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Unemployment Persistence without Full Hysteresis: Quarterly Evidence from Sri Lanka

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Abstract

This paper examines the natural-rate hypothesis and the Keynesian persistence mechanism using genuine quarterly macroeconomic data for Sri Lanka. Inflation expectations are constructed recursively with an AR(4) forecast using information available at time t . Inference combines classical OLS, HC3 covariance estimates, Newey-West HAC(4) intervals, and a 1,000-replication circular block bootstrap. Preliminary ADF, KPSS, and Zivot-Andrews tests show that unemployment, real GDP, CPI, and real consumption exhibit substantial persistence in levels, whereas year-on-year GDP growth and inflation are more compatible with stationarity. The direct levels model has weak fit ($R^2 = 0.155$), pronounced serial correlation, and parameter instability; its positive real-income coefficient is not significant under HAC or block-bootstrap inference. A dynamic model performs materially better ($R^2 = 0.541$) and estimates unemployment persistence at 0.705 (HAC 95% CI: 0.508-0.903), while rejecting the unit-persistence restriction $\rho = 1$ ($p=0.0045$). GDP growth has the Keynesian/Okun-consistent negative sign but is statistically imprecise. The consumption results do not support persistent fractional long memory: Engle-Granger cointegration is not supported ($p=0.793$), and the exploratory Hurst estimate for changes in log consumption is 0.395. Overall, the quarterly evidence supports state dependence and medium-run persistence, but not full hysteresis, a stable negative income coefficient in levels, or persistent fractional long memory. The results therefore motivate a more conditional interpretation of Keynesian policy effects in a small open economy.

Keywords: natural rate of unemployment; hysteresis; inflation expectations; Okun relation; quarterly time series; Sri Lanka; robust inference

JEL Classification: E24, E31, E32, C22, O53

1 Introduction

The natural-rate hypothesis remains one of the central organizing ideas in macroeconomics because it separates temporary demand-driven movements in employment from longer-run labor-market equilibrium. Keynes's emphasis on effective demand, nominal rigidities, and involuntary unemployment provides the classic alternative to automatic market clearing [1]. Friedman formalized the natural-rate argument in monetary policy, stressing that attempts to hold unemployment permanently below its equilibrium level would ultimately be absorbed into inflation expectations [2]. Phelps reached a closely related conclusion in an expectations-augmented Phillips-curve framework [3] and subsequently developed a microeconomic labor-market account of wage dynamics and equilibrium unemployment [4]. The empirical question is therefore not whether demand matters in the short run, but how persistent its real effects can be and whether the equilibrium unemployment benchmark itself is invariant to shocks and policy.

The persistence literature weakened the sharpest interpretation of a fixed natural rate. Blanchard and Summers showed that labor-market shocks may alter subsequent wage bargaining and thereby leave long-lived effects on unemployment, giving formal content to hysteresis [5]. Blanchard and Katz later emphasized that economists possess a useful conceptual framework for the natural rate but much less precise empirical knowledge of its determinants [6]. Stiglitz similarly argued that uncertainty about the natural rate and its movement over time should temper excessively rigid policy conclusions [7]. Gordon's time-varying NAIRU framework made this point operational by allowing the inflation-consistent unemployment benchmark to evolve [8], while Staiger, Stock, and Watson documented the substantial uncertainty surrounding conventional NAIRU estimates and their use in inflation forecasting [9]. These contributions imply that persistence, mean reversion, and full hysteresis should be treated as distinct empirical possibilities rather than assumed *ex ante*.

A second strand of research asks whether aggregate demand can influence unemployment beyond the conventional cyclical horizon. Ball's cross-country evidence argues that demand conditions can affect long-run unemployment and therefore challenges a strict separation between cyclical and structural unemployment [10]. Blanchard and Wolfers show that shocks interact with labor-market institutions, so identical disturbances can generate different degrees of persistence across economies [11]. Nickell's comparative evidence likewise cautions against treating labor-market 'rigidity' as a single sufficient explanation for unemployment differences [12]. The broader framework of Layard, Nickell, and Jackman places unemployment outcomes at the intersection of wage setting, price setting, institutions, and macroeconomic demand [13]. For an emerging economy exposed to external financing constraints and repeated shocks, these mechanisms are especially relevant because the observed unemployment path may reflect both structural features and the history of aggregate demand.

The output-unemployment relationship provides a direct empirical bridge between Keynesian demand mechanisms and labor-market outcomes. Okun's original analysis established a systematic association between changes in economic activity and unemployment [14]. A large subsequent literature finds that the strength of this relationship varies across countries and institutional settings; Ball, Leigh, and Loungani show that Okun's law remains empirically useful but that its coefficient is heterogeneous across national labor markets [15]. This motivates testing both a levels relation, which is closest to the paper's theoretical income equation, and a dynamic growth specification that is less vulnerable to common stochastic trends.

Inflation expectations form the second channel in the theoretical model. Modern Phillips-curve work emphasizes forward-looking price setting and the distinction between expected and realized inflation [16]. Forecasting evidence, however, shows that inflation is difficult to predict with stable reduced-form relationships, particularly when the inflation process itself changes [17]. Atkeson and Ohanian demonstrate that simple benchmarks can outperform some NAIRU-based inflation forecasts, underscoring the need to validate rather than assume an expectations proxy [18]. Accordingly, the present study constructs expected inflation recursively from an AR(4) model estimated only on information available at each date and separately checks its relationship with subsequent realized inflation.

