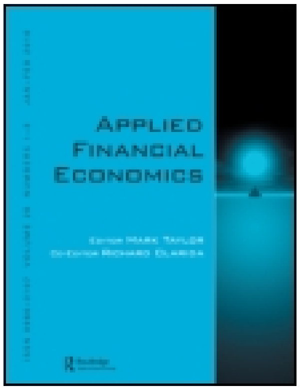

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What caused the equity withdrawal mechanism? An investigation using threshold cointegration and error correction

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What caused the equity withdrawal mechanism? An investigation using threshold cointegration and error correction

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This work investigates the mortgage equity withdrawal mechanism in the US economy from an empirical perspective. Using the threshold cointegration test of Enders and Siklos (2001), which allows for asymmetric adjustment, we find a cointegrating relationship among mortgage equity withdrawal, house prices and interest rates. In particular, we find that the speed of adjustment towards equilibrium is highly persistent above the appropriate estimated threshold, namely in the presence of favourable news. This finding is consistent with the theory of habit formation (Duesenberry, 1949) and conspicuous consumption (Veblen, 1899). Furthermore, this result helps to understand the complex issue of the consumption boom of the late 1990s.

Keywords: mortgage equity withdrawal; house prices; mortgage interest rates; threshold cointegration; habit consumption

JEL Classification: D12; C22; O51

I. Introduction

The Mortgage Equity Withdrawal (MEW) – defined as the amount of equity that consumers withdraw from their homes through home equity loans or lines of credit and cash-out refinancing – is considered an important variable in explaining the extraordinary consumption growth in the last 20 years. Hatzius (2006) argues that MEW has a statistically significant and large effect on consumer spending; between 50% and 62% of MEW flows into consumption. Smith and Searle (2008) maintain that MEW is the mechanism for transmitting the wealth effects of housing into the whole economy, in particular from the 1990s. A cross-country study by Catte *et al.* (2004) shows that 20% of MEW goes into consumption for the US economy. For understanding the contribution of home equity extraction on the economy, Riholtz (2009) calculates that MEW accounted for more than 75% of the Gross Domestic Product (GDP) growth from 2003 to 2006.

Notwithstanding the interest in studying the MEW contribution on consumption and output growth, to the best of our knowledge there are few studies in the literature that explain the theoretical and empirical aspects of the MEW mechanism.[1] In this study, we attempt to address the problem from an empirical

[1] The majority of the empirical works in the literature study the MEW using panel data (see, e.g. Banks *et al.*, 2004; Case *et al.*, 2005; Schwartz *et al.*, 2008).

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point of view by examining the variables that drive home equity extraction. Following the work of Duca and Kumar (2011), we identify house prices and mortgage interest rates as the key variables driving equity extraction. Our empirical methodology is based on the threshold cointegration technique of Enders and Siklos (2001), which allows for asymmetric adjustment. This is important because it is possible that homeowners will differ in the sensitivity (or velocity) of their reactions to positive or negative shocks. Our results show that departures from the long-run equilibrium, above (due to favourable news) and below (due to unfavourable news) an appropriate threshold, have different patterns of adjustment back to equilibrium. The adjustment process is highly persistent (almost 1 year) above the threshold and almost immediate (less than two quarters) below the threshold. If the MEW actually fuelled consumption expenditures in the past years (as explained by Greenspan and Kennedy, 2005; Hatzius, 2006; Smith and Searle, 2008), the sluggish adjustment to favourable shocks is consistent with the theories of habit formation (Duesenberry, 1949; Abel, 1990; Carroll *et al.*, 2000; Fuhrer, 2000) and conspicuous consumption (Veblen, 1899; Frank, 1997; O'Cass and McEwen, 2006). Habit consumption occurs because consumers are reluctant to reduce consumption; instead, conspicuous consumption occurs when consumers purchase goods and services to appear richer and acquire status and prestige in the society.

Our findings are also important for a better understanding of the complex issue of the consumption boom of the late 1990s.[2] FED was criticized for fixing interest rates too low for too long after the 2001 crisis, favouring the housing bubble (Schwartz, 2009; Taylor, 2009; Labonte, 2011) in a fear of a new recession/slowdown of the economy (Bernanke, 2010). These extraordinarily low interest rates accelerated the run-up in house prices and favoured the usage of MEW as an ATM machine for funding consumption spending. The easy way to obtain funds from the equity of their homes generated spending habits. The habit formation and the desire of higher consumption level, powered by MEW mechanism, puts the FED in a position of responsibility for having brought with its policy the consumption on unsustainable patterns. This position is in accordance with Taylor's (2009) view.

