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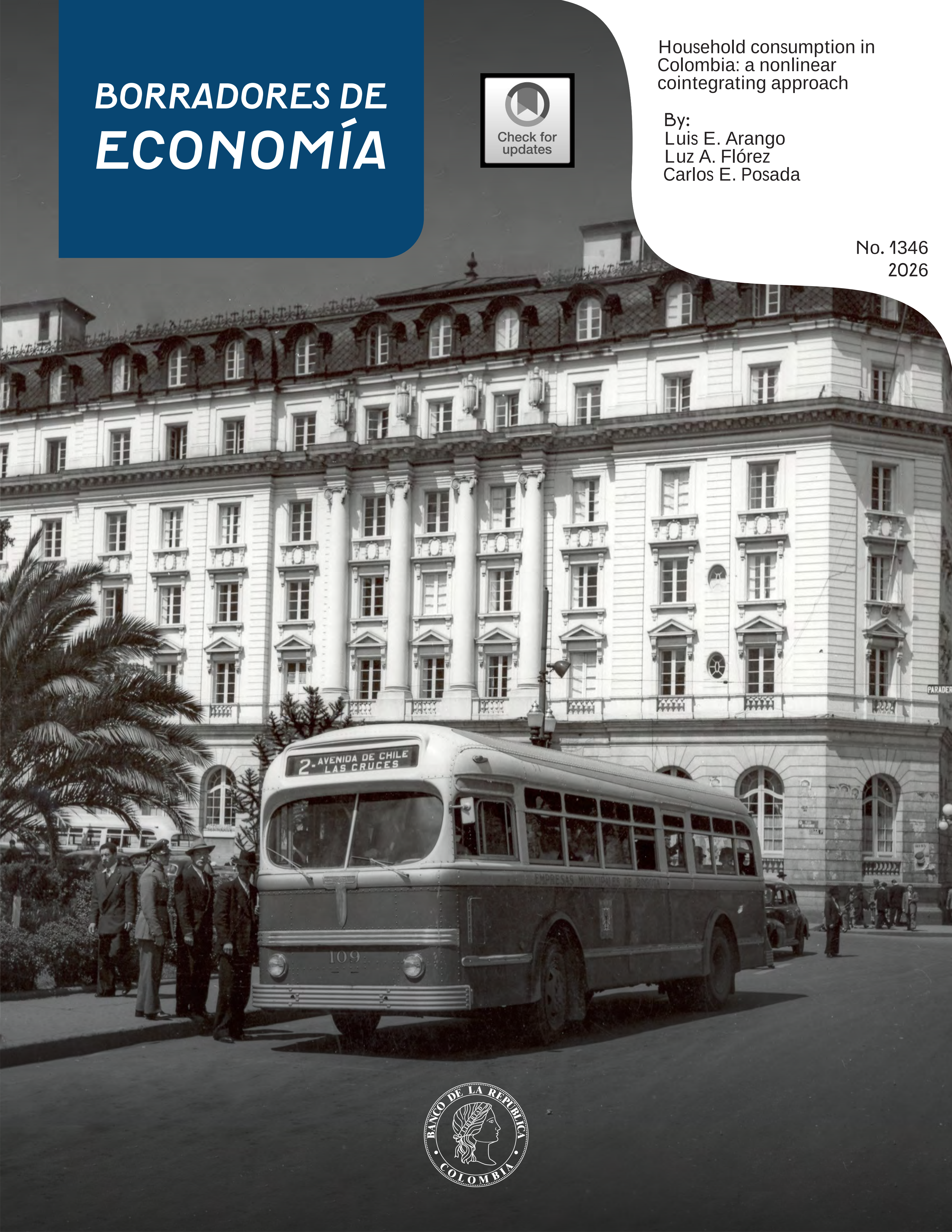

BANCO DE LA REPUBLICA
COLOMBIA

Household consumption in Colombia: a nonlinear cointegrating approach

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Abstract

This study analyzes the short- and long-run determinants of household consumption in Colombia using monthly data from April 2003 to June 2022. We estimate Nonlinear Autoregressive Distributed Lag (NARDL) models to examine asymmetric long-run responses of consumption to positive and negative changes in current income and interest rates. Consumption is measured using the retail trade index excluding motor vehicles and fuels. We find a cointegrating relationship linking consumption with current income (proxied by the Economic Situation Indicator, ESI), remittances, consumer credit, and the real interest rate on consumer credit. In the short run, consumer confidence and the age composition of the population also play a significant role in shaping consumption dynamics. The estimates reveal pronounced long-run asymmetries: income increases are associated with a long-run propensity to consume that is approximately 25 percent larger than that implied by income declines, while interest-rate reductions elicit long-run responses nearly 188 percent larger in absolute value than rate increases. These patterns are consistent with liquidity constraints, with the income-driven asymmetry delivering a superior fit—by standard model-comparison criteria—relative to the interest-rate asymmetry. The negative association between population aging and consumption is difficult to reconcile with a benchmark life-cycle model in the presence of capital-market frictions. It also underscores the need to anticipate headwinds to aggregate demand as population aging proceeds.

***JEL* Classification:** *E*21, *E*27.
Keywords: *Household consumption, nonlinearities, liquidity constraints, precautionary savings, remittances, aging population.*

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Consumo de los hogares en Colombia: un enfoque de cointegración no lineal

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Resumen

Este articulo analiza los determinantes de corto y largo plazo del consumo de los hogares en Colombia utilizando datos mensuales de abril de 2003 a junio de 2022. Estimamos modelos de rezagos distribuidos autorregresivos no lineales (NARDL) para examinar las respuestas asimétricas de largo plazo del consumo ante cambios positivos y negativos en el ingreso corriente y en la tasa de interés. El consumo se mide mediante el índice de comercio al por menor que excluye vehículos automotores y combustibles. Encontramos una relación de cointegración que vincula el consumo con el ingreso corriente (aproximado por el Indicador de Situación Económica, ESI), las remesas, el crédito de consumo y la tasa de interés del crédito de consumo. En el corto plazo, la confianza del consumidor y la composición etaria de la población también desempeñan un papel significativo en la dinámica del consumo. Las estimaciones revelan asimetrías pronunciadas de largo plazo: los incrementos del ingreso generan propensiones a consumir en el largo plazo aproximadamente 25 por ciento mayores que las asociadas a caídas del ingreso, mientras que las reducciones de la tasa de interés inducen respuestas de largo plazo casi 188 por ciento mayores (en magnitud) que los aumentos de la tasa. Estos patrones son coherentes con la presencia de restricciones de liquidez; además, la asimetría impulsada por el ingreso corriente presenta un ajuste superior —según criterios estándar de comparación de modelos— frente a la asimetría asociada a la tasa de interés. La asociación negativa entre envejecimiento de la población y consumo es difícil de conciliar con un modelo de ciclo de vida de referencia dadas las fricciones en el mercado de crédito. Asimismo, subraya la necesidad de anticipar presiones a la baja sobre la demanda agregada a medida que avanza el envejecimiento poblacional.

Clasificación *JEL*: *E*21, *E*27
Palabras clave: *Consumo de los hogares, no linealidades, restricciones de liquidez, ahorro precautelativo, remesas, envejecimiento de la población.*

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1. Introduction

Household consumption is central to macroeconomic analysis both because it directly affects welfare and because it constitutes a large share of GDP. Consequently, its short- and long-run behavior has been extensively studied (see, e.g., Keynes, 1936; Duesenberry, 1949; Modigliani and Brumberg, 1954; Friedman, 1957; Hall, 1978; Campbell and Mankiw, 1989). From a long-run perspective, a standard approach formalizes the relationship between consumption and income through cointegration (Engle and Granger, 1987). Under the benchmark permanent-income/life-cycle view, log consumption and log income are expected to share a common stochastic trend, with a cointegrating vector close to $(1, -1)'$ (see, e.g., King et al., 1991; Han and Ogaki, 1997). If these I(1) variables are cointegrated, the residual from regressing log consumption on log income should be stationary, indicating a stable long-run equilibrium relation. Empirical studies such as King et al. (1991) and Han and Ogaki (1997) provide evidence consistent with this hypothesis, underscoring the centrality of the income–consumption nexus in understanding consumption behavior.

Cointegration-based analyses of household consumption in Colombia have been prominent over the past three to four decades. Carrasquilla (1989) and Posada and Gaviria (1992) report that aggregate consumption shares a long-run common trend with GDP and income, respectively. Hernández (2006) documents a long-term relationship linking household consumption with the housing stock—used as a proxy for wealth—and the interest rate. According to López, Misas, and Oliveros (1996) and Hernández (2006), demographic dependency did not have a statistically significant effect on saving and consumption patterns in Colombia. More recently, Arias et al. (2023) found cointegration among consumption, disposable income, financial wealth, and the interest rate using quarterly data for 2005–2019.[1]

Empirical studies of consumption have traditionally relied on linear and log-linear specifications.[2] However, recent theoretical and empirical work underscores the importance of nonlinear dynamics and asymmetric comovements, challenging the traditional framework. In response, the econometric toolkit has expanded to accommodate these features—encompassing threshold cointegration and asymmetric error-correction models, ARDL bounds testing and its nonlinear extension (NARDL), and

[1] From a different standpoint, Muñoz (2004) analyzes data from the 1994–1995 Income and Expenditure of Households Survey to examine the microeconomic determinants of households' current expenditure on goods and services, emphasizing the role of income. Using an approach similar to that of Muñoz (2004) and closely related to Miles (1997), Bande, Riveiro, and Ruiz (2021) find that income uncertainty increases consumption, except when households have savings capacity. Arango et al. (2024) investigate the principal drivers of consumption growth—including the current income, consumer credit, consumer confidence, remittances, and the real interest rate—among others.

[2] The long-run relationship between consumption and income has been extensively studied within the cointegration framework (see, e.g., Engle and Granger, 1987).

nonlinear/semiparametric cointegration (see Balke and Fomby, 1997; Pesaran and Shin, 1998; Park and Phillips, 2001; Enders and Siklos, 2001; Shin, Yu, and Greenwood-Nimmo, 2013).

Motivated by the literature reviewed below, we employ both linear ARDL and nonlinear ARDL (NARDL) cointegration frameworks to examine how household consumption in Colombia adjusts to movements in current income—proxied by the Economic Situation Indicator (ESI)—remittances from abroad,[3] consumer credit, and interest rates over 2003–2022. The NARDL model (Shin, Yu, and Greenwood-Nimmo, 2013) accommodates short- and long-run asymmetries by decomposing each explanatory variable into positive and negative partial sums, allowing the response of consumption to differ for increases versus decreases (and by magnitude). This approach enables tests of several hypotheses about nonlinear consumption behavior in Colombia—including liquidity constraints, buffer-stock saving, income uncertainty, and mental accounting—which we discuss in the next section.

This article helps fill a gap in research on Colombian household consumption in the 21st century, with notable exceptions including Muñoz (2004), Hernández (2006), Iregui and Melo (2009), Bande, Riveira, and Ruiz (2021), Arias et al. (2023), and Arango et al. (2024a, 2024b). We analyze long-run nonlinear responses of household consumption to changes in current income, remittances, and real interest rates using monthly data for 2003–2022, a period that includes the COVID-19 pandemic. We also consider additional drivers emphasized by Arango et al. (2024)—consumer credit and consumer confidence. Finally, we examine how consumption is affected by Colombia's rapid population aging, set against the backdrop of the large influx of predominantly young Venezuelan migrants since 2015. To the best of our knowledge, no prior study has systematically assessed asymmetric responses of Colombian household consumption to changes in income, remittances, and interest rates using nonlinear cointegration methods.

Using monthly data for 2003–2022, we find that current income, remittances, consumer credit, and real interest rates share a long-run relationship with household consumption. We also uncover significant asymmetries: consumption responds more strongly to increases than to decreases in income, and it reacts differently to interest-rate hikes than to reductions.[4] In our preferred NARDL specification, the coefficient on positive income changes (0.9782) is approximately 25 percent larger than the coefficient on negative income changes (0.7832), indicating higher sensitivity to income expansions in the long run. This pattern is more consistent with liquidity constraints than with the buffer-stock saving model:

[3] Over 2000–2022, remittance inflows averaged 2.01 percent of GDP, according to data from Banco de la República.

[4] Our results—specifically, the sign of the estimated consumption responses—is consistent with Baugh, Ben-David, Park, and Parker (2021) and Coşkun, Apergis, and Coşkun (2022). While we analyze consumption within a long-run framework, these studies examine different settings and datasets: the former investigates temporary income changes arising from tax refunds and adjustments, and the latter analyzes stationary series under different tax regimes.

households appear not to accumulate precautionary savings during expansions but instead spend a large share of current income. The weaker response to adverse income movements further supports liquidity constraints in the sense of Shea (1995) and is consistent with the prominent role of remittances in relaxing borrowing constraints.

We find pronounced interest-rate asymmetries: the effect of real interest-rate cuts on consumption is about 188 percent larger (in magnitude) than the effect of rate hikes. This pattern is consistent with liquidity-constraint mechanisms (see Juster and Shay, 1964; Arango and Cardona-Sosa, 2023; Arango and Quevedo-Rocha, 2024). Consistent with Arango et al. (2024), the two-period-lagged, annual change in the consumer confidence index is a key short-run predictor of consumption. This index—published monthly by Fedesarrollo[5]—serves as an uncertainty proxy that conveys information beyond current income and remittances.