The consumption channel is also important because an income expansion affects employment only to the extent that it translates into aggregate expenditure. Hall's stochastic formulation of the life-cycle/permanent-income hypothesis highlights the role of expectations in consumption dynamics [19]. Campbell and Mankiw show that aggregate consumption may reflect a mixture of forward-looking behavior and current-income 'rule-of-thumb' behavior [20], while Deaton demonstrates how liquidity constraints can make consumption more

closely tied to current resources [21]. These results are particularly relevant for crisis-prone emerging economies, where borrowing constraints and losses in real income can strengthen the short-run connection between output, disposable resources, and household expenditure.

Fractional Brownian motion provides a formal stochastic structure in which the Hurst parameter governs long-range dependence [22]. Yet long-memory claims require caution: Lo shows that apparent persistence can disappear once short-range dependence is treated appropriately [23], and Beran's statistical treatment emphasizes that long-memory inference is highly sensitive to sample size and specification [24]. Accordingly, fractional long memory is treated as an empirical proposition rather than a maintained assumption. The analysis estimates an exploratory R/S Hurst statistic on changes in log real consumption and treats $H > 0.5$ as a diagnostic condition to be assessed.

Sri Lanka offers a demanding environment in which to evaluate these mechanisms. The economy entered a severe balance-of-payments and sovereign-debt crisis in 2022, followed by an IMF-supported adjustment and reform program; the 2024 Article IV consultation stresses that stabilization has advanced while macroeconomic vulnerabilities remain substantial [25]. The World Bank's 2024 development update likewise describes a recovery from the deep crisis but continued weakness in household incomes, employment conditions, and labor-force participation [26]. The Central Bank of Sri Lanka's Annual Economic Review 2023 documents the transition from the 2022 contraction toward a gradual rebound as inflation and external pressures eased [27]. Official labor-force statistics report an unemployment rate of 4.7% in 2022 but also show important demographic and participation dimensions of labor-market slack [28]. Recent Central Bank research using quarterly data from 1998 to 2024 confirms the usefulness of macroeconomic variables for unemployment forecasting and highlights the importance of dynamic econometric specifications [29]. The World Bank's 2023 development update further documents the large output losses, job destruction, poverty increase, and severe external constraints associated with the crisis [30]. Together, these sources make Sri Lanka a relevant case for distinguishing persistent but mean-reverting unemployment from full hysteresis and for assessing whether output and inflation expectations contribute independently to unemployment dynamics.

The empirical design has four main features. First, it uses only genuinely observed quarterly series and performs no frequency interpolation. Second, it separates a direct empirical counterpart of the theoretical levels equation from a dynamic specification that includes lagged unemployment and year-on-year real GDP growth. Third, it uses robust inference suited to quarterly macroeconomic data: HC3 covariance estimates, Newey-West HAC(4) intervals, and a circular moving-block bootstrap. Fourth, it treats the sign of the income term, the magnitude of unemployment persistence, the unit-persistence restriction associated with strong hysteresis, the existence of a consumption-income cointegrating relation, and the presence of long memory as testable propositions rather than maintained assumptions. The evidence indicates that unemployment is persistent but the restriction $\rho = 1$ is rejected in the dynamic model; the real-output coefficient has the expected negative sign only in the growth specification and remains statistically imprecise; and consumption exhibits strong short-run comovement with output but neither cointegration nor a Hurst estimate above 0.5. The contribution is therefore a disciplined assessment of the Keynesian persistence mechanism rather than a claim that expansionary policy mechanically eliminates a fixed natural rate.

2 Theoretical framework and testable implications

Let u_t denote the unemployment rate, u^* a benchmark natural-rate concept, $E_t[\pi_{t+1}]$ the expectation formed at time t of next-period inflation, and i_t real aggregate income/output. The baseline Keynesian specification is:

$$u_{t+1} = u^* + f E_t[\pi_{t+1}] + g i_{t+1} + \varepsilon_t \quad (1)$$

In (1), u_{t+1} is the unemployment rate in the next quarter; u^* is the benchmark natural-rate concept; $E_t[\pi_{t+1}]$ is the conditional expectation, formed at time t , of next-period inflation π_{t+1} using information available at that date; i_{t+1} is real aggregate income/output in the next quarter; f is the reduced-form coefficient on expected inflation; g is the reduced-form coefficient on real aggregate income/output; and ε_t is a mean-zero regression disturbance with finite variance. The index t denotes the current quarter. The model is intentionally stylized. Keynesian reasoning suggests that stronger real activity should ordinarily reduce unemployment, implying $g < 0$, whereas inflation expectations may raise measured unemployment in a supply-constrained or wage-adjustment environment if anticipated price pressure worsens real-income and employment conditions. The empirical analysis does not impose either sign and does not impose Gaussian disturbances.

A central identification clarification is necessary. In an unrestricted regression, the estimated constant is a conditional intercept, not automatically an estimate of u^* . Calling the intercept a structural natural rate would require normalization and exclusion restrictions strong enough to separate the equilibrium benchmark from omitted trend, institutional, and external-sector influences. The empirical analysis therefore reports α as an intercept and reserves u^* for the theoretical benchmark.