Our article is organized as follows. Section II describes the MEW equation that we estimated and the threshold cointegration technique. Section III presents the empirical results. In particular, after the data description, the classical Engle–Granger two-step cointegration and the threshold cointegration tests are presented. Then, the asymmetric error-correction model is estimated. This section concludes with a stability test of the estimated equation. Section IV details our conclusions.

II. MEW Equation Specification and Threshold Cointegration Test

The MEW equation specification

Very few papers in the literature describe the MEW mechanism. Duca and Kumar (2011) maintain that the propensity for withdrawing housing equity rises with house price appreciation and with lower interest rates. In the presence of increasing house prices, homeowners can withdraw housing equity by taking out a second mortgage or refinancing their old mortgage with a larger loan.[3] Due to the transaction costs of refinancing, the incentive to withdraw housing equity is enhanced if borrowers encounter low mortgage rates. This suggests that the main variables explaining the MEW mechanism are house prices and mortgage rates. A problem arises regarding the appropriate measure of interest rates, nominal or real. This is because mortgage debt is contracted in nominal terms, such as the payments owed, and thus households could be more sensitive to nominal interest rate movements. For this reason, we consider both real and nominal interest rate measures in the analysis and rely on the estimation results to decide which measure is more appropriate.

According to the above discussion, the MEW functions in the two interest rate formulations are

$$mew_t = \beta_0 + \beta_1 hp_t + \beta_2 imor_t + \mu_t \quad \text{(Model 1)} \quad (1)$$

$$mew_t = \beta_0 + \beta_1 hp_t + \beta_2 rimor_t + \mu_t \quad \text{(Model 2)} \quad (2)$$

where *mew*, *hp*, *imor* and *rimor* denote the mortgage equity withdrawal, house prices, nominal interest rates and real interest rates, respectively. μ_t denotes the residual of the MEW equation. In Equations 1

[2] For a review of this topic, see, for example, Guidolin and La Jeunesse (2007).
[3] An important distinction is between 'active' and 'passive' MEW. While the former consists of cash-out refinancing and home equity borrowing, which are discretionary actions to extract home equity, the latter is the equity released during housing turnover. In our analysis we consider the active MEW, and it is expressed in terms of disposable income. We consider this variable because it is considered by the literature as affecting consumption significantly (see Canner *et al.*, 2002; Hatzius, 2006), and for this reason it is very interesting.

and 2, MEW is measured as a ratio of disposable personal income and *hp* is measured in real (year-on-year) growth rates. Details on data constructions and sources are in Section III.

Threshold cointegration

This article employs the threshold cointegration approach introduced by Enders and Granger (1998) and further refined by Enders and Siklos (2001). Let $\{x_{it}\}_1^T$ denote the observable random variables integrated of order one, i.e. $I(1)$. The long-run equilibrium relationship is given by

$$x_{1t} = \beta_0 + \beta_2 x_{2t} + \cdots + \beta_n x_{nt} + \mu_t \tag{3}$$

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.[4] The second step procedure is given by

$$\Delta\mu_t = \rho\mu_{t-1} + \varepsilon_t \tag{4}$$

where ε_t is the white noise disturbance. If $-2 < \rho < 0$, the long-run equilibrium (3) 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, called the Threshold Autoregressive (TAR) model

$$\Delta\mu_t = I_t\rho_1\mu_{t-1} + (1 - I_t)\rho_2\mu_{t-1} + \varepsilon_t \tag{5}$$

$$I_t = \begin{cases} 1 & \text{if } \mu_{t-1} \geq \tau \\ 0 & \text{if } \mu_{t-1} < \tau \end{cases} \tag{6}$$

where I_t is the Heaviside indicator and τ is the value of the threshold. The threshold value τ, which is unknown, is estimated according to Chan's (1993) method, as suggested by Enders and Siklos (2001). The TAR model with a Consistent threshold is denoted as a TAR-C model.

Since the exact nature of nonlinearity 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 Equation 6 becomes

$$I_t = \begin{cases} 1 & \text{if } \Delta\mu_{t-1} \geq \tau \\ 0 & \text{if } \Delta\mu_{t-1} < \tau \end{cases} \tag{7}$$

This variant of the 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 known as the Momentum-TAR model with a Consistent (M-TAR-C) threshold. 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.