Finally, we test the hypothesis—often associated with the life-cycle model—that population aging is not a major driver of aggregate consumption under perfect capital markets. Contrary to this prediction, our estimates indicate that population aging significantly reduces household consumption. This finding is particularly salient given Colombia's rapid demographic transition and helps explain recent consumption dynamics. Because all these results can be interpreted through multiple theoretical lenses, we adopt an eclectic, integrative perspective in the discussion that follows.

The remainder of the paper is organized as follows: Section 2 reviews key perspectives on household consumption behavior, with emphasis on nonlinear responses to changes in its determinants. Section 3 describes the data. Section 4 outlines the empirical model. Section 5 presents and discusses the results. Section 6 concludes with preliminary lessons.

2. Asymmetric Consumption Responses: Theory and Evidence

Across a range of economies, studies document robust asymmetries in consumption behavior, particularly when consumption is measured in stationary (detrended/cyclical) terms rather than in nonstationary levels. Building on Hall (1978) for the United States and Boone et al. (1998) for Canada, Germany, Japan, the Netherlands, and the United Kingdom, Shirvani and Wilbratte (2000) find that consumption in Germany and the United States responds more to stock-price declines than to increases; however, they report that this effect is short-lived, dissipating after about seven quarters. Case (2005, 2011) presents evidence that declines in wealth—especially housing prices—do not translate one-for-

[5] This is of the leadings think-tanks in Colombia. Information about the consumer confidence index can be found at https://www.repository.fedesarrollo.org.co/handle/11445/36.

one into lower household consumption, while Apergis and Miller (2006) show that, in the short-run adjustment process, per capita consumption responds less to good news than to bad news. For the United Kingdom, MacDonald, Mullineaux, and Sensarma (2011) find that wealth reductions associated with a tightening monetary-policy stance have a smaller impact on consumption than comparable wealth increases.

Within the consumption–income nexus, Shea (1995) shows that consumption is more sensitive to predictable income declines than to increases, contradicting the symmetry implied by the life-cycle/permanent-income hypothesis (LC/PIH).[6] According to him, this pattern is inconsistent with both myopia—which would predict similar responses to income increases and decreases—and standard liquidity-constraint models, which imply stronger responses to anticipated income increases than to declines because borrowing constraints impede borrowing in downturns but not saving in upturns. Shea argues that the observed asymmetry is better explained by reference-dependent preferences with loss aversion (Bowman, Minehart, and Rabin, 1993), in which utility is concave for gains relative to a reference point and convex for losses, making consumption more reactive to adverse income news.

Souleles (1999) documents excess sensitivity of consumption to income-tax refunds. Responses are strongest among liquidity-constrained households, whose nondurable spending spikes upon receipt, while less-constrained households adjust primarily through durable purchases. Building on prospect theory, Bowman, Minehart, and Rabin (1999) model reference-dependent preferences with loss aversion: when income uncertainty is high, households resist cutting consumption in response to bad news about future income but adjust sharply when the loss materializes; the asymmetry is stronger for losses than gains. Using an internet-based survey of Dutch households, Christelis, Georgarakos, Jappelli, Pistaferri, and van Rooij (2017) find that consumers' responses to positive income surprises are smaller than to negative ones, with average marginal propensities to consume (MPCs) around 0.15–0.25 and larger for adverse shocks.[7] Consistent evidence from the United Kingdom shows marked asymmetry: Bunn, Le Roux, Reinold, and Surico (2018) estimate MPCs of 0.46–0.68 for negative income shocks versus 0.07–0.17 for positive shocks.

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[6] This finding contrasts with Altonji and Siow (1987), who provide empirical evidence that households anticipating income increases exhibit greater consumption sensitivity to predictable income changes than households anticipating income declines.
[7] The empirical estimates show that, when liquidity constraints bind, the marginal propensity to consume (MPC) out of a negative income shock exceeds the MPC out of a positive shock. The magnitude of the shock also matters: for large positive income increases, households are more likely to relax the constraint; consequently, the MPC is lower than for small increases. Note that the analysis of Christelis et al. (2017) focuses on unanticipated income increases, which contrasts with Shea (1995), who examines anticipated income changes (see also Altonji and Siow, 1987).

By incorporating income uncertainty and a broad HARA-class utility into a standard intertemporal optimization problem, Carroll and Kimball (1996) prove that the consumption function is concave in wealth (cash-on-hand): the marginal propensity to consume (MPC) out of wealth or transitory income declines as wealth rises. This contrasts with certainty-equivalence or perfect-certainty benchmarks—such as quadratic utility—in which the consumption function is linear, and the MPC is constant and independent of wealth. With prudence $u'''_{(c)} > 0$, income uncertainty induces precautionary saving: households reduce current consumption at a given wealth level to self-insure against future risk. This mechanism is consistent with Keynes's conjecture that the MPC falls with income, and it underpins buffer-stock saving behavior, naturally summarized by measures of absolute and relative prudence.

Carroll et al. (2019) stress the importance of incorporating nonlinearities to more effectively capture the complexities of consumer behavior under income uncertainty and credit constraints. The authors develop a model explaining consumption through three primary channels: the wealth effect, credit availability, and precautionary motives. Their framework accounts for the impact of non-financial risk—specifically, the risk of permanent unemployment—on intertemporal choice. In this context, once a worker becomes unemployed, re-employment is assumed to be impossible. The central intuition of the model is that, in the presence of income uncertainty, optimizing households target a specific ratio of wealth to permanent income. This target depends on factors such as time preference, prudence, the degree of labor income uncertainty, and credit availability. The model builds upon the approach of Carroll and Toche (2009), which assumes a constant relative risk aversion (CRRA) utility function. This assumption enables a closed-form solution for the target level of wealth, defined as the point at which prudence sufficiently counterbalances impatience.[8]

Shefrin and Thaler (1988) introduced the behavioral life-cycle model, which enhances the traditional life-cycle framework by incorporating three key elements: self-control problems, mental accounting, and framing. Regarding self-control, the theory posits that individuals often struggle with "akrasia," or weakness of will. Shefrin and Thaler conceptualize each person as divided between a "foresight planner," responsible for long-term decision-making, and a "shortsighted doer," who succumbs to immediate temptations. The planner can implement strategies—such as enrolling in savings plans with automatic

[8] With respect to uncertainty, three empirical studies for Colombia are noteworthy; these are Muñoz (2004), Bande, Riveiro, and Ruiz (2021), and Arango et al. (2024). The former concludes that income is the primary driver of household spending and that income uncertainty affects consumption more than the standard life-cycle model predicts. Bande et al. (2021) finds that income uncertainty increases consumption, but when households have savings capacity, the result aligns with the precautionary saving hypothesis, as uncertainty reduces consumption. The capacity to save is, to some extent, endogenous to households' financial conditions. Finally, the latter, emphasizes the distinct contribution of consumer confidence as an uncertainty measure that is separate from income-related uncertainty.

withdrawals—to help regulate the doer's impulsive tendencies. In terms of mental accounting, the model distinguishes between three types of financial resources: current income, current assets, and future or expected income. Current income is the most tempting to spend, current assets (accumulated savings and wealth) are less tempting, and future income is the least likely to be spent. Unlike the traditional view that treats all money as fungible, this model suggests that individuals mentally allocate funds to these distinct accounts, leading to nonfungibility. As a result, people are more likely to spend from their current income than to dip into their savings. For instance, an unexpected bonus may be perceived as "easy money" and spent quickly, in contrast to regular income. Finally, the concept of framing highlights how the presentation of financial information can influence decision-making. For example, a wage increase framed as a periodic rise is more likely to be spent, whereas the same increase presented as an annual bonus is more likely to be saved. This demonstrates how cognitive framing can significantly affect saving and spending behaviors.

Baugh, Ben-David, Park, and Parker (2021) employ the behavioral life-cycle model to explain the observed increase in household consumption following the receipt of anticipated tax refunds. They note that households behave as if they are liquidity constrained when spending these refunds yet tend to smooth consumption in years when they are required to make tax payments. Notably, the authors find that even households with fewer liquidity constraints exhibit similar consumption-smoothing behavior both when making payments and when spending refunds. This asymmetric response to positive and negative cash flows aligns with the mental accounting framework of Shefrin and Thaler (1988), which posits that different forms of wealth—namely, current income, current assets, and future income—are treated as nonfungible mental accounts as we mentioned above. This behavioral perspective suggests that, even in the absence of formal credit constraints, individuals allocate resources differently depending on the source and timing of funds, a finding supported by related research (see also Graham and Isaac, 2001).

In line with behavioral interpretations of asymmetric consumption responses, Coşkun, Apergis, and Coşkun (2022) employ LSTAR and STAR models to differentiate between expansionary and contractionary regimes. Their findings indicate that income is the most significant determinant of household consumption, exhibiting both time-varying and asymmetric effects. Notably, the impact of income on consumption is more pronounced during expansionary periods, particularly in Greece, Italy, Israel, the United States, Australia, and the United Kingdom. Additionally, the study reveals that household consumption in Italy, the United States, Australia, Canada, and Finland respond more strongly to declining interest rate regimes, whereas Sweden and Greece display heightened sensitivity during periods of rising interest rates. The authors further investigate the asymmetric influences of financial and housing wealth on consumption, extending the analysis of prior research (see also Apergis and Miller, 2006).

Asymmetrical responses of consumption to positive versus negative changes in real interest rates have also been the subject of study. In their analysis of consumer credit, Juster and Shay (1964) show that the credit demand of rationed (constrained) borrowers is less interest-rate elastic than that of unrationed borrowers. Unconstrained borrowers—who typically hold relatively high savings and liquid assets—report substantially lower subjective time-preference rates to additional borrowing than constrained customers and, consequently, are unwilling to pay high interest rates on new credit card debt (see Arango and Quevedo-Rocha, 2024; Arango and Cardona-Sosa, 2023).[9] Gourinchas and Rey (2018) further argue that the global consumption-to-wealth ratio and global real interest rates do not systematically comove; moreover, their predictive relation runs from the consumption-to-wealth ratio to future real interest rates, whereas our focus is the reverse channel—from real interest rates to consumption.

For Colombia, Arango et al. (2024) present evidence against the traditional life-cycle/permanent-income hypothesis (LC/PIH): monthly, quarterly, and annual fluctuations in household consumption are predictable from contemporaneous movements in current income, remittances, credit-market conditions, interest rates, and the consumer confidence index (used as an inverse proxy for uncertainty). Building on this result, we examine whether household consumption responds nonlinearly —i.e., asymmetrically— to positive versus negative changes in these drivers and discuss mechanisms that could account for such asymmetries. We also study the long run comovement among consumption, income, interest rates, and related variables. It is important to remark that, unlike the studies cited above, our data does not allow us to determine whether (transitory) changes in these variables are anticipated by consumers or not; instead, we observe permanent expansions and contractions of variables such current income that we exploit to test for asymmetric consumption responses or long run multipliers.

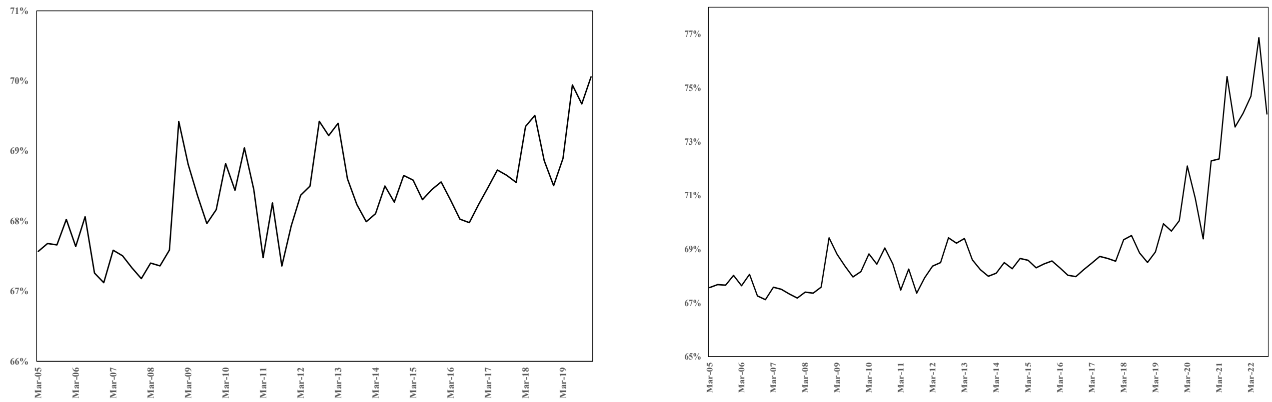
3. Household consumption behavior

From 2005 to 2022, household consumption in Colombia accounted, on average, for just under 70 percent of GDP. As Figure 1 shows, this share remained relatively stable through 2019; beginning around 2016, however, it edged upward and then accelerated after 2019, reaching roughly 76 percent.

[9] GRECO (2002, chap. 5, pp. 135 ff.) presents empirical evidence for 1954–1995 of liquidity constraints and their impact on the consumption of non-durable goods. See also on this topic: López-Mejía (1994) and López-Mejía and Ortega (1998).