The policy mechanism expresses real income as an accumulating function of an expansionary impulse A_t :

$$i_{t+1} = (1 + A_t)i_t \quad (2)$$

In (2), A_t denotes the proportional expansionary impulse applied between t and $t+1$, so that $1 + A_t$ is the corresponding gross income-expansion factor; i_t is real aggregate income/output at time t and i_{t+1} is its next-period value. Under the expansionary interpretation used here, A_t is non-negative. Iterating (2) from the current quarter t through the final policy-impulse quarter T yields the cumulative income path through quarter $T+1$. Rather than requiring literal zero unemployment, the framework allows a general policy target for unemployment at $T+1$, which may lie below the benchmark natural rate. A zero-unemployment target is the limiting special case in which the target equals zero. Evaluating (1) at T and combining it with the cumulative income path yields the policy-product identity:

$$\prod_{s=t}^T (1 + A_s) = \frac{u_{T+1}^{\text{target}} - u^* - f E_T[\pi_{T+1}]}{g i_t} \quad (3)$$

In (3), the product is taken over policy periods $s = t, \dots, T$; s is the running time index; T is the final policy-impulse quarter; A_s is the proportional expansionary impulse in period s ; u_{T+1}^{target} is the chosen unemployment target in quarter $T+1$; $E_T[\pi_{T+1}]$ is inflation in quarter $T+1$ as expected using information available at time T ; and i_t is real aggregate income/output at the initial date. The disturbance ε_T from (1) is omitted from the calibration identity, equivalently using its conditional mean of zero. The coefficients f and g retain the meanings defined after (1). The identity is well defined provided $g i_t \neq 0$. Equation (3) is a theoretical calibration identity, not an empirical causal estimate. It indicates the direction and magnitude of income expansion that would be required under the maintained structural model. Whether such a calibration is economically feasible depends on the sign and stability of g , the behavior of expectations, external constraints, and policy endogeneity. The quarterly evidence below therefore tests the reduced-form ingredients before drawing any policy conclusion.

2.1 Dynamic persistence and hysteresis

To distinguish persistence from full hysteresis, the empirical analysis augments the baseline relation with lagged unemployment and uses a stationary activity transformation:

$$u_{t+1} = \alpha + \rho u_t + f E_t[\pi_{t+1}] + \gamma \Delta_4 y_{t+1} + \sum_{q=2}^4 \delta_q D_{q,t} + \varepsilon_t \quad (4)$$

In (4), α is the conditional intercept; ρ is the coefficient on current unemployment u_t ; f retains its meaning from (1); γ is the coefficient on year-on-year real GDP growth; and $\Delta_4 y_{t+1}$ is that growth rate evaluated at $t+1$. For any quarter s , $\Delta_4 y_s \equiv 100(y_s/y_{s-4} - 1)$, where y_s denotes real GDP. $D_{q,t}$ is the seasonal indicator for quarter q at time t : it equals 1 when observation t belongs to quarter q and 0 otherwise. The index q runs over $\{2, 3, 4\}$, so quarter 1 is the omitted reference category; δ_q is the coefficient measuring the seasonal effect of quarter q relative to quarter 1; and the summation from $q = 2$ to 4 collects the complete set of quarterly seasonal controls. Finally, ε_t is the regression disturbance. A value $0 < \rho < 1$ indicates state dependence with eventual mean reversion; ρ close to one indicates strong persistence; and $\rho = 1$ corresponds to unit persistence in the conditional specification. The restriction $H_0: \rho = 1$ is therefore tested directly. Because the sample is moderate and residual dependence may remain, the paper reports HAC and block-bootstrap uncertainty rather than relying on a single conventional t statistic.

2.2 Consumption transmission and the long-memory hypothesis

The theoretical demand channel begins with a simple consumption-income relation:

$$C_t = C_0 + B i_t \quad (5)$$

In (5), C_t denotes real private consumption at time t ; C_0 is the baseline or autonomous consumption component; B is the reduced-form slope linking consumption to real aggregate income/output; and i_t is real aggregate income/output at time t . In this levels relation, B is measured in consumption units per unit of income and is not automatically an elasticity. For empirical work, levels of real GDP and real consumption are strongly trending, so a stable long-run relation must be established rather than assumed. The paper therefore tests log

consumption and log real GDP for Engle-Granger cointegration before interpreting any levels coefficient as a long-run elasticity. Short-run co-movement is then examined using changes in logs.

A stochastic extension can be written schematically as:

$$C_T = C_t + \ell(t) B_{T-t}^H \quad (6)$$

In (6), C_T is real consumption at the terminal horizon T ; C_t is current real consumption; ℓ_t is a time- t scale or loading parameter that converts the stochastic term into consumption units; B^H with horizon index $T-t$ denotes the fractional Brownian-motion term over the interval of length $T-t$; and $H \in (0,1)$ is the Hurst parameter governing dependence. The stochastic process B^H is distinct from the coefficient B in (5). Persistent long memory requires $H > 0.5$. This condition is not imposed empirically. The R/S estimate on $\Delta \log(C_t)$ is treated as exploratory evidence only; failure to obtain $H > 0.5$ means that the fractional-persistence extension is not supported by the present quarterly sample.

2.3 Empirical hypotheses

- H1: Expected inflation has a nonzero association with subsequent unemployment ($f \neq 0$).
- H2: Stronger real activity reduces unemployment ($g < 0$ in a valid levels relation or $\gamma < 0$ in the growth specification).
- H3: Unemployment is persistent ($\rho > 0$), but full hysteresis requires failure to reject $\rho = 1$.
- H4: Real consumption and real GDP share a stable long-run relation only if the cointegration null is rejected in favor of cointegration.

3 Data and empirical strategy

3.1 Data sources, frequency, and actual sample

The executed replication uses genuine quarterly observations and performs no interpolation, forward fill, backward fill, or synthetic frequency conversion. Although the code includes an optional ILOSTAT unemployment branch, the local execution reported in the supplied compilation did not have the pyreadr dependency installed; consequently, the actual unemployment series used in every reported estimate is the IMF International Financial Statistics fallback Q.LK.LUR_PT. Real GDP and real private consumption are non-seasonally adjusted IMF IFS constant-price national-currency series, and the CPI is a directly observed quarterly IFS index. The series were retrieved through the DBnomics API.