Moreover, Enders and Siklos (2001) have proposed tests when τ is known ($\tau = 0$). In this case, the above two models are called TAR and M-TAR, respectively. When the adjustment process (Equation 5) is serially correlated, Equation 5 is rewritten as

$$\Delta\mu_t = I_t\rho_1\mu_{t-1} + (1 - I_t)\rho_2\mu_{t-1} + \sum_{i=1}^{p}\gamma_i\Delta\mu_{t-p} + \varepsilon_t \tag{8}$$

To test for threshold cointegration, Enders and Siklos (2001) proposed the Φ-test statistic. 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 (6) and M-TAR specification of (7) are called Φ_μ and Φ_μ^*, respectively. The critical values to test the null hypothesis in the case of three variables are tabulated by Enders and Dibooglu (2001). 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.

III. Empirical Results

Data and unit root tests

The variables used in our analysis are active MEW, house prices and mortgage rates in nominal and real terms. MEW is the equity extracted from the existing homes via cash-out refinancing, home equity borrowing and housing turn-over. The data is taken from Greenspan and Kennedy's (2008) dataset. In our analysis, we consider the active MEW (see footnote 3 on this point) and it is expressed in terms of disposable income; we denote this variable as *mew*. House prices are the year-on-year growth rate of

[4] Enders and Siklos (2001) use this method in their procedure. For maintaining coherence with their procedure, we also use the OLS method.

Standard and Poor's/Case–Shiller home price index (deflated for the US consumer price index). This variable is denoted as *hp* in our analysis. The mortgage interest rate (taken from the Federal Reserve Bank of St. Louis data set (FRED)) is considered in nominal (*imor*) and real (*rimor*) terms for reasons explained in Section II. Time series plot of the data and descriptive statistics are reported in the Appendix.

The integrated properties of the variables are tested with Lee and Strazicich's (2003) two-break minimum Lagrange Multipliers (LM) unit root tests. This is because the period under investigation (1990Q2–2008Q2) is characterized by important changes and events. The three main events are the 'Dot-com' bubble (1998–2001), the house price bubble (2001–2007) and the 2008 stock price crash. Other important events are the technological progress in the 1990s, financial deregulations in the last 20 years (see, e.g. Sherman (2009) on this point) and the globalization of the markets. The break dates, in Lee and Strazicich's test, are endogenously determined and can be explained using a model allowing for two shifts in the intercept (Model A) and trend (Model C) as follows:

$$\text{Model A: } Z_t = [1, t, D_{1t}, D_{2t}]' \quad \left(D_{jt} = 1 \text{ for } t \geq T_{Bj} + 1, j = 1, 2 \text{ and 0 otherwise}\right) \tag{9}$$

$$\text{Model C: } Z_t = [1, t, D_{1t}, D_{2t}, DT_{1t}, DT_{2t}]' \quad \left(DT_{jt} = t - T_{Bj} \text{ for } t \geq T_{Bj} + 1, \; j = 1, 2, \text{ and 0 otherwise}\right) \tag{10}$$

where T_{Bj} denotes the break date. Equations 9 and 10 state the null (H_0) and alternative hypothesis (H_1) of the two models, respectively:

$$\begin{aligned} H_0&: y_t = \mu_0 + d_1 B_{1t} + d_2 B_{2t} + y_{t-1} + \nu_{1t}; \\ H_1&: y_t = \mu_1 + \gamma t + d_1 D_{1t} + d_2 D_{2t} + \nu_{2t} \end{aligned} \tag{11}$$

$$\begin{aligned} H_0&: y_t = \mu_0 + d_1 B_{1t} + d_2 B_{2t} + d_3 D_{1t} + d_4 D_{2t} + y_{t-1} + \nu_{1t}; \\ H_1&: y_t = \mu_1 + \gamma t + d_1 D_{1t} + d_2 D_{2t} + d_3 DT_{1t} + d_4 DT_{2t} + \nu_{2t}; \end{aligned} \tag{12}$$

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:

$$\Delta y_t = \delta' \Delta Z_t + \phi \overline{S}_{t-1} + \mu_t \tag{13}$$

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 null hypothesis of a unit root against otherwise. The optimal lag lengths (from a maximum of eight lags) are selected using the *t*-sig method of Ng and Perron (1995).