Figure 1. Household consumption as a share of GDP, quarterly data

2005: Q1 – 2019: Q4 | **2005: Q1 – 2022: Q3**

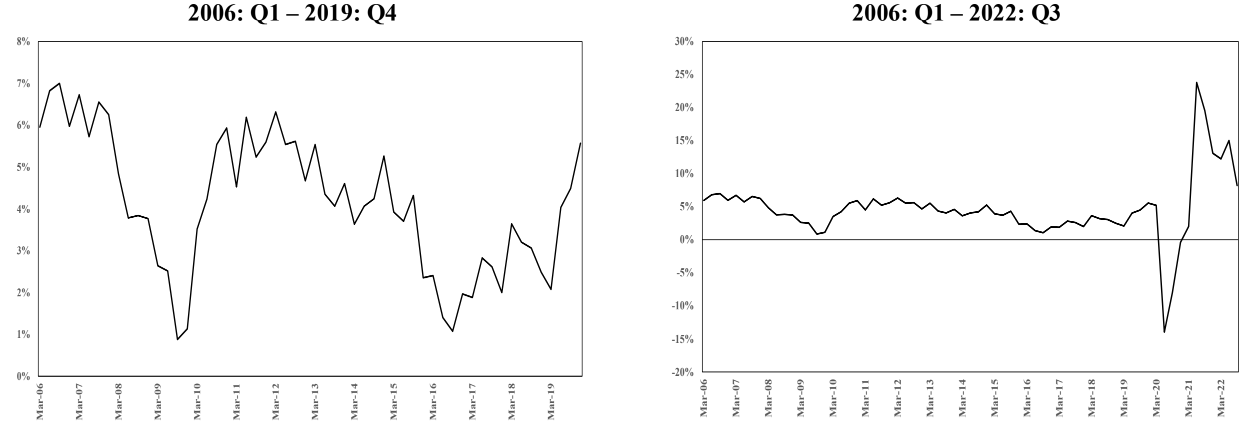


Source: DANE– national accounts. Chained volume series with reference year 2015. Data adjusted for seasonal and calendar effects. Authors' calculations.

Figure 2 plots annual growth in household consumption at a quarterly frequency for two samples: March 2005–December 2019 and March 2005–September 2022. Over 2005Q1–2019Q4, the average annual growth rate of consumption was 4.1 percent; over the full period including the COVID-19 pandemic episode (through 2022Q3), it averaged 4.6 percent. For comparison, GDP grew at average annual rates of 3.9 percent and 4.4 percent over the same windows, respectively.

Figure 2. Annual variation of household consumption

2006: Q1 – 2019: Q4 | **2006: Q1 – 2022: Q3**



Source: DANE – national accounts. Chained volume series with reference year 2015. Data adjusted for seasonal and calendar effects. Authors' calculations.

However, our analysis of the determinants of household consumption does not rely on the national accounts' consumption series. Following Arango et al. (2024a, 2024b), we instead use the monthly retail trade index compiled by the National Administrative Department of Statistics (DANE, by its Spanish acronym), which provides higher-frequency coverage and richer information content. This choice enables us to exploit monthly variation and details that are not observable in the national accounts.

This study investigates the short- and long-run determinants of household consumption in Colombia using monthly data from April 2003 to June 2022 (N = 230). Following Arango et al. (2024a, 2024b), our baseline empirical specification is as follows:

$$c = \alpha_1 + \alpha_y y + \alpha_{rem} rem + \alpha_{ccr} ccr + \alpha_r r + \alpha_{cci} cci + \varepsilon_t, \quad (1)$$

where c denotes the log of the retail trade index excluding fuels and vehicles—our proxy for real household consumption; y denotes current income (proxied by the Economy Situation Indicator, ESI); rem represents remittance inflows to Colombian households;[10] ccr is consumer credit; r corresponds to the consumer-credit interest rates; and cci denotes the consumer confidence index, which we use as an inverse proxy for uncertainty and as a measure of consumer sentiment (Carroll, Fuhrer, and Wilcox, 1994; Ludvigson, 2004).

Introducing regressors beyond current income in equation (1) already implies a departure from the LC/PIH, insofar as it allows for additional determinants of aggregate household consumption (see Arango et al., 2024a). We qualify this claim in Section 5. For now, Figures 3 and 4 plot the variables included in equation (1) over 2003–2022. Figure 3 reports log levels of the retail trade index excluding fuels and vehicles (our proxy for real household consumption), the Economic Situation Indicator (ESI, our proxy for current income), and real remittance inflows. We follow DANE's (2002) definition of retail trade as the resale (without transformation) of merchandise intended exclusively for personal or household use. For the purposes of DANE's retail trade statistics, transactions in pawn shops, telemarketing, lottery outlets, mobile stands, and home-based businesses are excluded, as are maintenance and repair services and sales of used goods. Throughout, we use the retail trade index excluding fuels and vehicles.[11]

The ESI corresponds to the series published by DANE, with 2005 as the base year, expressed in real, seasonally adjusted terms at a monthly frequency from January 2003 onward. Remittances refer to the monthly series published by Banco de la República in nominal U.S. dollars; we convert them to Colombian pesos using the representative exchange rate, also published by the Central Bank, and deflate

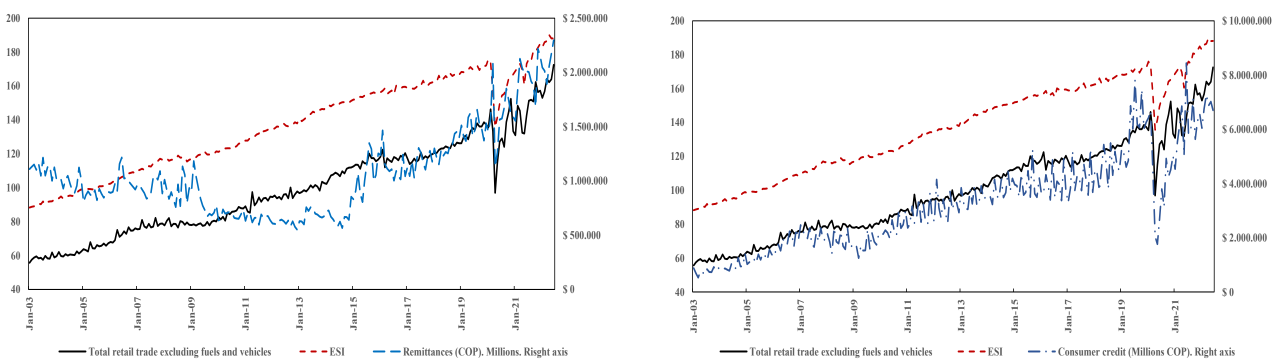
[10] Banco de México (2021) indicates that the pandemic-era contraction in private consumption would have been more severe had remittance inflows not continued to rise in 2020.

[11] We use monthly data from surveys conducted by Colombia's National Administrative Department of Statistics (DANE). Specifically, we draw on the Monthly Retail Trade Sample (MMCM; Spanish acronym), launched in 1989 and redesigned in 2003, from which comparable records are available. In 2013, DANE again redesigned the instrument, renaming it the Monthly Retail Trade Survey (EMCM). The most recent redesign occurred in 2019, when the survey became the Monthly Trade Survey (EMC), with the principal change being the inclusion of both retail and wholesale trade. Our analysis focuses exclusively on retail trade. DANE provides two consistent spans for retail sales indices: (i) January 2003–December 2019, covering MMCM and EMCM; and (ii) January 2013–June 2022. To extend the MMCM-based series through 2022, we splice (chain-link) it by using EMC growth rates.

them with the consumer price index (CPI). Panel A of Figure 3 shows that, following the COVID-19 pandemic, household consumption converged toward the ISE series owing to its faster growth, a pattern also evident in Figure 1, Panel B. Remittances, by contrast, after trending downward through 2015, began to rise and have since moved in the same direction as consumption. Panel B of Figure 3, in addition to consumption and the ESI, plots total real consumer credit, which we obtain from Banco de la República, seasonally adjusted, and deflated using the CPI.

Figure 3. Household consumption, Economic Situation Indicator (ESI), remittances and consumer credit. Monthly data

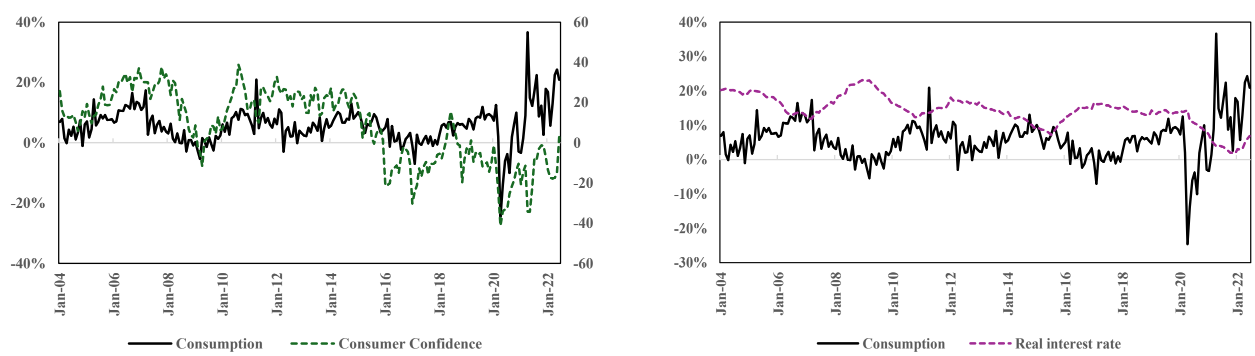
A. Consumption, ESI and remittances **B. Consumption, ESI and consumer credit**



Fuente: Dane; Banco de la República. Authors' calculations.

Figure 4. Real interest rate and annual variations of household consumption and consumer confidence index. Monthly data

A. Consumption and consumer confidence **A. Consumption and real interest rate**



Fuente: Dane; Fedesarrollo; Banco de la República; Authors' calculations.

Panels A and B of Figure 4 plot annual changes in household consumption, the Consumer Confidence Index (cci), and the real consumer-credit interest rate. The cci, published monthly by Fedesarrollo since 2001, comprises the Index of Consumer Expectations (ice) and the Index of Economic Conditions (iec), both constructed from responses to 14 survey questions. We construct the real consumer-credit interest rate from the nominal series published by Banco de la República, applying the ex-ante Fisher relation under a perfect-foresight assumption (i.e., expected inflation equals realized inflation).[12] Visual inspection indicates pronounced long-run comovement among household consumption, the Economic Situation Indicator (ESI), consumer credit, and—toward the end of the sample—remittances, consistent with Panels A and B of Figure 3. Moreover, the annual growth rate of consumption appears closely associated with both the CCI and the real interest rate (Figure 4, Panels A and B).

4. The Autoregressive Distributed Lag (ARDL) and Nonlinear Autoregressive Distributed Lag (NARDL) empirical models

The autoregressive distributed lag (ARDL) framework of Pesaran, Shin, and Smith (2001) is widely used to analyze both short-run dynamics and long-run relationships among time-series variables. A central advantage of the ARDL methodology is its ability to accommodate regressors with mixed orders of integration—stationary (I(0)) and nonstationary (I(1))—provided that no series is integrated of order two or higher (I(2)). The specification allows different lag lengths across variables and admits a conditional error-correction representation, facilitating inference on short-run adjustments and on the existence of a long-run equilibrium via the bounds-testing procedure. These features make ARDL particularly appealing in empirical settings with limited samples and variables of heterogeneous persistence.

As discussed earlier, several studies emphasize incorporating nonlinearities to capture consumer behavior under income uncertainty and credit constraints. While the ARDL framework is widely used to estimate long-run relationships, its linear specification may obscure asymmetric responses. To address this, Shin, Yu, and Greenwood-Nimmo (2013) propose the nonlinear ARDL (NARDL) model, which extends ARDL to allow both short-run and long-run asymmetries. The method applies a partial-sum decomposition of each regressor into positive and negative changes, permitting increases and decreases to exert distinct effects on the dependent variable and enabling dynamic-multiplier analysis of asymmetric adjustment paths. Relative to other nonlinear approaches—such as smooth transition autoregressive (STAR) models and threshold vector error-correction models (TVECMs)—NARDL is parsimonious, straightforward to estimate, and typically less prone to convergence difficulties. It retains the practical advantages of ARDL, including accommodation of mixed orders of integration (I(0)/I(1)), a bounds-testing strategy for level relationships, and an error-correction representation, provided no

[12] We apply the same procedure to the statutory usury rate published by Colombia's Financial Superintendence (Superintendencia Financiera de Colombia, SFC), which we use in several regressions. This rate is the legal ceiling on remunerative (ordinary) and default interest that financial institutions may charge. It is calculated as 1.5 times the average current bank lending rate for consumer credit.

series is I(2). Empirical evidence indicates robust performance in small samples (see Shin, Yu, and Greenwood-Nimmo, 2014; Jareño et al., 2020; Sharma et al., 2023; Uğurlu-Yıldırım et al., 2021).[13]

We employ the nonlinear ARDL (NARDL) framework (Shin, Yu, and Greenwood-Nimmo, 2013) to examine short-run and long-run asymmetric responses of household consumption in Colombia to positive and negative changes in the covariates discussed above. We start by specifying the model:

$$c_t = \sum_{j=1}^{J} \beta_{c,j} c_{t-j} + \sum_{k=1}^{K} \sum_{l=0}^{L} \beta_{k,l} x_{k,t-l} + \sum_{m=0}^{M} \beta_m z_{m,t-1} + \varepsilon_t \tag{2}$$

which has a well-known conditional error correction representation:

$$\Delta c_t = -\rho c_{t-1} + \sum_{k=0}^{K} \lambda_k x_{k,t-1} + \sum_{j=1}^{J} \varphi_{c,j} \Delta c_{t-j} + \sum_{k=1}^{K} \sum_{l=0}^{L} \varphi_{k,l} \Delta x_{k,t-l} + \sum_{m=0}^{M} \beta_m z_{m,t-1} + \varepsilon_t \tag{3}$$

where x comprises y, ccr, rem, and, sometimes, r while z includes cci, and, sometimes, r.