Table 1 Data sources and raw quarterly coverage

Variable	Series code	Provider	First	Last	N
Unemployment rate, u_t (%)	Q.LK.LUR_PT	IMF International Financial Statistics via DBnomics	1990Q1	2022Q4	104
Real GDP, Y_t (constant-price LCU)	Q.LK.NGDP_R_NSA_XD C	IMF International Financial Statistics via DBnomics	2010Q1	2025Q1	61
Real private consumption, C_t (constant-price LCU)	Q.LK.NCP_R_NSA_XDC	IMF International Financial Statistics via DBnomics	2010Q1	2025Q1	61
Consumer price index, P_t	Q.LK.PCPI_IX	IMF International Financial Statistics via DBnomics	1959Q1	2025Q2	266

The merged data frame contains 142 calendar quarters from 1990Q1 onward. Missing values remain missing. Joint regression samples are shorter because real GDP/consumption begin in 2010 and unemployment ends in 2022Q4.

The individual-series coverage is therefore much longer than the joint estimation window. Model A uses 50 observations from 2009Q4 to 2022Q3 after aligning u_{t+1} , i_{t+1} , seasonal dummies, and the recursively constructed expectation. Model B uses 45 observations from 2010Q4 to 2022Q3 because year-on-year GDP growth additionally requires four quarters of lagged GDP. This distinction is important: the paper has long historical coverage for unemployment and prices, but the multivariate identification sample is approximately twelve to thirteen years, not three decades.

3.2 Transformations and inflation expectations

Real GDP is scaled by 10^{12} in the levels regression solely for numerical conditioning. Because the underlying IFS series is in national-currency units, the magnitude of the levels coefficient is not economically interpretable without reversing that scaling; the sign and statistical uncertainty are the relevant objects. Year-on-year real GDP growth is computed as $100(y_t/y_{t-4} - 1)$, and year-on-year inflation as $100(CPI_t/CPI_{t-4} - 1)$. Quarter dummies Q2, Q3, and Q4 are included because the national-accounts series are non-seasonally adjusted.

The theoretical term $E_t[\pi_{t+1}]$ is generated recursively. At each quarter t , an AR(4) inflation equation is fitted using only inflation observations available through t , subject to a minimum training window of 24 observations, and a one-quarter-ahead forecast is produced. Thus no realized future inflation enters the expectation. Model A2 provides a transparent robustness check using adaptive expectations, $E_t[\pi_{t+1}] = \pi_t$.

3.3 Preliminary tests and robust inference

Before regression estimation, each series is examined using descriptive statistics and 95% confidence intervals for the sample mean. Stationarity is assessed with the augmented Dickey-Fuller test [31] and KPSS test [32], which use opposite null hypotheses, and with the Zivot-Andrews test allowing an endogenous structural break [33]. This combination is especially important for crisis data, where unit-root tests without breaks can be misleading.

Regression inference is deliberately multi-layered. Classical OLS results are retained for comparability, Newey-West HAC standard errors with four lags address serial dependence and heteroskedasticity in quarterly residuals [34], and HC3 covariance estimates guard against leverage-sensitive heteroskedasticity [35]. To reduce reliance on asymptotic approximations in a moderate sample, 1,000 circular moving-block bootstrap replications with an eight-quarter block are also reported, following the logic of dependent-data resampling [36]. Model diagnostics include Durbin-Watson, Breusch-Godfrey, Breusch-Pagan [38], Ljung-Box [39], CUSUM stability tests based on Brown, Durbin, and Evans [40], Jarque-Bera [41], and Ramsey RESET [42]. In the executed run, RESET statistics were not returned for Models A, A2, and B and are therefore reported as unavailable rather than reconstructed.

3.4 Consumption tests

The consumption module first applies the Engle-Granger procedure [37] to log real consumption and log real GDP. Only if cointegration is supported can the coefficient from the levels equation be interpreted as a stable long-run elasticity. The code also estimates a residual-based dynamic regression in changes; because the cointegration test ultimately fails, this equation is reported for completeness but is not interpreted as a valid error-correction mechanism. Finally, an exploratory R/S Hurst statistic is computed on changes in log real consumption, consistent with the caution in the long-memory literature [22-24].

4 Preliminary time-series evidence

4.1 Descriptive statistics

Table 2 Descriptive statistics for the quarterly series

Series	N	Mean	95% CI for mean	Std. dev.	Minimum	Maximum
Unemployment rate, u_t (%)	104	7.189	[6.454, 7.924]	3.777	3.818	16.586
Real GDP, Y_t	61	2.717×10^6	$[2.549 \times 10^6, 2.886 \times 10^6]$	6.572×10^5	1.415×10^6	3.563×10^6
Real GDP growth, ΔY_t (YoY, %)	57	5.448	[2.336, 8.560]	11.728	-16.975	52.060
Consumer price index, P_t	142	98.706	[85.227, 112.184]	81.246	13.667	322.468
Inflation, π_t (YoY, %)	138	9.513	[7.799, 11.226]	10.179	-9.867	69.329
Real private consumption, C_t	61	1.731×10^6	$[1.636 \times 10^6, 1.825 \times 10^6]$	3.692×10^5	1.059×10^6	2.313×10^6
Consumption growth, ΔC_t (YoY, %)	57	5.068	[2.322, 7.814]	10.350	-19.283	45.009

Real GDP and real consumption are shown in their IMF IFS source units. Full quartiles, skewness, and excess kurtosis are reported in Appendix A.