The results are reported in Table 1. 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 hypothesis cannot be rejected at the 5% level. For the first differences of these variables the unit root null is rejected at the 5% level with the exception of *hp* for Model A, even if the statistic value (−3.78) is very close to the 5% critical value (−3.84) and well above the 10% critical value (−3.50). In the 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 cover different periods of 1990s and 2000s as expected, since, as we said, the period under investigation comprises various and important changes. Anyway, it is interesting to note that the second break appears to occur after the period 2004–2006 for *mew* and *hp*, which coincides with the so-called house prices bubble and the peak of subprime lending.

Symmetric cointegration test

Having established that the variables under investigations – *mew*, *hp*, *imor* and *rimor* – are $I(1)$, a cointegrating relationship is possible. The estimated long-run equilibrium relationships (using Engle–Granger method) in the two versions are

$$mew_t = \underset{(9.74)*}{8.252} + \underset{(5.29)*}{0.110} hp_t - \underset{(-7.28)*}{0.819} imor_t + \mu_t \quad \text{(Model 1)} \tag{14}$$

$$mew_t = \underset{(14.21)*}{6.092} + \underset{(9.97)*}{0.166} hp_t - \underset{(-9.50)*}{0.892} rimor_t + \mu_t \quad \text{(Model 2)} \tag{15}$$

where parentheses show *t*-statistics and * denotes significance at the 1% level. Table 2 presents the results of tests for cointegration.

Since the test results are largely influenced by the number of lags chosen (Haug, 1996), we employ two criteria for sensitivity check: SIC and the data-dependent method (*t*-sig) recommended by Ng and Perron (1995). The maximum lag order ($\rho_{\max}$) is set according to the Schwert (1989) 'rule of thumb'. According to this rule, $\rho_{\max} = \text{int}[12(T/100)^{1/4}]$, where T is the number of observations and *int*

Table 1. Two-break minimum LM unit root test 1990Q2–2008Q2

	Model A				Model C			
	Level		First difference		Level		First difference	
Variables	Test statistic	Break dates	Test statistic	Break dates	Test statistic	Break dates	Test statistic	Break dates
mew	−3.133 [6]	2001Q1; 2004Q2	−5.950 [2]	2003Q1; 2006Q4	−5.140 [3]	2000Q4; 2006Q1	−11.051 [2]	2003Q2; 2005Q2
hp	−2.092 [8]	1994Q1; 2006Q2	−3.782 [0]	2000Q2; 2005Q1	−4.794 [4]	1998Q2; 2004Q3	−6.570 [7]	1994Q1; 2005Q1
imor	−3.446 [3]	1996Q1; 2001Q1	−5.981 [1]	1993Q1; 2000Q4	−4.421 [1]	1993Q4; 2003Q1	−7.087 [3]	1994Q3; 1999Q3
rimor	−3.296 [7]	1997Q1; 2005Q2	−4.378 [1]	1999Q3; 2002Q4	−4.552 [5]	1998Q1; 2006Q2	−7.086 [7]	1994Q3; 2006Q3

Notes: The 5% and 10% critical values for Model A are −3.842 and −3.504, respectively. The 5% and 10% critical values for Model C are −5.286 and −4.989, respectively. The number in square brackets indicates the optimal number of lagged first-differenced terms included in the unit root test. The trimming region is (0.15 *T*, 0.85 *T*), where *T* is the sample size. Critical values are taken from Lee and Strazicich (2003). Kumar and Webber (2013) provide more details on this test. RATS 7.2 was used to perform this test.

Table 2. The results of classical residual-based cointegration test

	μ_t Model 1			μ_t Model 2		
$\rho_{\max}$	Schwert's criteria (lag = 11)	Schwert's criteria (lag = 11)	8	Schwert's criteria (lag = 11)	Schwert's criteria (lag = 11)	8
Optimal ρ	SIC (lag = 1)	*t*-sig (lag = 9)	*t*-sig (lag = 7)	SIC (lag = 0)	*t*-sig (lag = 11)	*t*-sig (lag = 7)
EG test value	−2.694	−1.747	−1.447	−4.524*	−0.537	−1.629

Notes: Schwert's (1989) criteria assumes that $\rho_{\max} = \text{int}[12(T/100)^{1/4}] = 11$. Critical values are −4.05, −3.42 and −3.10 at 1%, 5% and 10% significance level, respectively (MacKinnon, 1991). SIC = Schwartz Information Criterion.
*Denotes significance at the 1% level.

denotes the integer portion of the content in brackets. In our dataset, where $T = 73$, the maximum lag length suggested by this rule is 11. Schwert's rule is a frequent choice in empirical research (Perron and Qu, 2007), but might be regarded as being rather conservative. If the ρ selected is too large then the power of the test suffers, therefore we fix a smaller $\rho_{\max}$ equal to 8, in accordance with Breitung *et al*. (2004) and Maki and Kitasaka (2006), for quarterly data. These authors consider 2 years a sufficiently long period to capture most of the correlation in the residuals.