If we further define $EC_t = c_{t-1} - \sum_{k=0}^{K} \frac{\lambda_k}{\rho} x_{k,t-1}$, then the error correction form will be given by the familiar equation:

$$\Delta c_t = -\rho EC_t + \sum_{j=1}^{J} \varphi_{c,j} \Delta c_{t-j} + \sum_{k=1}^{K} \sum_{l=0}^{L} \varphi_{k,l} \Delta x_{k,t-l} + \sum_{m=0}^{M} \beta_m z_{m,t-1} + \varepsilon_t \tag{4}$$

If, as we mentioned above, the empirical model considers the asymmetric effects of y, rem, or r on consumption, following standard practice, some selected regressors are decomposed into positive and negative partial sums:

$$\Delta c_t = -\rho c_{t-1} + \sum_{k=0}^{K} \lambda_k^+ x_{k,t-1}^+ + \sum_{k=0}^{K} \lambda_k^- x_{k,t-1}^- + \sum_{j=1}^{J} \varphi_{c,j} \Delta c_{t-j} + \sum_{k=1}^{K} \sum_{l=0}^{L} \varphi_{k,l} \Delta x_{k,t-l} + \sum_{m=0}^{M} \beta_m z_{m,t-1} + \varepsilon_t \tag{5}$$

where,

$$x_{k,t}^+ = \sum_{i=1}^{t} \Delta x_{k,t}^+ = \sum_{i=1}^{t} max\left(\Delta x_{k,t}, 0\right) \tag{6}$$

$$x_{k,t}^- = \sum_{i=1}^{t} \Delta x_{k,t}^- = \sum_{i=1}^{t} min\left(\Delta x_{k,t}, 0\right) \tag{7}$$

If, as before in the linear model, we further define de asymmetric error correction term as $AEC_t = c_{t-1} - \sum_{k=0}^{K} \left(\frac{\lambda_k^+}{\rho} x_{k,t-1}^+ + \frac{\lambda_k^-}{\rho} x_{k,t-1}^- \right)$, then we can write the NARDL model as:

$$\Delta c_t = -\rho AEC_t + \sum_{j=1}^{J} \varphi_{c,j} \Delta c_{t-j} + \sum_{k=1}^{K} \sum_{l=0}^{L} \varphi_{k,l} \Delta x_{k,t-l} + \sum_{m=0}^{M} \beta_m z_{m,t-1} + \varepsilon_t \tag{8}$$

[13] Galindo and Steiner (2022) employ a nonlinear ARDL (NARDL) model to analyze pass-through from the monetary policy rate to retail interest rates in Colombia, documenting significant asymmetries in adjustments to policy rate increases and decreases.

Short-run asymmetric effects can also be incorporated, leading to the following specification:

$$\Delta c_t = -\rho AEC_t + \sum_{j=1}^{J} \varphi_{c,j} \Delta c_{t-j} + \sum_{k=1}^{K} \sum_{l=0}^{L} \left(\varphi_{k,l}^{+} \Delta x_{k,t-l}^{+} + \varphi_{k,l}^{-} \Delta x_{k,t-l}^{-}\right) + \sum_{m=0}^{M} \beta_m z_{m,t-1} + \varepsilon_t \qquad (9)$$

According to Shin, Yu, and Greenwood-Nimmo (2013), the choice of an appropriate lag structure in equation (9) will render the model free from residual serial correlation and account perfectly for weak endogeneity of any nonstationary explanatory variables.[14] The OLS estimators of the long-run parameters computed as $\hat{\lambda}_k^{+}/\hat{\rho}$ and $\hat{\lambda}_k^{-}/\hat{\rho}$ are T-consistent and follow the mixture normal distribution while all the short-run dynamic parameters in (9) are $\sqrt{T}$-consistent and have the asymptotic normal distribution. The null hypotheses of a symmetric long-run relationship or symmetric short-run coefficients can be tested using the Wald statistic following an asymptotic χ^2 distribution.[15]

Equations (8) and (9) set out the empirical specifications used to test for asymmetric responses of household consumption to positive and negative changes in y, rem, and r, with each driver analyzed in a separate model. Table A1 in the appendix reports unit-root test results for the variables used in the estimations. Because most series are integrated of order one, I(1), a NARDL specification is appropriate for modeling both long-run relationships and short-run dynamics, provided that no variable is integrated of order two.

5. Results and discussion

We begin by estimating symmetric error-correction models (ECMs) as specified in Equation (4). For the consumer-credit interest rate, we consider two samples—2003:01–2019:12 and 2003:01–2022:06—to enable a pre-/post-pandemic comparison. For the usury real interest rate, the estimation covers 2004:01–2022:06. The contrast between the first two periods is intended to gauge the effect of the COVID-19 pandemic on consumption behavior. Arango et al. (2024) report that, in the short-run dynamics of retail trade (excluding fuels and vehicles), the coefficient on the Consumer Confidence Index increased after the pandemic; however, as shown below, this pattern does not persist in our consumption estimates.

As shown in Equation (3), we identify y, rem, and ccr as variables that enter the cointegrating relationship with household consumption, c. Table 1 reports symmetric error-correction specifications estimated within the ARDL framework, which serve as benchmarks for the asymmetric NARDL models analyzed later. Column (1)—estimated through 2019:12 to avoid the effects of the COVID-19

[14] Moreover, for Shin, Yu, and Greenwood-Nimmo (2013) NARDL shares desirable attributes with fully modified estimation (Phillips and Hansen, 1990) and the dynamic corrections in ARDL (Pesaran and Shin, 1998) within a parametric framework capable of modeling asymmetric adjustment.

[15] This follows from Theorems 3.1 and 3.2 in Pesaran and Shin (1998), and requires that assumption (2) in Shin, Yu, and Greenwood-Nimmo (2013), holds.

pandemic—includes the interest rate, r, and the Consumer Confidence Index, cci, as additional covariates. The cci proxies consumer sentiment and is expected to be positively associated with consumption, insofar as higher confidence is typically accompanied by higher spending.

Table 1. Symmetric models of household consumption.

Coefficient	(1)	(2)	(3)	(4)†	(5)†
Speed of adjustment, ρ	-0.4832***	-0.4539***	-0.4893***	-0.2020***	-0.2296***
	(0.0717)	(0.0727)	(0.0874)	(0.0329)	(0.0344)
Number of cointegrating variables	3	3	4	3	4
Long run coefficients, $-\lambda_k/\rho$					
y_{t-1}	1.1230***	0.9632***	0.9513***	1.1849***	1.1408***
	(0.1020)	(0.1179)	(0.1076)	(0.2928)	(0.2511)
rem_{t-1}	0.0657***	0.0953***	0.0953***	0.1251***	0.1196***
	(0.0120)	(0.0168)	(0.0152)	(0.0466)	(0.0397)
ccr_{t-1}	0.0438	0.0838**	0.0879**	0.1035	0.1070
	(0.0330)	(0.0394)	(0.0358)	(0.0911)	(0.0786)
r_{t-1}			-0.8962***		-1.0730***
			(0.1170)		(0.3486)
Short run coefficients					
Number of lags of symmetric coefficients					
Δc_{t-i}	6	3	3	12	11
Δy_{t-i}	1	12	10	8	8
Δrem_{t-i}	1	11	11	7	10
Δccr_{t-i}	11	12	12	12	11
Δr_{t-i}	NA	NA	0	NA	0
Other variables (fixed)					
cci_{t-2}	0.0278***	0.0097	0.0092	0.0432***	0.0378***
	(0.0079)	(0.0086)	(0.0184)	(0.0103)	(0.0093)
r_{t-1}	-0.2252***	-0.4135***		-0.2694***	
	(0.0387)	(0.0698)		(0.0485)	
Statistics					
Sample period	**2003:07 –2019:12**	**2003:05–2022:06**	**2003:05–2022:06**	**2004:02–2022:06**	**2004:01–2022:06**
Observations	198	230	230	221	222
Adjusted R^2	0.6953	0.7787	0.7804	0.7824	0.7825
***F*-statistic**	23.4727	22.7751	20.8482	20.7768	20.3960
***F*-statistic (prob)**	0.000	0.000	0.0000	0.000	0.0000
Durbin–Watson	2.000	1.9704	1.9562	1.8878	1.9988
Bounds tests - F-statistic	8.8883	7.6355	7.7547	7.3578	7.2118
Serial correlation order six					
***F*-statistic (prob)**	0.0567	0.6551	0.7809	0.1228	0.1970
χ^2 (prob)	0.0314	0.5258	0.6782	0.0510	0.0916
Jarque-Bera (prob)	0.0059	0.0007	0.0017	0.0009	0.0003

Notes: standard errors in parenthesis; *, **, and * denote significance at 99%, 95% and 90% respectively. † The model in Column (3) includes usury instead of consumer credit interest rate.**
Source: Authors' calculations.

In the symmetric specification, the bounds test indicates a long-run relationship among consumption, c, current income, y, real consumer credit, ccr, and remittances, rem. Coefficients display the expected signs and are statistically significant in both the long-run relation and the short-run dynamics. The model attains an adjusted R^2 of 0.695 and a Durbin–Watson statistic of 2.00, consistent with the absence of first-order residual autocorrelation. However, the estimated long-run coefficient on current income (y) is 1.123, exceeding unity and implying a more-than-proportional long-run response of consumption to income. This pattern persists in columns (4) and (5), where the usury real interest rate replaces the consumer-credit real interest rate.

Models (1), (2), and (4) include the interest rate to capture the intertemporal-substitution channel in household consumption. Model (2) mirrors Model (1) but extends the estimation sample through 2022:06. In this extended sample, the estimated long-run propensity to consume is close to unity; this, however, does not by itself validate the Life-Cycle Permanent-Income Hypothesis (LC-PIH), because the specification uses current rather than permanent income. Interpretation should also account for the roles of other covariates in the long-run relation and the short-run dynamics. As reported in Table—Models (3) and (5)—multiple lags of Δc, Δy, Δrem, Δccr, and Δr enter significantly in explaining monthly changes in consumption (Δc), consistent with the specification in Model (4).

Drawing on the time-series properties reported in Table A1 (appendix) and the potential wealth effects of the interest rate, Model (3) in Table 1 includes the real interest rate, r, in the cointegrating relationship.[16] The long run response of consumption to the interest rate (-0.8962) is statistically significant and exhibits the expected sign; current income, remittances, and real consumer credit are likewise significant and report the expected signs. Conversely, the Consumer Confidence Index is not statistically significant at conventional levels. Model (3) explains 78 percent of the variation in monthly household consumption (R^2 = 0.78). For comparison, Arango et al. (2024) report adjusted R^2 values up to 0.73 for models of annual changes. Because these specifications use different transformations of the dependent variable, the fit statistics are not strictly comparable; nevertheless, Model (3) exhibits a strong in-sample fit. Moreover, the overall F-statistic is significant at conventional levels, the Durbin–Watson statistic is close to 2, and the ARDL bounds test supports a long-run relationship, all of which indicate a well-specified model. According to the F and χ^2 tests the error processes of both models in Table 1 do not exhibit order-six autocorrelation but the null hypothesis of normally distributed errors is rejected at usual levels according to Jarque-Bera test.

Building on Section 2, we examine the maintained assumption of symmetry in the effects of key determinants on household consumption. Specifically, we ask whether, in the long run, consumption responds differently to positive versus negative changes in the drivers of consumption. The canonical LC-PIH implies symmetry; however, this benchmark may not hold in practice, particularly because our specification uses current rather than permanent income. Moreover, evidence for Colombia indicates that liquidity constraints are salient (e.g., Arango and Cardona-Sosa, 2023; Arango and Quevedo-Rocha,

[16] We allow the interest rate to either enter the cointegrating vector or, alternatively, be restricted to the short-run dynamics in order to probe distinct transmission channels: (i) the intertemporal-substitution (interest-rate) channel, typically identified in short-run dynamics, which changes the relative price of current versus future consumption and affects debt-service costs; (ii) the credit channel, whereby borrowing capacity and loan supply/terms shape households' ability to smooth consumption, often with asymmetric effects; and (iii) the wealth channel, whereby changes in expected returns revalue assets and hence net worth (the net present value of wealth).

2024), a feature that can generate asymmetric adjustment. Accordingly, we now test for asymmetries in consumption long run responses.

We consider current income as a key driver that may exert nonlinear effects on consumption. In particular, we allow the consumption response to differ between permanent expansions and contractions in y. As discussed in Section 2, such asymmetries can signal market frictions and characteristics of consumers. Plausible mechanisms include liquidity constraints affecting a subset of households,[17] buffer-stock savings, and precautionary saving behavior under income risk, and reference-dependent preferences such as loss version. Consistent with this rationale, we implement asymmetry by decomposing income into positive and negative partial sums within the NARDL framework.

We estimate the two asymmetric specifications in Table 2 over 2003:01–2019:12 to characterize pre-pandemic determinants of household consumption. Column (1) allows for asymmetric long-run responses to current income: the estimated long-run response on positive income changes is 0.781, whereas the corresponding coefficient on negative changes is 0.3697. A Wald test of long-run symmetry ($H_0: \lambda_y^+/\rho = \lambda_y^-/\rho = \lambda_y/\rho$) yields a p-value of 0.0012, leading us to reject symmetry. This pattern suggests that households allocate a larger share of permanent income gains to consumption while adjusting more cautiously when income permanently falls.