Unemployment averages 7.189% over the 104 observed quarterly points, but this average combines the high-unemployment early 1990s with the markedly lower post-2010 regime. The unemployment distribution is right-

skewed (1.224). Year-on-year GDP growth averages 5.448% but has a standard deviation of 11.728 and a maximum of 52.060%, consistent with substantial volatility and a break around the mid-2010s. Inflation is even more non-Gaussian: year-on-year inflation has skewness 3.361 and excess kurtosis 15.746, reflecting the extreme crisis episode. These features justify robust rather than purely Gaussian inference.

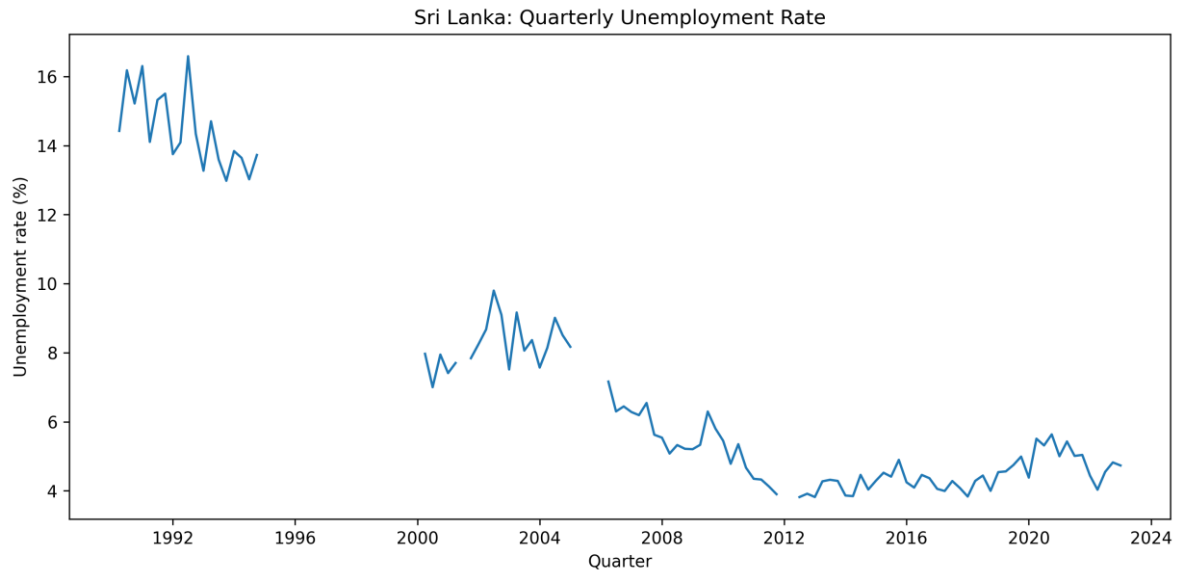


Figure 1 Quarterly unemployment rate in Sri Lanka. Missing quarters correspond to unavailable source observations and are left untreated to preserve the original data structure. Source: IMF International Financial Statistics.

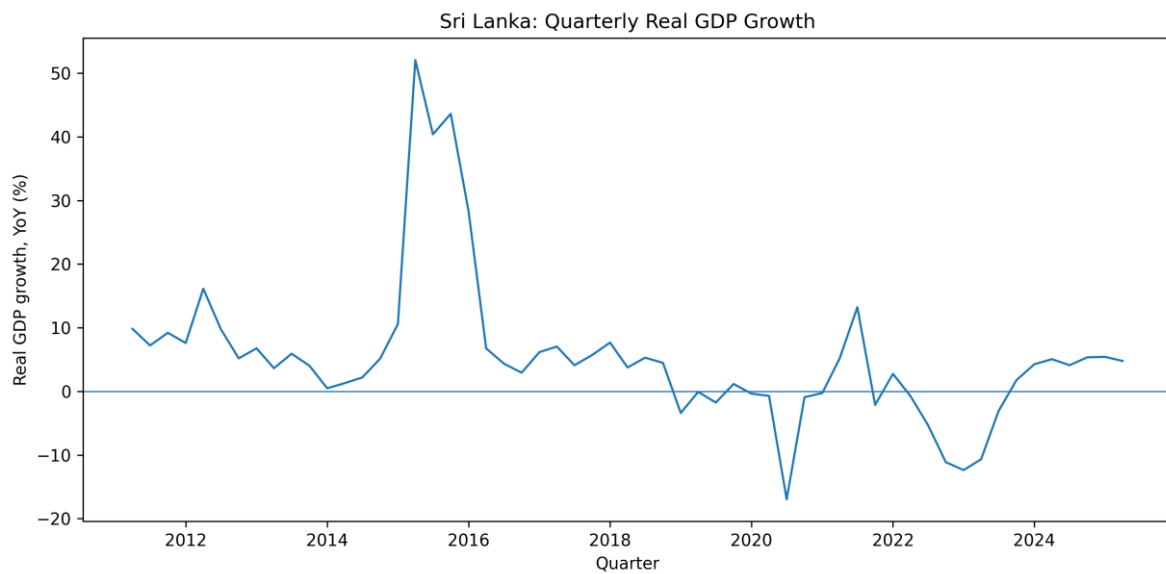


Figure 2 Year-on-year quarterly real GDP growth. The pronounced mid-2010s movement is consistent with the structural-break evidence in the GDP series.