From Table 2, it emerges that, for model 2 (real mortgage interest rate), only in one case (maximum lag equal to Schwert's rule of thumb and optimal lag selected according to SIC criteria) we can reject the null of no cointegration at the 1% significance level. In other cases we cannot reject the null hypothesis of no cointegration. This suggests that we cannot refuse the hypothesis of no cointegration. Though there is no cointegration in the symmetric EG cointegration test, there might be a possibility of uncovered asymmetric cointegration in the two models.

Asymmetric cointegration test

Table 3 presents the results for cointegration with TAR, M-TAR, TAR-C and M-TAR-C for the nominal (Model 1) and real (Model 2) interest rates. The null hypothesis of no cointegration is rejected for real mortgage rates in all threshold versions, but only M-TAR-C version passes the $\rho_1 = \rho_2$ test. In addition, this formulation exhibits the lower SIC criteria (results available upon request by the author) and for this reason it is preferred. The Φ-statistic for this formulation is 11.6 and is significant at the 1% level. The symmetry test using the *F*-statistic rejects the symmetry of the two adjustment parameters; the negative adjustment parameter ρ_2 is larger (in

Table 3. The results of asymmetric cointegration tests

	Model	Lag	ρ_1	ρ_2	Φ or Φ^*	$\rho_1=\rho_2$	BG(1,4)
Model 1	TAR	1	−0.246 (0.123)*	−0.268 (0.137)*	3.626	0.016	[0.14,0.64]
	M-TAR	1	−0.285 (0.137)**	−0.231 (0.127)*	3.664	0.086	[0.12,0.61]
	TAR-C ($\tau = -1.171$)	1	−0.208 (0.111)*	−0.370 (0.167)**	3.996	0.686	[0.12,0.52]
	M-TAR-C ($\tau = -0.764$)	1	−0.182 (0.093)*	−1.068 (0.306)***	7.906*	7.752***	[0.46,0.96]
Model 2	TAR	0	−0.390 (0.124)***	−0.569 (0.168)***	10.709**	0.734	[0.69,0.70]
	M-TAR	0	−0.392 (0.132)***	−0.531 (0.154)***	10.410**	0.473	[0.32,0.87]
	TAR-C ($\tau = -0.672$)	0	−0.372 (0.116)***	−0.665 (0.188)***	11.370***	1.757	[0.55,0.96]
	M-TAR-C ($\tau = -0.605$)	0	−0.382 (0.108)***	−0.786 (0.238)***	11.633***	2.386*	[0.62,0.97]

Notes: SEs of the coefficients are in parentheses; the statistics reported in square brackets are p-values for serial correlation test. τ is the threshold level endogenously determined according to Chan's (1993) method. BG(p) = Breusch–Godfrey test for serial correlation of order p. $\Phi$is the F-statistic for the null hypothesis of no threshold cointegration. Critical values for Φ are tabulated by Enders and Dibooglu (2001). Critical values for Model 1 (one lag) are 6.24 (10%), 7.49 (5%), 10.15 (1%) for TAR and 6.72 (10%), 7.92 (5%), 10.66 (1%) for MTAR. Critical values for Model 2 (zero lag) are 6.53 (10%), 7.80 (5%), 10.90 (1%) for TAR and 6.98 (10%), 8.30 (5%), 11.4 (1%) for MTAR. $\rho_1 = \rho_2$ is the F-statistic that the two coefficients are equal.
***, ** and * denote significance at the 1, 5 and 10% levels, respectively.

absolute value) than the positive adjustment parameter ρ_1. Diagnostic checking on residuals shows that there is no autocorrelation (Breusch–Godfrey test for autocorrelation conducted at lag one and four), and therefore the model is correctly specified.

