} + d_3 DT_{1t} + d_4 DT_{2t} + v_{2t}; \end{aligned} \tag{A4}$$

v_{1t} and v_{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:

$$\Delta y_t = \delta' \Delta Z_t + \phi \overline{S}_{t-1} + \mu_t \tag{A5}$$

[14] http://www.federalreserve.gov/newsevents/speech/duke20101202a.htm

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 $\overline{\delta}; \overline{\psi}_x = y_1 - Z_t\overline{\delta}$ 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).

Carrion-i-Silvestre, Kim and Perron test

To assess robustness of our unit root results, we also employ the unit root test procedure developed by Carrion-i-Silvestre *et al.* (2009). This test allows for multiple breaks in the level and/or slope of the trend function under both the null and alternative hypotheses. In addition, this test adopts the quasi–generalized least squares detrending method advocated by Elliot *et al.* (1996) that allows tests to have local asymptotic power functions close to the local asymptotic Gaussian power envelope. Carrion-i-Silvestre *et al.* (2009) consider a variety of tests, including the feasible point optimal statistic of Elliot *et al.* (1996). For our purpose, we utilize the feasible point optimal statistic given as follows:

$$P_T^{gls}(\lambda^0) = \left\{S(\overline{\alpha}, \lambda^0) - \overline{\alpha}S(1, \lambda^0)\right\}/s^2(\lambda^0) \tag{A6}$$

where λ is the estimate of the break fraction, $\overline{\alpha} = 1 + \overline{c}/T$ ($\overline{c}$ is the non-centrality parameter) and $s^2(\lambda^0)$ is an estimate of the spectral density at frequency zero of v_t.

Asymmetric tests

Breitung tests

Breitung's (2001) proposed rank test for non-linear cointegration. The null hypothesis of no cointegration is tested against the alternative of cointegration in either linear or non-linear form. A good exposition of this test can be found in Haug and Basher (2011) and Liew *et al.* (2009). To test for cointegration among k + 1 series y_t, x_{1t},…, x_{kt}, the following multivariate rank statistic is computed:

$$\Xi_T^* = \frac{T^{-3}\sum_{t=1}^{T}\left(\tilde{u}_t^R\right)^2}{\hat{\sigma}_{\Delta\tilde{u}}^2} \tag{A7}$$

where $\tilde{u}_t^R = R(y_t) - \sum_{j=1}^{k}\tilde{b}_jR(x_{jt})$, 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. $\hat{\sigma}_{\Delta\tilde{u}}^2$ 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). Furthermore, Breitung (2001) also developed a score test statistic that assesses the linearity nature of the cointegrating relationship. The score test statistic is given as follows:

$$\tilde{u}_t = c_0 + c_1x_t + c_2R(x_t) + e_t \tag{A8}$$

where R^2 is the estimate of the determination in equation (A8) and T is the sample size. Using the Stock and Watson's (1993) dynamic ordinary least squares (DOLS) method, the errors ($\tilde{u}_t$) are corrected for serial correlation and endogeneity.

Breitung (2002) proposed the non-parametric test for cointegration. The idea is quite consistent to Johansen's vector error correction model. In order to test for cointegration, the following problem about the $n \times n$ matrix A_T, B_T is considered:[15]

$$\left|\lambda_j B_T - A_T\right| = 0 \tag{A9}$$

where $A_T = \sum_{t=1}^{T} \hat{u}_t \hat{u}_t'$, $B_T = \sum_{t=1}^{T} U_t U_t'$ and $U_t = \sum_{j=1}^{t} \hat{u}_t$ represent the n-dimensional partial sum concerning $\hat{u}_t$. The problem is equivalent to solving the eigenvalue of $R_T = A_T B_T^{-1}$. The solution is:

$$\lambda_j = \frac{(\eta_j' A_T \eta_j)}{(\eta_j' B_T \eta_j)} \tag{A10}$$

where η_j is the eigenvalue of λ_j. $T^2\lambda_j$ diverges to infinity when the vectors of the stochastic trends are less than q. Given that the stochastic trends are associated with each other, this implies the existence of a cointegrating vector. Hence, the test statistic is the following:

$$\Lambda_q = T^2 \sum_{j=1}^{q} \lambda_j \tag{A11}$$

where $\lambda_1 \leq \lambda_2 \leq \ldots \leq \lambda_n$ is the ordered eigenvalues of R_T. This statistic tests whether a q-dimensional stochastic component is rejected at the significance level.

Enders and Siklos test

Enders and Siklos (2001) proposed a test for threshold cointegration among the series. The threshold autoregressive (TAR) model is specified as follows:

$$\Delta\mu_t = I_t \rho_1 \mu_{t-1} + (1 - I_t)\rho_2 \mu_{t-1} + \varepsilon_t \tag{A12}$$

$$I_t = \begin{cases} 1 & \text{if } \mu_{t-1} \geq \tau \\ 0 & \text{if } \mu_{t-1} < \tau \end{cases} \tag{A13}$$

where μ_t is the disturbance term, ρ is the coefficient of $\mu_{t\text{-}1}$, I_t is the Heaviside indicator and τ is the value of the threshold.

According to Enders and Siklos, it is also probable to let the adjustment to depend on the change in μ_{t-1} ($\Delta\mu_{t-1}$). This is because we do not have information about the exact nature of non-linearity in the relationship. To this end, equation (A13) is rewritten as follows:

[15] This test is well-discussed in Holmes and Panagiotidis (2009). They employed this approach to test asymmetries associated with the US current account.

$$I_t = \begin{cases} 1 & \text{if } \Delta\mu_{t-1} \geq \tau \\ 0 & \text{if } \Delta\mu_{t-1} < \tau \end{cases} \tag{A14}$$

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 is serially correlated, equation (A12) is rewritten as:

$$\Delta\mu_t = I_t\rho_1\mu_{t-1} + (1 - I_t)\rho_2\mu_{t-1} + \sum_{i=1}^{p} \gamma_i \Delta\mu_{t-p} + \varepsilon_t \tag{A15}$$

The test statistic Φ is utilized to test for threshold cointegration. To compute Φ statistic, an F-statistic is used which tests the null hypothesis $\rho_1 = \rho_2 = 0$. Φ_μ is the F-statistic for the null hypothesis $\rho_1 = \rho_2 = 0$ in the TAR formulation. On the other hand, Φ_μ^* is the F-statistic in the M-TAR formulation. The critical values to test the null hypothesis of cointegration are tabulated in Enders and Siklos (2001) and Wane *et al.* (2004).

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