4.2 Stationarity and structural breaks

Table 3 Unit-root, stationarity, and structural-break tests

Series	N	ADF statistic	p-value	KPSS statistic	p-value	Z-A statistic	p-value	Break quarter
Unemployment rate, u_t	104	-1.435	0.850	0.395	0.010	-4.280	0.326	1994Q3
Real GDP, Y_t	61	-1.570	0.804	0.284	0.010	-5.788	0.005	2014Q4
Real GDP growth, ΔY_t	57	-3.725	0.021	0.077	0.100	-4.586	0.176	2015Q4
CPI, P_t	142	0.343	0.996	0.271	0.010	-2.540	0.994	2020Q1
Inflation, π_t	138	-3.589	0.031	0.050	0.100	-4.605	0.170	2013Q1
Real private consumption, C_t	61	-0.850	0.961	0.270	0.010	n.a.	n.a.	n.a.

Series	N	ADF statistic	p-value	KPSS statistic	p-value	Z-A statistic	p-value	Break quarter
Consumption growth, $\Delta \Delta C_t$	57	-3.077	0.112	0.073	0.100	-4.561	0.187	2016Q1

ADF and Zivot-Andrews: null of a unit root. KPSS: null of stationarity. Tests use trend specifications as implemented in the code. The Zivot-Andrews statistic for real consumption was not returned in the run.

The tests strongly reject treating all variables as stationary in levels. For unemployment, the ADF test fails to reject a unit root ($p=0.850$), KPSS rejects stationarity ($p=0.010$), and Zivot-Andrews also fails to reject the unit-root null ($p=0.326$). Real GDP has the same ADF/KPSS pattern, although Zivot-Andrews rejects the unit-root null when allowing a break at 2014Q4 ($p=0.0048$), indicating that structural change materially affects inference. CPI is clearly nonstationary. By contrast, year-on-year GDP growth is more compatible with stationarity (ADF $p=0.0208$; KPSS $p=0.10$), as is year-on-year inflation (ADF $p=0.0308$; KPSS $p=0.10$). Consumption growth is mixed: ADF does not reject at conventional levels ($p=0.112$), whereas KPSS does not reject stationarity ($p=0.10$). This preliminary evidence is the main reason the dynamic specification uses output growth rather than the real-GDP level.

4.3 Pairwise associations

The pairwise results reinforce the distinction between levels and growth rates. In the 50-quarter overlap, unemployment and real GDP have a modest positive Pearson correlation ($r=0.255$, $p=0.074$), whereas unemployment and year-on-year GDP growth have a negative Pearson correlation ($r=-0.151$, $p=0.317$) and a stronger negative Spearman correlation ($\rho = -0.422$, $p=0.0035$). The difference suggests that the rank association between cyclical activity and unemployment is more robust than a simple linear levels relation. The very strong negative level correlation between unemployment and CPI ($r=-0.759$) should not be interpreted structurally because both level series are nonstationary. Real GDP and real consumption are highly correlated in levels ($r=0.944$) and their year-on-year growth rates are also strongly correlated ($r=0.856$), motivating the consumption-channel analysis. All 21 pairwise Pearson and Spearman estimates are reproduced in Appendix A.

4.4 Validation of recursively expected inflation

Table 4 Recursive AR(4) inflation-expectation validation

Observations (N)	MAE (p.p.)	RMSE (p.p.)	Corr.(forecast, realized)	95% mean-CI coverage
114	2.931	4.884	0.899	42.1%

Coverage refers to the confidence interval for the forecast mean generated by the code, not to a 95% predictive interval for future realized inflation.

The recursive expectation tracks realized next-quarter inflation closely in direction: the forecast-realized correlation is 0.899 over 114 observations. The MAE is 2.931 percentage points and RMSE 4.884. Realized inflation falls inside the code's 95% confidence band for the forecast mean only 42.1% of the time. This low figure should not be read as a failed 95% predictive-coverage test because the code stores mean-confidence intervals rather than observation/prediction intervals. It nevertheless confirms that uncertainty around realized inflation is much wider than uncertainty about the estimated conditional mean, especially during the 2022-2024 inflation episode.

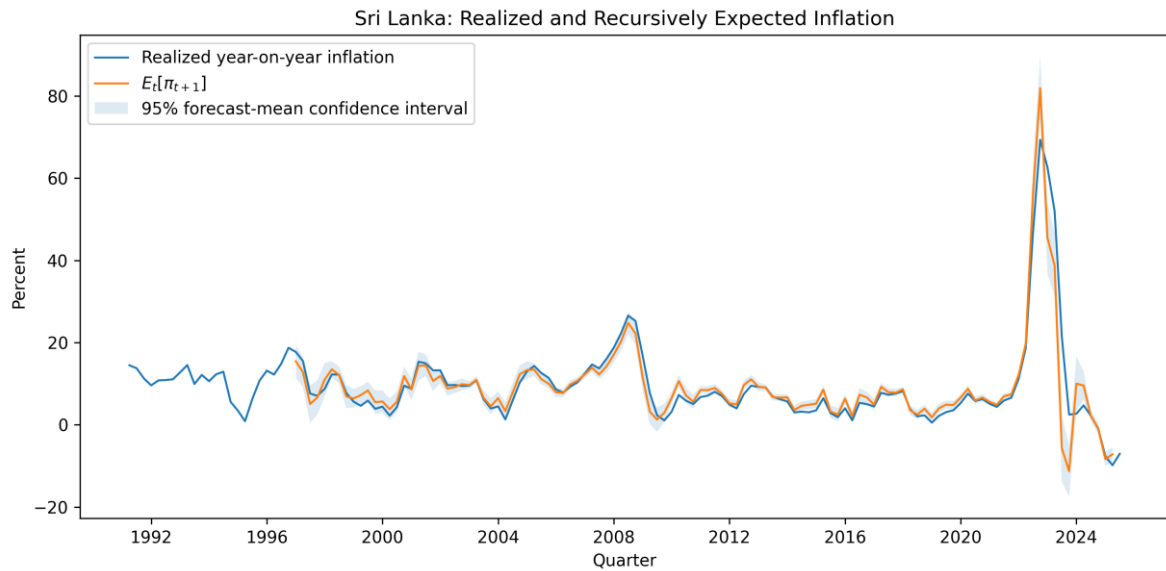


Figure 3 Realized year-on-year inflation and recursively generated one-quarter-ahead expected inflation, with the code-generated 95% confidence interval for the forecast mean.