Threshold error-correction test

The positive finding of a long-run equilibrium relation with asymmetric adjustment justifies estimating the error-correction representation in the M-TAR-C framework. The three estimated error-correction equations with threshold are given by

$$\begin{aligned}\Delta mew_t = &\underset{(-1.09)}{-0.054} - \underset{(-3.72)***}{0.321 I_t \mu_{t-1}} - \underset{(-5.16)***}{0.704(1-I_t)\mu_{t-1}} \\ &+ \underset{F_{11}<0.01***}{A_{11}(L)\Delta mew_{t-1}} + \underset{F_{12}=0.80}{A_{12}(L)\Delta hp_{t-1}} \\ &+ \underset{F_{13}=0.96}{A_{13}(L)\Delta rimor_{t-1}} + \underset{(4.65)***}{2.142 D_{1t}} \\ &+ \underset{(5.11)***}{2.121 D_{2t}} + \zeta_{1t}\end{aligned} \tag{16}$$

$$\bar{R}^2 = 0.69 \quad JB = 2.21\ [0.33]$$
$$BG(1,4) = [0.75, 0.10] \quad BPG = 2.22\ [0.07]$$

$$\begin{aligned}\Delta rimor_t = &\underset{(-0.18)}{-0.011} - \underset{(-2.39)**}{0.267 I_t \mu_{t-1}} - \underset{(-1.28)}{0.220(1-I_t)\mu_{t-1}} \\ &+ \underset{F_{21}=0.15}{A_{21}(L)\Delta mew_{t-1}} + \underset{F_{22}=0.02**}{A_{22}(L)\Delta hp_{t-1}} \\ &+ \underset{F_{23}=0.11}{A_{23}(L)\Delta rimor_{t-1}} - \underset{(-3.26)***}{1.774 D_{3t}} + \zeta_{2t}\end{aligned} \tag{17}$$

$$\bar{R}^2 = 0.22 \quad JB = 0.37\ [0.83]$$
$$BG(1,4) = [0.52, 0.06] \quad BPG = 0.51\ [0.86]$$

$$\begin{aligned}\Delta hp_t = &\underset{(0.11)}{0.013} - \underset{(-1.08)}{0.221 I_t \mu_{t-1}} + \underset{(0.48)}{0.153(1-I_t)\mu_{t-1}} \\ &+ \underset{F_{31}=0.07*}{A_{31}(L)\Delta mew_{t-1}} + \underset{F_{32}<0.01***}{A_{32}(L)\Delta hp_{t-1}} \\ &+ \underset{F_{33}=0.08*}{A_{33}(L)\Delta rimor_{t-1}} - \underset{(-4.36)***}{4.356 D_{3t}} + \zeta_{3t}\end{aligned} \tag{18}$$

$$\bar{R}^2 = 0.60 \quad JB = 0.57\ [0.75]$$
$$BG(1,4) = [0.07, 0.11] \quad BPG = 2.75\ [0.01]$$

where the estimation period is 1990Q2–2008Q2, $\zeta_t \sim I.I.D.(0, \sigma^2)$, and F equals the p-value for the null hypothesis that the coefficients on the two lags of each variable used are equal to zero. The t-statistics of coefficient estimations are reported in parentheses, while ***, ** and * denote significance at the 1, 5 and 10% levels, respectively; p-values of residual diagnostic tests are in square brackets. D_i are impulse dummies: $D_1 = 2004Q2$ corresponds to a huge increase in the house prices (+11% compared to the same period of the previous year); $D_2 = 2003Q3$ relates to the peak value (billions of $) of mortgage refinancing in that year; $D_3 = 2007Q4$ is the beginning of the US recession (Hamilton, 2010). It is interesting to note that these dummies are in line with the breaks detected by Lee and Strazicich's unit root test, with the exception of the D_3 dummy. This is plausible since Lee and Strazicich's test is unable to detect the break very close to the end of the sample. There is no evidence of serial correlation on the residuals (Breusch–Godfrey (BG) test, conducted at lag 1 and 4, shows p-values larger than the standard significance level of 0.05 in all equations); the residuals appear normal (Jarque–Bera (JB) test exhibits p-values > 0.05), and have a constant variance (Breusch–Pagan–Godfrey (BPG) test for heteroscedasticity with p-values > 0.05) with the exception of house prices error-correction equation (16).