Our estimates indicate asymmetric long-run adjustment: consumption responds more to income gains than to income losses. Rather than reflecting "prudence during expansions," this pattern would be consistent with mechanisms such as buffer-stock behavior or precautionary saving under income risk. The role of uncertainty is nuanced—Bande et al. (2021) document that in Colombia higher uncertainty coincides with higher consumption—implying that precautionary saving is not the sole channel. A plausible reconciliation, following Shea (1995) for transitory anticipated income changes, is that borrowing constraints bind asymmetrically: households cannot advance consumption ahead of expected income gains (making consumption more sensitive when gains materialize) but can smooth anticipated declines through prior saving (dampening the response to contractions). We therefore interpret the larger pass-through of positive income changes—and the smaller response to negative changes—as evidence of asymmetric adjustment rather than uniform prudence in expansions.

[17] We emphasize that liquidity-constrained consumers likely constitute only a subset of households. Although Colombia's financial sector extends consumer credit, the amounts supplied may fall short of households' desired borrowing (Iregui and Melo, 2009).

Table 2. Asymmetric models of household consumption 2003:01 – 2019:12

Coefficient	(1)	(2)
Speed of adjustment, ρ	-0.5459***	-0.6516***
	(0.0646)	(0.0879)
Number of cointegrating variables	3	4
Long run coefficients $-\lambda_k/\rho$		
y_{t-1}		0.4912**
		(0.1945)
y^+_{t-1}	0.7810***	
	(0.0984)	
y^-_{t-1}	0.3697*	
	(0.2223)	
rem_t	0.0443***	
	(0.0113)	
rem_{t-1}		0.0491***
		(0.0097)
ccr_{t-1}	0.0874***	0.0964***
	(0.0250)	(0.0272)
r^+_{t-1}		-0.2257
		(0.1400)
r^-_{t-1}		-0.9802***
		(0.1805)
Short run coefficients		
Number of lags of symmetric coefficients		
Δc_{t-i}	6	3
Δy_{t-i}	1	1
Δrem_{t-i}	1	1
Δccr_{t-i}	4	11
Other variables (fixed)		
cci_{t-2}	0.0196**	0.0230***
	(0.0080)	(0.0070)
r_{t-1}	-0.3856***	NA
	(0.0736)	
Statistics		
Sample period	**2003:07-2019:12**	**2003:04-2019:12**
Observations	198	201
Adjusted R^2	0.6800	0.6876
***F*-statistic**	35.8883	21.9615
***F*-statistic (prob)**	0.0000	0.0000
Durbin–Watson	2.0348	1.9873
Bounds tests - F-statistic*	11.5885	7.5812
Serial correlation order six		
***F*-statistic (prob)**	0.3883	0.8664
χ^2 (prob)	0.3223	0.8122
Jarque-Bera (prob)	0.0000	0.0000
***F*-statistic (p-value) of long-run symmetry**		
y	0.0012	
r		0.0072

Notes: standard errors in parenthesis; *, **, * denote significance at 90%, 95% and 99% respectively, respectively. Source: Authors' calculations.**

When income falls, households may smooth consumption through international remittances and—where borrowing constraints are not binding—by drawing on consumer credit. In our estimates, both remittances and consumer credit are statistically significant with the anticipated signs. Consumer confidence and the interest rate are likewise significant, with the former positively associated with consumption and the latter negatively so. We include the Consumer Confidence Index to capture residual household sentiment and uncertainty not fully explained by income, credit, or remittances.

Motivated by evidence of credit-market frictions in Colombia (Arango and Cardona-Sosa, 2023; Arango and Quevedo-Rocha, 2024), we test for long run asymmetric responses of household consumption to positive and negative permanent movements in the interest rate. These studies indicate that credit-financed spending is less responsive to rate increases than to decreases. Accordingly, Model (2) in Table 2 introduces nonlinear interest-rate effects by including the positive and negative partial sums of r in the cointegrating relationship, rather than restricting r to the short-run dynamics. The resulting long-run vector comprises four determinants— y, rem, ccr, and r^+/r^- —alongside c. The NARDL bounds test yields a statistic of 7.5812, supporting the existence of a long-run relationship at conventional significance levels.

Consistent with theory, the estimated long-run coefficients on the positive and negative partial sums of the interest rate are both negative. Consumption is substantially more responsive to interest-rate decreases (−0.9802) than to increases (−0.2257), with the latter not statistically significant. Given that $r-$ is defined as the cumulative sum of negative changes, a negative λ_r^- implies that rate cuts raise consumption permanently, and the larger magnitude indicates a stronger response to decreases than to increases. A Wald test of long-run symmetry ($H_0\colon \lambda_r^+/\rho = \ \lambda_r^-/\rho = \lambda_r/\rho$) yields a p-value of 0.0072, leading us to reject symmetry and confirming asymmetric interest-rate effects on household consumption.

In Model (2), the estimated long-run multiplier on income is 0.4912, which—despite being statistically significant—is far below unity and contrasts with the coefficient exceeding one in the symmetric specification of Model (1) (Table 1), indicating sensitivity of the income effect to model specification. Although Model (2) in Table 2 slightly improves in-sample fit relative to Model (1), neither specification matches the adjusted R^2 of 0.6953 reported for the symmetric ARDL model in Table 1.[18]

Building on the pre-pandemic analysis, we extend the sample through 2022:06 and estimate asymmetric NARDL models; results are reported in Table 3. Columns (1)–(2) allow for asymmetric responses to current income, column (3) to remittances, and columns (4)–(5) to the real interest rate, with column (5) using the real usury interest rate. In each specification, the ARDL bounds test supports the presence of a cointegrating relationship among the included variables.

[18] According to the F and χ^2 tests the error processes of both models in Table 2 do not exhibit order-six autocorrelation but the null hypothesis of normally distributed errors is rejected at usual levels according to Jarque-Bera test.

Table 3. Asymmetric models of household consumption 2003- 2022

Coefficient	(1)	(2)	(3)	(4)	(5)†
Speed of adjustment, ρ	-0.6598***	-0.6466***	-0.5611***	-0.6405***	-0.4112***
	(0.0768)	(0.0846)	(0.0827)	(0.0851)	(0.0927)
Number of cointegrating variables	3	3	3	4	4
Long run coefficients $-\lambda_k/\rho$					
y_{t-1}			0.7017***	0.6007***	0.7289***
			(0.1248)	(0.1369)	(0.1952)
y^+_{t-1}	0.8929***	0.8940***			
	(0.0780)	(0.0804)			
y^-_{t-1}	0.7306***	0.7315***			
	(0.0874)	(0.0884)			
rem_{t-1}	0.0619***	0.0631***		0.0724***	0.0918***
	(0.0116)	(0.0115)		(0.0130)	(0.0193)
rem^+_{t-1}			0.0827***		
			(0.0134)		
rem^-_{t-1}			0.0631***		
			(0.0159)		
ccr_{t-1}	0.0924***	0.0912***	0.1007***	0.1110**	0.1227***
	(0.0260)	(0.0270)	(0.0317)	(0.0278)	(0.0445)
r^+_{t-1}				-0.3581*	-0.3885**
				(0.1838)	(0.1890)
r^-_{t-1}				-0.8837***	-0.7394***
				(0.0876)	(0.1239)
Short run coefficients					
Number of lags of symmetric coefficients					
Δc_{t-i}	3	3	3	3	3
Δy_{t-i}	8	10	10	10	9
Δrem_{t-i}	7	1	10	11	10
Δccr_{t-1}	12	12	12	12	12
Δr_{t-i}	NA	NA	NA	1	1
Lag number of asymmetric coefficients					
Δrem_{t-i}	0	1			
Asymmetric coefficients					
Δrem^+_t		0.0146			
		(0.0220)			
Δrem^-_t		0.0805***			
		(0.0274)			
Δrem^+_{t-1}		-0.0402*			
		(0.0234)			
Δrem^-_{t-1}		-0.0147			
		(0.0271)			
Other variables (fixed)					
cci_{t-2}	0.02416***	0.0219**	0.0234**	0.0256***	0.0184
	(0.0083)	(0.0088)	(0.0090)	(0.0085)	(0.0201)
r_{t-1}	-0.3344***	-0.3165***	-0.4017***		
	(0.0439)	(0.0480)	(0.0627)		
Sample period	2003:05-2022:06	2003:05-2022:06	2003:05-2022:06	2003:05-2022:06	2004:02-2022:06
Observations	230	230	230	230	221
Adjusted R^2	0.7912	0.7855	0.7857	0.7937	0.7741
***F*-statistic**	28.9893	28.9528	24.3295	25.4716	19.3867
***F*-statistic (prob)**	0.0000	0.0000	0.0000	0.0000	0.0000
Durbin–Watson	2.0515	2.0022	1.9654	1.9237	1.9416
Bounds tests - F-statistic*	11.9779	9.4897	7.4693	7.8402	4.6321
Serial correlation order six					
***F*-statistic (prob)**	0.2895	0.6671	0.7118	0.7947	0.7685
χ^2 (prob)	0.1866	0.5599	0.5923	0.6931	0.6569
Jarque-Bera (prob)	0.0214	0.0008	0.0061	0.0197	0.0063
F-statistic (*p*-value) of long-run symmetry					
y	0.0000	0.0000			
rem			0.0086		
r				0.0094	0.0866
F-statistic (*p*-value) of short-run symmetry: sum					
Δrem		0.1086			
F-statistic (*p*-value) of short-run symmetry: lag					
Δrem		0.2211			

Notes: standard errors in parenthesis; *, **, and * denote significance at 99%, 95% and 90% respectively. † The model in Column (3) includes usury instead of consumer credit interest rate. Source: Authors' calculations.**

Based on Wald tests (F-statistics) reported across the models, we reject the null of long-run symmetry in the coefficients (p-values < conventional thresholds). Columns (1) and (4) yield the highest adjusted R^2 values, with column (1) also displaying the larger overall F-statistic. By contrast, the Wald test provides no evidence of short-run asymmetry for remittances (column 3), whereas long-run asymmetry in income is statistically significant.

Column (1) of Table 3 shows that the long-run response of household consumption to positive changes in current income is 0.8929, exceeding the propensity to negative changes (0.7306). This asymmetry aligns with mechanisms such as precautionary saving and liquidity constraints. Remittances, consumer credit, and consumer confidence (the latter treated as exogenous to the cointegrating relationship) are positively associated with consumption, whereas increases in the real interest rate—also exogenous—reduce consumption.

The NARDL bounds test supports a cointegrating relationship in Model (3) of Table 3. Also estimates nonlinear long-run effects of remittances on consumption by allowing asymmetric responses to positive and negative remittance changes: Wald test rejects long-run symmetry at conventional levels. The long-run coefficient on positive changes $(\lambda_{rem}^{+}/\rho = 0.0827)$ exceeds that on negative changes $(\lambda_{rem}^{-}/\rho = 0.0631)$, mirroring the pattern observed for long-run effects of current income. Consumer confidence is positively associated with consumption, and the real interest rate enters with the expected negative sign, with both covariates statistically significant. Nevertheless, this specification does not achieve the highest adjusted R^2 among the alternatives.

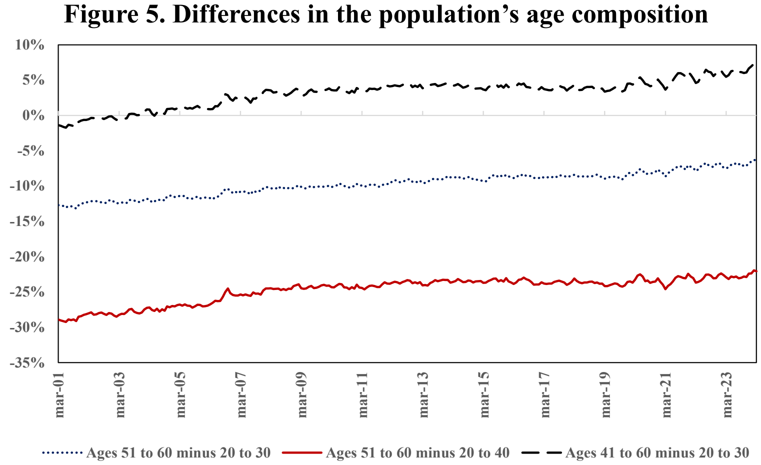
Columns (1) and (4) exhibit strong in-sample fit and well-behaved diagnostics but feature different sources of nonlinearity: column (1) allows asymmetric long-run effects of current income, whereas column (4) introduces asymmetry through the interest rate by including r^{+} and r^{-} in the cointegrating relationship (as in Table 2, column (2)). The long-run coefficient on positive rate changes is −0.3581, smaller in magnitude than the coefficient on negative changes (−0.8837), indicating that consumption responds more strongly to rate cuts than to hikes—consistent with liquidity-constraint mechanisms discussed before (Juster and Shay, 1964). In this specification, the interest-rate channel appears to operate in tandem with the credit channel. As in Table 2, column (2), the estimated long-run coefficient on current income (0.6007) is relatively small compared with earlier estimates. The model in column (5), which uses the real usury interest rate, yields a similar asymmetry suggestive of liquidity constraints, although its adjusted R^2 is not the highest in Table 3. Note that in this table, as in the two previous tables, although the errors do not exhibit autocorrelation of order six, the hypothesis of normally distributed errors is rejected.

5.1 Population aging and Government transfers

The NARDL framework allows us to model the long-run adjustment dynamics of key determinants of household consumption—current income, remittances, consumer credit, and the interest rate. We also find evidence of asymmetric (nonlinear) consumption responses to positive and negative changes in our proxies for current income and the interest rate. Consistent with Arango et al. (2024), the consumer confidence index —our proxy for consumer sentiment—is an important exogenous short-run driver of household consumption. Nevertheless, there remains scope to examine whether demographic variables can be incorporated into this nonlinear framework, allowing population age structure to help explain consumer behavior or not as predicted by life-cycle model. Accordingly, we investigate whether age affects consumption decisions, particularly in light of the recent demographic shift toward population aging in Colombia.