5 Unemployment regressions

5.1 Model A: direct empirical counterpart of the theoretical levels equation

Table 5 Model A: Newey-West HAC(4) estimates

Regressor	Coefficient	HAC std. error	p-value	95% CI
Intercept (α)	3.993525	0.462602	<0.001	[3.061211, 4.925838]
Expected inflation, $E_t[\pi_{t+1}]$	0.005078	0.001892	0.010	[0.001265, 0.008891]
Real GDP, $Y_{t+1} / 10^{12}$	2.174×10^5	1.803×10^5	0.234	$[-1.460 \times 10^5, 5.807 \times 10^5]$
Quarter 2 indicator	-0.121341	0.095152	0.209	[-0.313108, 0.070426]
Quarter 3 indicator	-0.380034	0.110050	0.001	[-0.601826, -0.158242]
Quarter 4 indicator	-0.160060	0.111240	0.157	[-0.384250, 0.064131]

$N=50$, 2009Q4-2022Q3. The GDP coefficient is for real GDP scaled by 10^{12} and should not be interpreted as a unit-free elasticity. Full OLS, HC3, and HAC output is in Appendix B.

Model A does not provide support for a stable negative income-unemployment relation in levels. The HAC estimate on recursively expected inflation is positive, 0.00508 ($p=0.010$; 95% CI 0.00126 to 0.00889), but the income coefficient is also positive, 217,393 in the scaled units, and statistically insignificant under HAC inference ($p=0.234$; 95% CI -145,951 to 580,737). Classical OLS makes the GDP coefficient appear significant ($p=0.0339$), while HC3 inference moves it to $p=0.0666$ and HAC inference to $p=0.234$. This sensitivity is a warning against treating the positive levels coefficient as a structural effect.

The dependent-data bootstrap is even more cautious. Across 1,000 circular block replications, the expected-inflation coefficient has a bootstrap mean of 0.00224, SE 0.01607, and 95% interval [-0.03522, 0.04148], while the scaled GDP coefficient has a bootstrap mean of 202,518, SE 227,027, and interval [-176,578, 633,583]. Both bootstrap intervals include zero. Consequently, the apparent HAC significance of expected inflation should be described as covariance-estimator dependent rather than as unambiguous evidence.

Model fit and diagnostics are weak. $R^2 = 0.155$ and adjusted $R^2 = 0.059$; AIC=69.664 and BIC=81.136. Durbin-Watson is only 0.605. Breusch-Godfrey strongly rejects no serial correlation ($p=0.000018$) and Ljung-Box gives the same conclusion ($p=1.21e-10$). Breusch-Pagan does not reject homoskedasticity ($p=0.509$), and Jarque-Bera does not reject residual normality ($p=0.166$). CUSUM rejects parameter stability at 5% ($p=0.025$). Thus the levels equation is best viewed as a theoretical benchmark whose empirical residual dynamics are inadequately captured by a static regression.