The asymmetric adjustment process of the *mew* mechanism (16) shows that the adjustment is almost immediate (the speed of adjustment coefficient is equal to -0.71) for negative deviations from equilibrium (actual *mew* deviation from equilibrium below the appropriate threshold). This negative disequilibrium (caused by negative news) is absorbed in less than two quarters, whereas a positive disequilibrium (actual *mew* discrepancy from equilibrium above the threshold, due to positive shocks) is eliminated after almost 1 year (the speed of adjustment coefficient is -0.32). This means that good news are not absorbed immediately; the cash-out mechanism is only gradually restored when good news happen. If the MEW is used as an ATM machine (in the words of Klyuev and Mills, 2007) to finance consumption,[5] the result (i.e. $\rho_1 < \rho_2$) is compatible with the theories of consumption of habit formation (Abel, 1990; Carroll *et al.*, 2000; Fuhrer, 2000) and conspicuous consumption (Veblen, 1899; Frank, 1997; O'Cass and McEwen, 2006). According to the first theory, consumption patterns are subject to habit and are slow to fall in the face of negative shocks. 'The fundamental psychological postulate underlying our argument is that it is harder for a family to reduce its expenditure from a higher level than for a family to refrain from making high expenditures in the first place (Duesenberry, 1949, pp. 84–85)'. According to the second theory, '[t]he things we feel we "need" depend on the kinds of things that others have, and our needs thus grow when we find ourselves of others who have more than we do. Yet when all of us spend more, the new, higher spending level simply becomes the norm (Frank, 1997, p. 1840)'. If the MEW is an ATM for consumption expenditures, then it is plausible that consumers cannot reduce the needed cash withdrawals and, eventually, they tend to increase the cash-out over the time. This effect can be exacerbated in the presence of liquidity constraints.

For completeness of the analysis, we also report the threshold error-correction equations for *rimor* and *hp*. Real mortgage rate changes (17) respond to a positive disequilibrium; a rise in *mew*, caused by positive news, implies a complete adjustment to the interest rates after less than four quarters (the speed of adjustment coefficient is -0.27). In addition, the current changes in *rimor* are influenced by lagged changes in *hp* (A_{22}). Changes in house prices (18) are not influenced by the long-run relation, but are influenced by lagged changes in *mew* (A_{31}) and *rimor* (A_{33}).

Table 4. Quandt–Andrews structural break tests, 1990Q2–2008Q2

Statistics	Value	Break	Probability
Max LR F-stat.	2.3124	2003Q3	1.0000
Max Wald F-stat.	28.8115	2003Q3	0.1800
Exp LR F-stat.	0.6850	–	1.0000
Exp Wald F-stat.	8.3359	–	0.1007
Ave LR F-stat.	1.3059	–	1.0000
Ave Wald F-stat.	11.7532	–	0.1549

Notes: Probabilities are calculated using Hansen's (1997) method. Eviews 7.0 was used to perform this test.

Structural breaks and stability

We investigate the stability of our estimated equation (16) in the light of the important changes that occurred in the 20 years under investigation. In doing so, we subject the threshold error-correction equation (16) to the 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 4 displays the Quandt–Andrews test results. Results show that all the test statistics do not reject the null of no structural breaks at the 1% level. The detected break date, although statistically insignificant, is 2003Q3. This break date is not unexpected, since it corresponds to the period of peak value (billions of $) of mortgage refinancing as stated previously. Based on these results, we infer that our estimated equation is stable and robust.

IV. Conclusions

This article investigates a neglected issue in the empirical financial literature: the variables that have driven the MEW mechanism in the last 20 years of the US economy. Following the work of Duca and Kumar (2011), we select house prices and interest rates as key explanatory variables. Then we study the

[5] See Greenspan and Kennedy (2005), Hatzius (2006) and Smith and Searle (2008).

long-run relationship between MEW and the two above-mentioned variables using the threshold cointegration test developed by Enders and Siklos (2001). The threshold cointegration approach provides strong evidence of the long-run relation characterized by asymmetric adjustment. In particular, we show that the adjustment process towards the equilibrium is highly persistent (almost 1 year) above the appropriately estimated threshold; that is, in the presence of favourable news. Comparatively, the adjustment back to equilibrium is quick (two quarters) when disequilibrium is below the appropriate threshold; that is, in the presence of unfavourable news. This finding is consistent with the theories of habit formation (Duesenberry, 1949; Abel, 1990; Carroll *et al*., 2000; Fuhrer, 2000) and conspicuous consumption (Veblen, 1899; Frank, 1997; O'Cass and McEwen, 2006). According to these theories, consumption patterns are affected by 'habit' or 'status and prestige', and for this reason they rarely reduce but, eventually, tend to increase. If the MEW is an ATM for consumption expenditures, then it is plausible that consumers cannot reduce the needed cash withdrawals.