To that end, we partition the population aged 20–60 into four 10-year cohorts and compute each cohort's share of the total population: 20–30, 31–40, 41–50, and 51–60. Figure 5 plots differences in shares between older and younger groups: the black line plots the share of those aged 51–60 minus the share of those aged 20–30; the green line plots the share of those aged 51–60 minus the combined share of those aged 20–40; and the red line plots the share of those aged 41–60 minus the share of those aged 20–30. All series trend upward, indicating population aging: the relative share of older cohorts has increased over time.

Figure 5. Differences in the population's age composition



Source: Dane; Banco de la República; Authors' calculations

In the Table 4 models, the age-structure variables—defined as differences in cohort shares as in Figure 5—enter as exogenous (short-run) regressors outside the cointegrating space. All three demographic variables are statistically significant, and their negative coefficients indicate that a higher share of older cohorts, relative to the 20–30 and 20–40 groups, is associated with lower household consumption. Table 4 also reports results that allow for nonlinear consumption responses to income and the consumer-credit interest rate. Consistent with columns (4) and (5) of Table 3, the consumer-credit real interest rate provides greater explanatory power than the usury real interest rate.

Table 4. Asymmetric models of household consumption 2003- 2022

Coefficient	**(1)**	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**
	Asymmetric responses to ISE			**Asymmetric responses to interest rate**		
Speed of adjustment, ρ	-0.6811***	-0.7227***	-0.7196***	-0.6405***	-0.6511***	-0.6571***
	(0.0868)	(0.0900)	(0.0929)	(0.0929)	(0.0898)	(0.0911)
Number of cointegrating variables	3	3	3	4	4	4
y^+_{t-1}	0.8882***	0.9646***	0.9125***			
	(0.0756)	(0.0797)	(0.0722)			
y^-_{t-1}	0.7085***	0.7823***	0.7449***			
	(0.0868)	0.0029	(0.0804.0)			
y_{t-1}				0.6006***	0.5784***	0.5686***
				(0.1434)	(0.1320)	(0.1304)
ccr_{t-1}	0.1024***	0.0902***	0.1002***	0.1111***	0.1159***	0.1216***
	(0.0266))	(0.0237)	(0.0249)	(0.0302)	(0.0274)	(0.0282)
rem_{t-1}	0.0535***	0.0564***	0.0529***	0.0724***	0.0675***	0.0637***
	(0.0133))	(0.0111)	(0.0124)	(0.0150)	(0.0130)	(0.0142)
r^+_{t-1}				-0.3580*	-0.3163*	-0.3107*
				(0.1892)	(0.1770)	(0.1753)
r^-_{t-1}				-0.8837***	-0.9111***	-0.8885***
				(0.0904)	(0.0960)	(0.0858)
Number of lags of symmetric coefficients						
Δc_{t-i}	3	3	3	3	3	3
Δy_{t-i}	8	9	9	10	10	10
Δrem_{t-i}	7	7	7	11	10	10
Δccr_{t-i}	12	12	12	12	12	12
Δr_{t-i}				0	0	0
Other variables (fixed)						
cci_{t-2}	0.0240***	0.0265***	0.0269***	0.0256	0.0263***	0.02661***
	(0.0085)	(0.0084)	(0.0086)	(0.0192)	(0.0470)	(0.0088)
r_{t-1}	-0.3342***	-0.3837***	-0.3627***			
	(0.0438)	(0.0603)	(0.0529)			
$dif\_tpop4160-2030$	-0.3185***			-0.0004		
	(0.0999)			(0.2701)		
$dif\_tpop5160-2030$		-0.9183***			-0.3490***	
		(0.0603)			(0.0470)	
$dif\_tpop5160-2040$			-0.4713***			-0.2689***
			(0.0550)			(0.0359)
Sample period	**2003:05 – 2022:06**	**2003:05 – 2022:06**	**2003:05 – 2022:06**	**2003:05 – 2022:06**	**2003:05 – 2022:06**	**2003:05 – 2022:06**
Observations	230	230	230	230	230	230
Adjusted R^2	0.7914	0.7936	0.7930	0.7859	0.7930	0.7934
***F*-statistic**	28.1559	27.6887	27.5813	20.5516	25.3765	25.4282
***F*-statistic (prob)**	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000
Durbin–Watson	2.0504	1.9613	1.9865	1.9236	1.9093	1.9206
Bounds tests - F-statistic*	9.9940	10.4733	9.7426	6.5731	7.2702	7.1986
Serial correlation order six						
***F*-statistic (prob)**	0.2210	0.3240	0.2872	0.7965	0.7576	0.7401
χ^2 (prob)	0.1305	0.2095	0.1793	0.6927	0.6455	0.6237
Jarque-Bera (prob)	0.0402	0.0450	0.0416	0.0197	0.0365	0.0439
F-statistic (*p*-value) of long-run symmetry						
y	0.0000	0.0000	0.0000			
r				0.0154	0.0041	0.0036

Notes: standard errors in parenthesis; *, **, * represents significance at 10%, 5% and 1%, respectively. Source: Authors' calculations.**

Columns (1)–(3) of Table 4 document long run asymmetric responses of household consumption to positive and negative permanent changes in current income, whereas columns (4)–(6) report nonlinear coefficients for positive and negative fluctuations in the consumer-credit interest rate. We cannot jointly estimate asymmetries in both income and the interest rate with this dataset: when both are modeled simultaneously, the bounds/Wald test fails to reject long-run symmetry with respect to the interest rate. Based on the diagnostics, the model in column (2) performs well (adjusted R^2, F-statistic, and bounds test). According to the diagnostic tests reported in the lower panel of Table 4, the residuals of models (1)–(3) show no evidence of autocorrelation through order six, and there is some evidence that they are approximately normally distributed. The Jarque–Bera tests yield the highest p-values observed so far, which supports the normality assumption of Shin et al. (2013).

It features a sizable speed of adjustment to the long-run equilibrium (−0.7227) and indicates that consumption responds more to income increases (0.9646) than to decreases (0.7823). Remittances and consumer credit, which enter the cointegrating vector, are statistically significant with the expected signs. Outside the cointegrating relationship, consumer confidence, the interest rate, and the age-structure variable—defined as the difference between the population shares aged 51–60 and 20–30—are statistically significant. This pattern is inconsistent with the strict life-cycle hypothesis, which predicts smooth consumption over time under perfect capital markets. Instead, the estimates indicate lower consumption among older cohorts relative to younger ones, consistent with income declines at retirement in a country, like Colombia, where persistent labor informality limits pension coverage to a small fraction of the population. Moreover, the asymmetries reported in Table 4 are consistent with capital-market frictions. Accordingly, once liquidity constraints are taken seriously, the sign and statistical significance of the age-structure coefficients are to be expected.

Columns (4)–(6) estimate asymmetric consumption responses to positive and negative changes in the consumer-credit interest rate. The bounds test indicates cointegration in all three specifications. Consumption reacts less (in absolute value) to rate increases than to decreases, a pattern consistent with liquidity constraints. These models share several features. First, the demographic variables are either statistically insignificant (e.g., column 4—consistent with the life-cycle model) or smaller in magnitude than in columns (1)–(3). Second, the speeds of adjustment back to the long-run equilibrium are similar across the three models. Third, the current income coefficient is relatively low (up to 0.6006), which reduces comparability with the models in columns (1)–(3) and with prior estimates. Given the latter characteristic, in what follows we include the interest rate symmetrically in the cointegrating vector and allow long run nonlinear consumption responses only with respect to current income.

Among the three models in columns (1)–(3) of Table 5, the specification in column (2) exhibits the strongest diagnostics. Within the cointegrating (long-run) relationship, all variables—including consumer confidence and the age-structure difference (the 51–60 share minus the 20–30 share)—are statistically significant with the expected signs. Columns (4)–(6) add the lagged annual percentage change in government transfers to households, as exogenous, sometimes interpreted as support for household consumption. When significant, its coefficient is small. In those cases, the speed of adjustment slows materially (see columns (5) and (6)); at the same time, the long-run coefficients on consumer credit and remittances increase relative to columns (1)–(3), while the interest-rate coefficient—now included in the cointegrating vector—declines.

Table 5. Asymmetric models of household consumption including transfers

Coefficient	**(1)**	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**
Speed of adjustment, ρ	-0.7620***	-0.7757***	-0.7632***	-0.7592***	-0.4952***	-0.4991***
	(0.0893)	(0.0851)	(0.0856)	(0.1043)	(0.0617)	(0.0636)
Number of cointegrating variables	4	4	4			
y^{+}_{t-1}	0.9115***	0.9782***	0.9230***	0.7639***	0.9830***	0.9387***
	(0.0672)	(0.0732)	(0.0671)	(0.0832)	(0.1388)	(0.1280)
y^{-}_{t-1}	0.7226***	0.7832***	0.7462***	0.6348***	0.7829***	0.7554***
	(0.0772)	(0.0759)	(0.0747)	(0.0879)	(0.1326)	(0.1326)
ccr_{t-1}	0.0878***	0.0796***	0.0887***	0.1527***	0.1012***	0.1051***
	(0.0240)	(0.0222)	(0.0237)	(0.0300)	(0.0380)	(0.0395)
rem_{t-1}	0.0480***	0.0488***	0.0477***	0.0639***	0.0619***	0.0630***
	(0.0124)	(0.0107)	(0.0118)	(0.0138)	(0.0176)	(0.0195)
r_{t-1}	-0.5873***	-0.6209***	-0.5939***	-0.4644***	-0.4988***	-0.4855***
	(0.0983)	(0.0954)	(0.0971)	(0.1019)	(0.1609)	(0.1581)
Number of lags of symmetric coefficients						
Δc_{t-i}	3	3	3	3	9	9
Δy_{t-i}	9	9	9	8	8	8
Δrem_{t-i}	7	7	7	7	2	3
Δccr_{t-i}	12	12	12	12	2	2
Δr_{t-i}	5	5	5	0	0	0
Other variables (fixed)						
cci_{t-2}	0.0273***	0.0259***		0.0225***	0.0470***	0.0468***
	(0.0078)	(0.0081)		(0.0195)	(0.0103)	(0.0102)
$dif\_tpop4160-2030$	-0.3346***			0.3283		
	(0.0735)			(0.4077)		
$dif\_tpop5160-2030$		-1.0005***			-0.4860***	
		(0.1065)			(0.0545)	
$dif\_tpop5160-2040$			-0.0263***			-0.1553***
			(0.0080)			(0.0174)
$\Delta trans_{t-1}$				0.0084	0.0114*	0.0113*
				(0.0062)	(0.0063)	(0.0062)
Sample period	**2003:05 – 2022:06**	**2003:05 – 2022:06**	**2003:05 – 2022:06**	**2007:03 – 2022:06**	**2007:03 – 2022:06**	**2007:03 – 2022:06**
Observations	230	230	230	184	184	184
Adjusted R^2	0.8015	0.8039	0.8022	0.8090	0.7942	0.7937
***F*-statistic**	25.9927	26.3803	26.1112	21.4101	31.7121	31.6074
***F*-statistic (prob)**	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000
Durbin–Watson	2.0102	1.9899	2.0143	2.0112	1.7585	1.7671
Bounds tests - F-statistic*	10.0777	11.4994	10.9869	9.0565	8.8650	8.4630
Serial correlation order six						
***F*-statistic (prob)**	0.1676	0.1543	0.1589	0.7201	0.0126	0.0145
χ^2 (prob)	0.0816	0.0733	0.0761	0.5777	0.0044	0.0052
Jarque-Bera (prob)	0.1117	0.1346	0.1299	0.2192	0.1231	0.1359
F-statistic (*p*-value) of long-run symmetry						
y	0.0000	0.0000	0.0000	0.0026	0.0016	0.0016

Notes: standard errors in parenthesis; *, **, * represents significance at 10%, 5% and 1%, respectively. Source: Authors' calculations.**

Given the magnitudes of the estimated coefficients on positive and negative changes in current income, the specification reported in Table 5, column (2), merits an interpretation in terms of asymmetric consumption responses.[19] Were the analysis based on permanent rather than current income, a rejection of long-run symmetry would contradict the strict LC-PIH. However, because the regressors capture current income this concern does not arise. The coefficient on current income increases is close to one, implying near one-for-one pass-through when current income rises. The result is also difficult to reconcile with standard buffer-stock and precautionary-saving models, which typically imply some smoothing of both positive and negative fluctuations. By contrast, the asymmetric pattern—where the estimated coefficient is larger for income increases than for decreases—suggests credit frictions or hand-to-mouth behavior on the upside, alongside partial smoothing of negative fluctuations, and is broadly consistent with evidence reported by Bande et al. (2021). Overall, the smaller coefficient when income declines relative to when it increases indicates asymmetric adjustment consistent with liquidity constraints and related frictions in household behavior.[20]

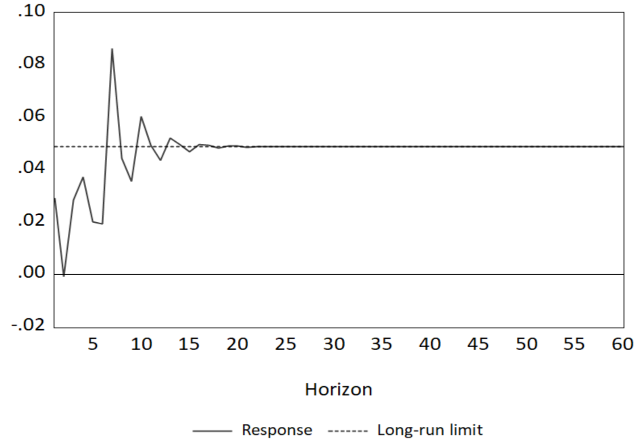
Figure 6 reports the cumulative dynamic multipliers for our preferred specification (Table 5, column 2). Following a change, the system converges to its long-run equilibrium within roughly 20 months, consistent with the estimated speed of adjustment (Panels A–D). Panel B displays the consumption responses to positive and negative fluctuations to current income; the differential response (positive minus negative) converges to 0.195 and is statistically significant by the 60-month horizon. The pass-through from income increases to consumption is approximately 25 percent larger than the pass-through from income decreases. Stability diagnostics in Figure 7 indicate no structural change in the intercept (CUSUM, Panel A); this is consistent with mean parameter stability. The CUSUMSQ test suggests that the error variance is stable over most of the sample, with a possible episode of persistent variance instability between 2015 and 2020 in the column (2) specification, after which stability resumes. This brief excursion beyond the bounds is consistent with transient variance instability.

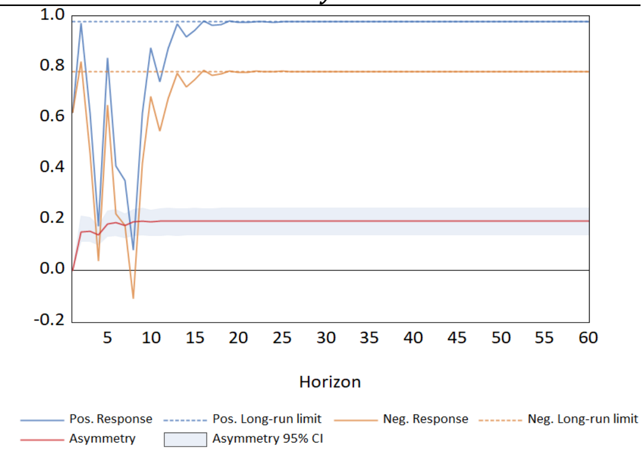
[19] Based on the diagnostic tests in the lower panel of Table 5, the residuals for model (2) show no evidence of autocorrelation up to lag six. The Jarque–Bera test fails to reject normality, and the remaining diagnostics do not raise concerns, indicating an overall well-specified model.

[20] An alternative explanation for nonlinear consumption responses—best examined in a framework with durable and nondurable goods—is as follows. When current income rises and borrowing constraints ease, the share of expenditure allocated to durables increases. Conversely, when income declines and constrained households face tighter credit supply, spending on durables contracts sharply and the durables share of total expenditure falls relative to nondurables. In this mechanism, saving is predominantly precautionary: in normal times, households save against the risk of a future downturn; in recessions, they save in anticipation of further deterioration. During booms, credit supply expands, inducing dissaving (a drawdown of net financial assets) and a pronounced rise in durables spending.

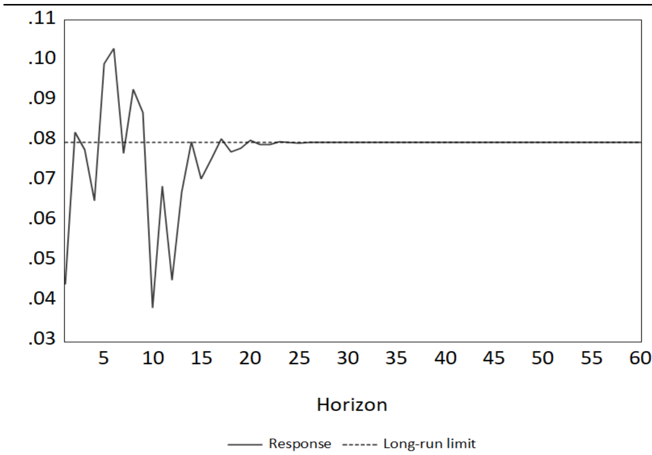
Figure 6. Model in Column (2) of Table 5: cumulative dynamic multiplier

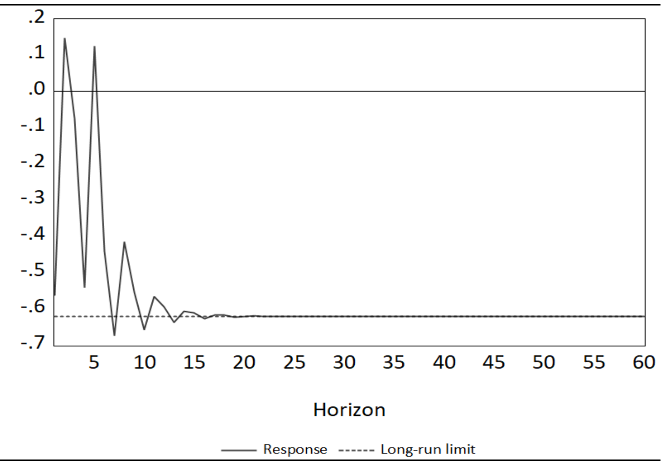
A. $\boldsymbol{rem}$	**B.** $\boldsymbol{y}$





C. $\boldsymbol{ccr}$	**D.** $\boldsymbol{r}$

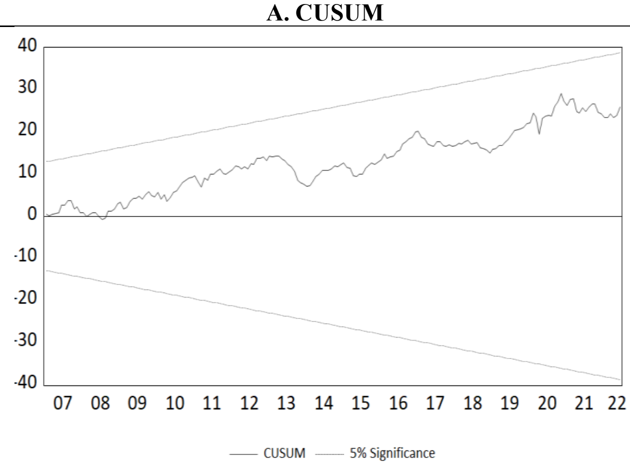


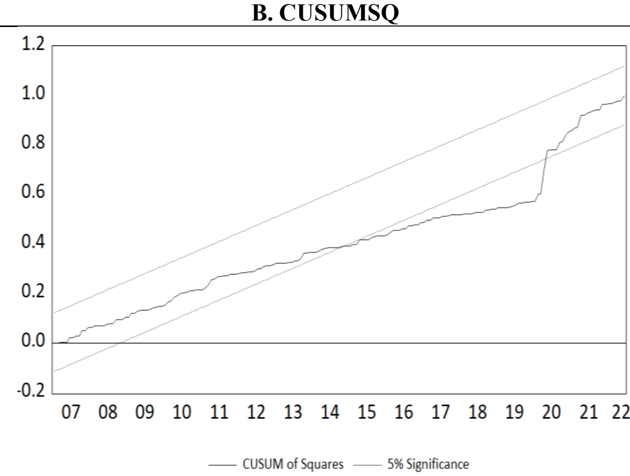


Source: Authors' calculations.

Figure 7. CUSUM and CUSUM squares of models of Table 5

A. CUSUM	**B. CUSUMSQ**





Note: 5% confidence interval. Source: Authors' calculations.

6. Conclusions

This article provides evidence on asymmetries in Colombian household consumption for 2003–2022 using nonlinear cointegration methods. Estimating both linear ARDL and nonlinear ARDL (NARDL) models, we document pronounced asymmetric adjustments in consumption dynamics, proxied by the retail sales index excluding motor vehicles and fuels. Using monthly data and following Arango et al. (2024), we find that current income (proxied by ESI), remittances, consumer credit, the interest rate, consumer confidence, and several measures of population aging jointly explain up to 80 percent of the variation in consumption.

In first place, we find that symmetric ARDL models for 2003–2022, exhibit satisfactory fit by conventional goodness-of-fit criteria. Nevertheless, specifications that allow consumption to respond asymmetrically to changes in current income and the interest rate outperform the symmetric benchmarks.

According to the results, interest-rate cuts elicit long-run consumption responses nearly 188 percent larger than interest rate increases [see Model (5) in Table 4]. We interpret the observed asymmetry in consumption responses as consistent with binding liquidity constraints and with recent evidence that the interest-rate channel operates alongside the credit channel (Arango and Cardona-Sosa, 2023; Arango and Quevedo-Rocha, 2024). In particular, Arango and Cardona-Sosa (2023), show that the credit demand of rationed borrowers is less responsive to interest rates than that of unconstrained borrowers, reflecting quantity rationing. By contrast, unconstrained households—who typically hold higher savings and liquid assets and thus have substantially lower subjective yields —are less willing to accept high rates on new consumption credit, implying a stronger interest-rate sensitivity.

We also find that, in the long run, consumption's response to positive income fluctuations is about 25 percent larger (in absolute value) than its response to declines. This pattern points to credit-market frictions: households devote a larger share of income gains to consumption while moderating spending cuts when income falls, consistent with evidence on liquidity-constrained households.

Although asymmetries in long-run consumption responses to permanent changes in current income and the interest rate are both plausibly attributable to liquidity constraints, the income-driven asymmetry provides a better fit—by conventional model-comparison metrics—than the interest-rate asymmetry.

Remittances and consumer credit also emerge as significant long-run determinants of household consumption. These results add to the standard consumption literature by underscoring the role of cross-border transfers. The presence of remittances in the cointegrating relationship alongside current income suggests that external transfers are incorporated into Colombian households' accounts—

consistent with features of a remittance-dependent, migration-linked economy. In addition, consumer credit, especially the portion corresponding to payroll-deduction loans[21], has gained much importance since 2008.

The estimated sign of the consumer confidence index and the significance of its coefficient are consistent with the view that higher confidence proxies lower perceived uncertainty or consumer sentiment. However, the uncertainty captured by the confidence index is distinct from that of income. When current income rises, the long-run consumption response is close to unity, suggesting that income uncertainty does not attenuate spending out of permanent positive income changes.

An additional contribution of this study is to identify demographic effects in the short-run dynamics of household consumption. Across specifications, population aging—measured by cohort-share differentials—is associated with lower consumption: coefficients on the gap between older (51–60) and younger (20–30) cohort shares are negative and close to −1. This result is salient in light of rapid demographic transition of Colombia and stands in contrast to López, Misas, and Oliveros (1996), who had reported no demographic effects on consumption. The negative association between aging and consumption that we find in this paper is difficult to reconcile with a benchmark life-cycle model under perfect capital markets, which implies that shifts in age composition should not materially affect aggregate consumption conditional on income and interest rates.

The estimates linked to population aging are consistent with income declines at retirement in a country, like Colombia, where persistent labor informality limits pension coverage to a small fraction of the population. Moreover, the asymmetrical long run responses of consumption that we report are consistent with capital-market frictions. Accordingly, if liquidity constraints are taken seriously, the negative sign and statistical significance of the age-structure coefficients should not surprise the analysts. The finding also underscores the need to anticipate headwinds to aggregate demand as population aging advances.

Future research should investigate in greater depth the mechanisms underlying these asymmetries, with particular attention to household heterogeneity and the interaction between demographic change and consumption behavior. Regional analyses within Colombia could also elucidate spatial heterogeneity in consumption dynamics.

[21] In Colombia, a payroll-deduction loan (crédito de libranza) is a form of credit in which installment payments are automatically deducted from the borrower's wages or pension benefits. The employee or pensioner executes an irrevocable authorization (for the duration of the debt) permitting the employer or paying entity to withhold the installment and remit it directly to the creditor.

References

Altonji, J.G., and Siow, A. (1987), Testing the Response of Consumption to Income Changes with (Noisy) Panel Data, The Quarterly Journal of Economics, 102, 293-328.

Apergis, N., and Miller, S. M. (2006). Consumption asymmetry and the stock market: Empirical evidence. *Economics Letters, 93*(3), 337–342. https://doi.org/10.1016/j.econlet.2006.06.002

Arango, L. E., and Cardona-Sosa, L. M. (2023). Consumer credit in an emerging economy: Demand, supply, and liquidity restrictions. *Emerging Markets Review, 54*, 100949. https://doi.org/10.1016/j.ememar.2022.100949

Arango, L. E., and Quevedo, I. K. (2024). Credit constraints in Colombia: Evidence from the use of credit cards among low- and middle-income individuals. *International Journal of Emerging Markets, 19*(10), 2991–3010. https://doi.org/10.1108/IJOEM-08-2021-1241

Arango, L. E., Flórez, L. A., Marín, N. J., and Posada, C. E. (2024). Consumption of households in Colombia: What do the retail trade indices tell us? *Banco de La República, Borrador de Economía*, *1275*.

Arias, F., Lozano, I., Granger, C., Vásquez, D., Vargas, C., and Rodríguez, N. (2023). Dinámica y determinantes del consumo de los hogares en Colombia durante la postpandemia del COVID-19. Revista De Economía Del Rosario, 26(2), 1–36. https://doi.org/10.12804/revistas.urosario.edu.co/economia/a.13515

Attanasio, O.P. (1999). Consumption, in J.B. Taylor J.B. and Woodford M. (Eds) *Handbook of Macroeconomics*, Volume 1, Chapter 11, 742 – 805.

Bande, R., Riveiro, D. and Ruiz, F. (2021): *Does Uncertainty Affect Saving Decisions of Colombian Households? Evidence on Precautionary Saving. MPRA, Paper No. 106771. https://mpra.ub.uni-muenchen.de/106771/*.