Our main finding (asymmetric adjustment of MEW with respect to long-run equilibrium level) is important for a better understanding of the complex issue of the consumption boom of the late 1990s. After the 2001 recession, the FED – in a fear of a new recession/slowdown of the economy (Bernanke, 2010) – fixed interest rates too low for too long (until the end of 2004), and the resulting low mortgage rates helped to inflate the bubble (Schwartz, 2009; Taylor, 2009; Labonte, 2011). The economy was very slow in recovering from the 2001 recession. The weakness of the recovery led the Federal Reserve Board to continue to cut interest rates, pushing the federal funds rate to 1% in the summer of 2003, a 50-year low. Mortgage interest rates followed the federal funds rate down. The average interest rate on 30-year fixed rate mortgages fell to 5.25% in the summer of 2003, also a 50-year low. These extraordinarily low interest rates accelerated the run-up in house prices, favouring the usage of MEW as an ATM machine for funding consumption spending. The easy way to obtain funds from the equity of their homes generated the desire of higher consumption (caused by the desire to appear richer). In addition, the extraordinary low levels of interest rates altered the consumption path favouring the formation of habit behaviour (the reluctance in reducing consumption as you reach a higher level) on unsustainable high patterns. The habit formation and the desire to appear richer, fuelled by MEW mechanism, puts the FED in a position of responsibility for having brought the consumption on unsustainable patterns with its policy. This position is in accordance with Taylor's (2009) view, who explains that the FED held the federal funds rate too low for too long during the critical years between 2002 and 2005. He argues that, if the FED instead had followed the Taylor Rule, the boom and bust largely would have been avoided. The literature on this topic is expanding rapidly, but more empirical research is still needed.

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Appendix

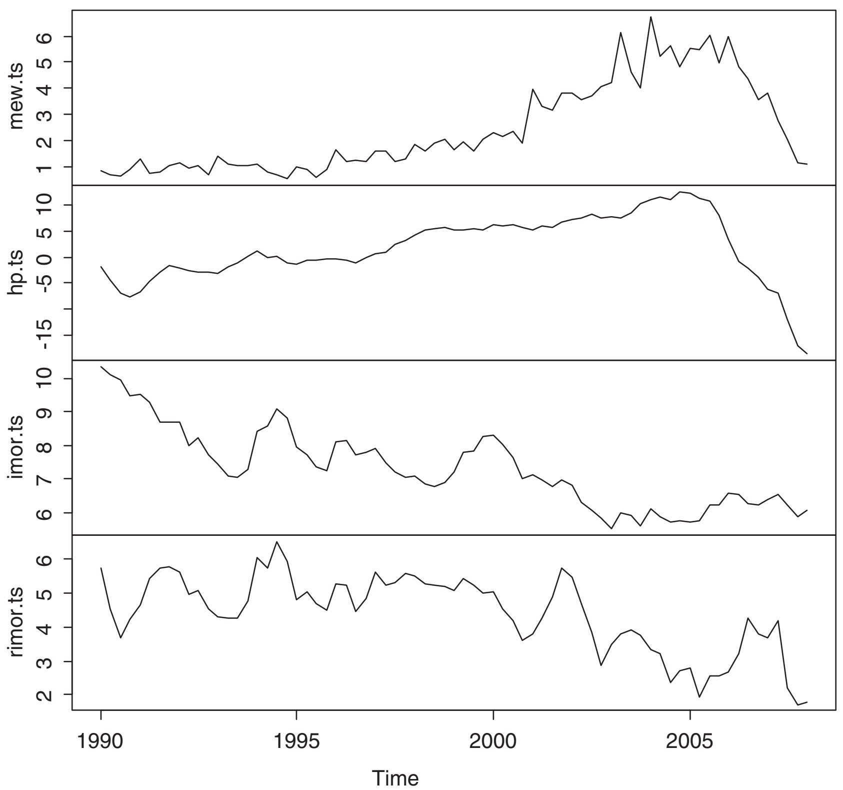

Fig. A1. ***mew*, *hp*, *imor*, *rimor* (1990Q2–2008Q2)**

Table A1. Descriptive statistics

Variable	Mean	SD	Min	Max
mew	2.446	1.737	0.535	6.746
imor	7.316	1.181	5.510	10.340
rimor	4.404	1.131	1.720	6.510
hp	1.705	6.369	−18.472	12.427

Notes: Min = minimum value and Max = maximum value. Data period is from 1990Q2 to 2008Q2.