Balke, N. S., and Fomby, T. B. (1997). Threshold cointegration. *International Economic Review, 38*(3), 627–645. https://doi.org/10.2307/2527284

Baugh, B., Ben-David, I., Park, H., and Parker, J. A. (2021). Asymmetric consumption smoothing. *American Economic Review*, *111*(1), 192-230.

Boone, L., C. Giorno and P. Richardson (1998), “Stock Market Fluctuations and Consumption Behaviour: Some Recent Evidence”, OECD Economics Department Working Papers, No. 208, OECD Publishing, Paris, https://doi.org/10.1787/853414662323.

Bowman, D., Minehart, D., and Rabin, M. (1993). Loss aversion in a savings model. Journal of Economic Behavior and Organization, 38, 2, 155-178.

Bunn, P., Le Roux, J., Reinold, K., and Surico, P. (2018). The consumption response to positive and negative income shocks. J*ournal of Monetary Economics*, *96*, 1-15.

Campbell, J. Y., and Mankiw, N. G. (1989). *Consumption, income, and interest rates: Reinterpreting the time series evidence* (NBER Working Paper No. 2924). National Bureau of Economic Research. https://doi.org/10.3386/w2924

Carrasquilla, A., (1989). “La Asignación Intertemporal del Consumo en Colombia”, Revista ESPE - Ensayos sobre Política Económica, Banco de la República de Colombia, vol. 8(16), pages 67-82, December.

Carroll, C. D., Fuhrer, J. C., and Wilcox, D. W. (1994). Does consumer sentiment forecast household spending? If so, why? *The American Economic Review*, *84*(5), 1397-1408.

Carroll, C., and Toche, P. (2009). *A Tractable Model of Buffer Stock Saving* (Working Paper 15265). National Bureau of Economic Research. https://doi.org/10.3386/w15265

Carroll, C., Slacalek, J., and Sommer, M. (2019). *Dissecting Saving Dynamics: Measuring Wealth, Precautionary, and Credit Effects*. National Bureau of Economic Research. https://doi.org/10.3386/w26131

Carroll, C. D., and Kimball, M. S. (1996). On the concavity of the consumption function. *Econometrica: Journal of the Econometric Society*, 981-992.

Christelis, D., Georgarakos D., Jappelli T., Pistaferri L., and van Rooij, M. (2017), "AsymmetricConsumption Effects of Transitory Income Shocks," DNB Working Paper, No 551.

Coşkun, Y., Apergis, N., and Coşkun, E. (2022). Nonlinear responses of consumption to wealth, income, and interest rate shocks. *Empirical Economics*, *63*(3), 1293-1335.

Deaton, A. (1992). Understanding Consumption. Clarendon Lectures in Economics, Oxford, Clarendon Press.

Duesenberry, J. S. (1949). *Income, saving, and the theory of consumer behavior*. Harvard University Press.

Enders, W., and Siklos, P. L. (2001). Cointegration and Threshold Adjustment. *Journal of Business and Economic Statistics*, *19*(2), 166–176. http://www.jstor.org/stable/1392161

Engle, R. F., and Granger, C. W. J. (1987). Co-Integration and Error Correction: Representation, Estimation, and Testing. *Econometrica*, *55*(2), 251–276. https://doi.org/10.2307/1913236

Friedman, M. (1957). *A theory of the consumption function*. Princeton University Press.

Galindo, A. J., and Steiner, R. (2022). Asymmetric interest rate transmission in an inflation-targeting framework: The case of Colombia. *Latin American Journal of Central Banking, 3*(3), 100069. https://doi.org/10.1016/j.latcb.2022.100069.

Graham, F., and Isaac, A. G. (2001). The behavioral life-cycle theory of consumer behavior: survey evidence. *Journal of economic behavior and organization*, *48*(4), 391-401.

Gohar, R., Bagadeem, S., Chang, B. H., and Zong, M. (2022). Do the income and price changes affect consumption in the emerging 7 countries? Empirical evidence using quantile ARDL model. *Annals of Financial Economics*, *17*(04). https://doi.org/10.1142/S2010495222500245.

Gourinchas, P. O., and Rey, H. (2019). Global real rates: A secular approach. BIS Working Papers, No 793.

GRECO – Grupo de Estudios del Crecimiento Económico (A. Pontón, C. E. Posada y M. Urrutia). (2002). El crecimiento económico colombiano en el siglo XX. Banco de la República – Fondo de Cultura Económica, Bogotá.

Hall, R. E. (1978). Stochastic implications of the life cycle-permanent income hypothesis: Theory and evidence. *Journal of Political Economy, 86*(6), 971–987. http://www.jstor.org/stable/1840393.

Han, H.-L., and Ogaki, M. (1997). Consumption, income and cointegration. *International Review of Economics and Finance, 6*(2), 107–117. https://doi.org/10.1016/S1059-0560(97)90019-8.

Hernández A., J. N. (2006). Revisión de los determinantes macroeconómicos del consumo total de los hogares para el caso colombiano. *Revista ESPE - Ensayos sobre Política Económica, 24*(52), 80–109. https://doi.org/10.32468/Espe.5202

Iregui, A. M., and Melo, L. (2009). La transmisión de la política monetaria sobre el consumo en presencia de restricciones de liquidez, in Jalil M. and Mahadeva L. (eds), Mecanismos de transmisión de la política monetaria en Colombia, Universidad Externado and Banco de la República, Chapter 5, 247-296.

Jain, S., Banerjee, R., and Sharma, R. W. (2023). Meaning-oriented consumption: A systematic review and research agenda. *International Journal of Consumer Studies, 47*(6), 2305–2334. https://doi.org/10.1111/ijcs.12927.

Japelli, T. and Pistaferri, L. (2017). The Economics of Consumption, Theory and Evidence. Oxford University Press.

Jareño, F., González, M. de la O., Tolentino, M., and Sierra, K. (2020). Bitcoin and gold price returns: A quantile regression and NARDL analysis. *Resources Policy, 67*, 101666. https://doi.org/10.1016/j.resourpol.2020.101666.

Juster, T. and Shay, R. P. (1964). Consumer sensitivity to finance rates, in T. Juster and R. P. Shay (eds), Consumer Sensitivity to Finance Rates: An Empirical and Analytical Investigation. National Bureau of Economic Research, Cambridge, MA, 6–46.

Keynes, J.M. (1936) The General Theory of Employment, Interest, and Money. John Maynard Keynes.

King, R. G., Plosser, C. I., Stock, J. H., and Watson, M. W. (1991). Stochastic trends and economic fluctuations American Economic Review, 81,4, 819-840.

López-Mejía, A. (1994). "La teoría del ingreso permanente y las restricciones de liquidez en Colombia", en *Estabilización y Crecimiento*. Fedesarrollo – Tercer Mundo Editores. Bogotá.

López-Mejía, A., Misas A., M., and Oliveros Camacho, H. (1996). *Understanding consumption in Colombia* (Borradores Semanales de Economía No. 58, pp. 1–29). Banco de la República.

López-Mejía, A. y Ortega, J. R. (1998). "Private Saving in Colombia". IMF Working Paper 171.

Ludvigson, S. C. (2004). Consumer confidence and consumer spending. *Journal of Economic perspectives*, *18*(2), 29-50.

MacDonald, G., Mullineux, A., and Sensarma, R. (2011). Asymmetric effects of interest rate changes: the role of the consumption-wealth channel, *Applied Economics*, 43, 16, pages 1991-2001.

Modigliani, F., and Brumberg, R. (1954). Utility Analysis and the Consumption Function: An Interpretation of Cross-Section Data. In Kenneth K. Kurihara, ed., *Post-Keynesian Economics*, 388 436. New Brunswick, NJ: Rutgers University Press.

Muñoz C., M. (2004). Determinants of household income and expenditures. Revista de Economía Institucional, 6(10), 183–199. https://www.scielo.org.co/scielo.php?script=sci_arttext&pid=S0124-59962004000100008.

Park, J. Y., and Phillips, P. C. B. (2001). Nonlinear regressions with integrated time series. *Econometrica, 69*(1), 117–161. https://doi.org/10.1111/1468-0262.00180.

Pesaran, M. H., and Shin, Y. (1998). An autoregressive distributed-lag modelling approach to cointegration analysis. *Econometrics and Economic Theory in the 20th Century: The Ragnar Frisch Centennial Symposium*, 31, 371–413 http://dx.doi.org/10.1017/CCOL0521633230.011.

Pesaran, M. H., Shin, Y., and Smith, R. J. (2001). Bounds Testing Approaches to the Analysis of Level Relationships. *Journal of Applied Econometrics*, *16*(3), 289–326. http://www.jstor.org/stable/2678547.

Phillips, P. C. B. and Hansen, B.E., 1990, Statistical Inference in Instrumental Variables Regressions with I(1) Processes, Review of Economic Studies, 57, 99-125.

Posada, Carlos E., and A. Gaviria, 1992, El consumo en Colombia: revisión de la evidencia empírica. Coyuntura Económica. Vol. XXII, No. 4, pp. 137-149. Fedesarrollo, Bogotá - Colombia

Shea, J. (1995). Myopia, liquidity constraints, and aggregate consumption: a simple test. *Journal of money, credit and banking*, *27*(3), 798-805.

Shefrin, H. M., and Thaler, R. H. (1988). The behavioral life-cycle hypothesis. *Economic inquiry*, *26*(4), 609-643.

Shirvani, H., and Wilbratte, B. (2000). Does consumption respond more strongly to stock market declines than to increases? *International economic journal*, *14*(3), 41-50.

Shin, Y., Yu, B., and Greenwood-Nimmo, M. (2013). Modelling Asymmetric Cointegration and Dynamic Multipliers in a Nonlinear ARDL Framework. In *Festschrift in Honor of Peter Schmidt* (pp. 281–314). Springer New York. https://doi.org/10.1007/978-1-4899-8008-3_9.

Souleles, N. (1999). The Response of Household Consumption to Income Tax Refunds, *American Economic Review*, 89, 4, 947 – 958.

Ugurlu-Yildirim, E., Kocaarslan, B., and Ordu-Akkaya, B. M. (2021). Monetary Policy Uncertainty, Investor Sentiment, and Us Stock Market Performance: New Evidence from Nonlinear Cointegration Analysis. *International Journal of Finance and Economics, 26, 1724-1738*. https://doi.org/10.1002/ijfe.1874.

Appendix. Unit root tests for the main variables

Based on the comprehensive unit root analysis presented in Table A1, our testing strategy employs three complementary methodologies—ADF, KPSS, and Phillips-Perron tests—to establish robust integration order classifications essential for ARDL and NARDL validity. The results provide strong empirical support for our methodological approach by confirming the theoretically expected integration patterns. Consumption, c, income, y, and remittances, rem, exhibit clear I(1) behavior across all specifications, failing to reject unit root nulls in levels but showing stationarity in first differences. Conversely, interest rates, r, consumer credit growth, Δccr, and consumer confidence, cci, exhibit I(0) properties, with ADF statistics approaching or exceeding critical values in levels and KPSS statistics falling well below rejection thresholds.

Table A1. Unit root tests

Variables	ADF				KPSS				Phillips -Perron			
	Levels		First differences		Levels		First differences		Levels		First differences	
	No trend	With trend	No trend	With trend	No trend	With trend	No trend	With trend	No trend	With trend	No trend	With trend
c	0.1796 (0.9708)	-2.1432 (0.5184)	-6.6518 (0.000)	-6.7473 (0.000)	2.0325	0.1421	0.0583	0.0561	-0.2532 (0.9282)	-7.4974 (0.00)	-26.1966 (0,00)	-26.1488 (0.00)
y (2000:01-2022:06)	-0.4576 (0.8957)	-2.8082 (0.1956)	-8.3817 (0.000)	-8.3658 (0.000)	2.1556	0.3614	0.0699	0.0440	-0.7649 (0.82)	-2.8721 (0.17)	-15.8896 (0.000)	-15.8367 (0.000)
rem (2000:01-2022:066)	-0.7351 (0.83)	-1.1960 (0.9086)	-12.0847 (0.000)	-12.0682 (0.000)	1.0254	0.3445	0.1379	0.1353	-2.4445 (0.1306)	-3.2409 (0.0788)	-27.0751 (0.000)	-28.3037 (0.000)
ccr (2000:05-2022:06)	-2.7722 (0.0640)	-2.7643 (0.2123)	-8.4118 (0.000)	-8.4156 (0.000)	0.1274	0.0427	0.1033	0.0608	-8.8115 (0.000)	-8.8820 (0.000)	-29.4332 (0.000)	-29.3416 (0.000)
cci (2001:11-2022:06)	-2.0707 (0.2568)	-2.9459 (0.1501)	-3.9272 (0.0022)	-3.9054 (0.0133)	0.8895	0.3308	0.1702	0.1047	-2.8683 (0.0506)	-3.5969 (0.0321)	-17.5625 (0.000)	-17.9380 (0.000)
r	-2.0822 (0.2523)	-5.6997 (0.000)	-4.7660 (0.000)	-4.6764 (0.000)	1.1884	0.0805	0.0520	0.0362	-1.4593 (0.5527)	-2.5525 (0.3027)	-12.1326 (0.000)	-12.1143 (0.000)
$usury\ r$ (2004:01-2022:12)	-3.4573 (0.0101)	-3.6438 (0.0285)	-6.2200 (0.000)	-6.2161 (0.000)	0.2042	0.0970	0.0479	0.0435	-2.5764 (0.0994)	-2.6783 (0.2467)	-14.5348 (0.000)	-14.5180 (0.000)

Notes: *p < 0.10, **p < 0.05, ***p < 0.01. Source: Authors' calculations.
Source: Authors' calculations