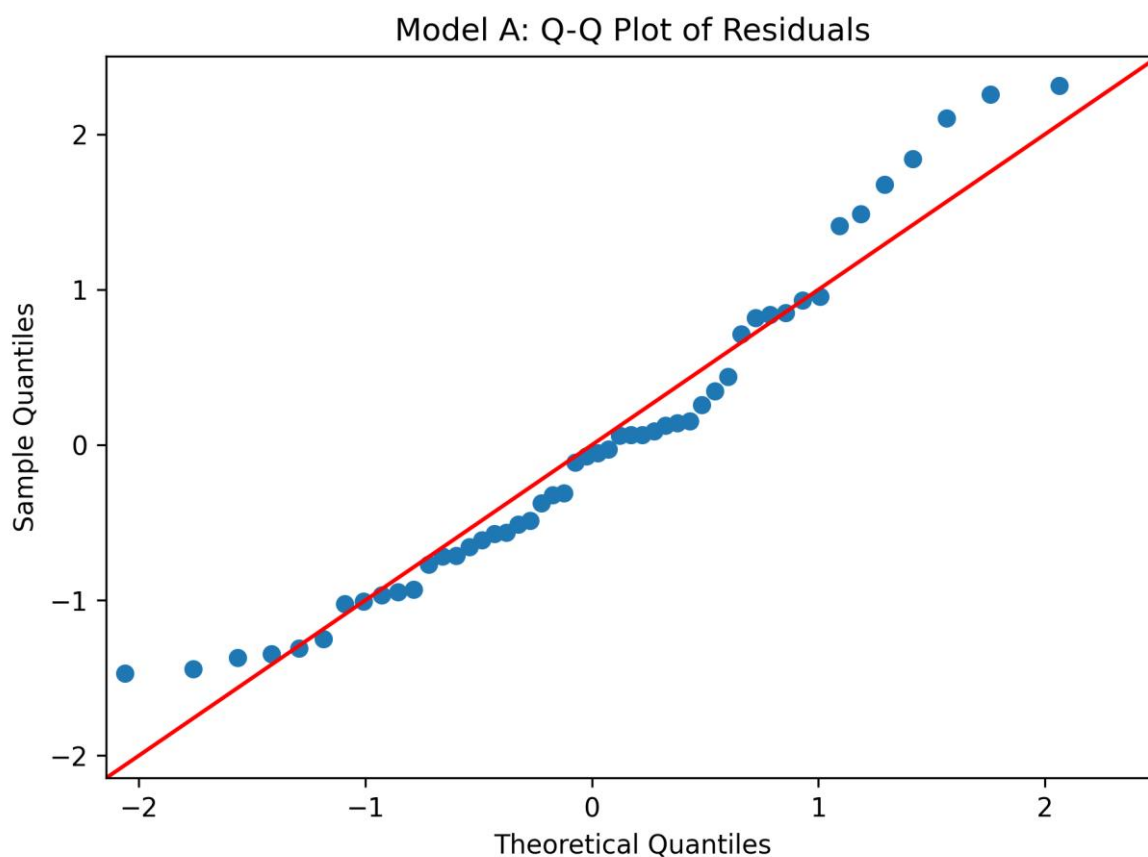


Figure 4 Model A residual Q-Q plot. Visual approximation to normality does not resolve the much stronger serial-correlation and stability failures.

5.2 Model A2: adaptive-expectations robustness

Table 6 Model A2: adaptive-expectations HAC(4) estimates

Regressor	Coefficient	HAC std. error	p-value	95% CI
Intercept (α)	3.999630	0.468468	<0.001	[3.055494, 4.943765]
Adaptive expectation, $E_t[\pi_{t+1}] = \pi_t$	0.005326	0.002193	0.019	[0.000907, 0.009746]
Real GDP, $Y_{t+1} / 10^{12}$	2.166×10^5	1.811×10^5	0.238	$[-1.484 \times 10^5, 5.817 \times 10^5]$
Quarter 2 indicator	-0.119057	0.094879	0.216	[-0.310274, 0.072159]
Quarter 3 indicator	-0.376680	0.110397	0.001	[-0.599171, -0.154190]
Quarter 4 indicator	-0.159116	0.111311	0.160	[-0.383448, 0.065216]

Adaptive expectations set $E_t(\pi_{t+1}) = \pi_t$. Full OLS, HC3, and HAC output is in Appendix C.

Replacing the recursive forecast with adaptive expectations does not change the basic levels-model conclusion. The HAC inflation coefficient is 0.00533 ($p=0.0193$; 95% CI 0.00091 to 0.00975), while the scaled real-GDP coefficient is positive and insignificant (216,634; $p=0.238$; 95% CI -148,406 to 581,674). Fit remains weak ($R^2 = 0.151$; adjusted $R^2 = 0.055$), serial correlation remains severe (Breusch-Godfrey $p=0.000021$; Ljung-Box $p < 0.001$), and CUSUM again rejects stability ($p=0.0229$). The similarity between Models A and A2 suggests that the central problem is not merely the choice of expectations proxy; the static levels specification itself is fragile.

5.3 Model B: dynamic persistence/hysteresis specification

Table 7 Model B: dynamic unemployment specification with HAC(4) inference

Regressor	Coefficient	HAC std. error	p-value	95% CI
Intercept (α)	1.409559	0.453844	0.004	[0.490801, 2.328318]
Lagged unemployment, u_t	0.705276	0.097545	<0.001	[0.507806, 0.902745]

Regressor	Coefficient	HAC std. error	p-value	95% CI
Expected inflation, $E_t[\pi_{t+1}]$	0.003150	0.001725	0.076	[-0.000342, 0.006642]
Real GDP growth, ΔY_{t+1} (YoY, %)	-0.001916	0.003836	0.620	[-0.009681, 0.005850]
Quarter 2 indicator	-0.104914	0.127665	0.416	[-0.363358, 0.153531]
Quarter 3 indicator	-0.360794	0.099982	<0.001	[-0.563196, -0.158391]
Quarter 4 indicator	0.025176	0.178419	0.889	[-0.336014, 0.386367]

N=45, 2010Q4-2022Q3. Full OLS, HC3, HAC, and bootstrap results are in Appendix D.

The dynamic specification changes the empirical picture substantially. Lagged unemployment has coefficient $\rho = 0.7053$ under all covariance estimators; with HAC inference its SE is 0.0975, $p < 0.001$, and 95% CI 0.5078 to 0.9027. The point estimate therefore indicates substantial state dependence: roughly 70% of a conditional unemployment deviation carries into the next quarter, holding the other regressors constant. Expected inflation remains positive but is only marginal under HAC inference (0.00315; $p=0.0757$; 95% CI -0.00034 to 0.00664). Year-on-year real GDP growth has the Keynesian/Okun-consistent negative sign (-0.00192), but its HAC p-value is 0.620 and CI (-0.00968, 0.00585) is wide. The data therefore do not identify a precise short-run output effect in this particular conditional regression.

The direct hysteresis restriction is rejected: $H_0: \rho = 1$ produces a test statistic of -3.021 and $p=0.0045$. In the conditional Model B sample, the evidence is therefore more consistent with persistent but mean-reverting unemployment than with exact unit persistence. This conclusion must be juxtaposed with the univariate full-sample stationarity tests, which fail to reject a unit root in unemployment. The difference reflects changes in sample length, conditioning variables, and test design and argues against categorical claims that Sri Lankan unemployment is either fully stationary or fully hysteretic.

Table 8 Model B circular block-bootstrap inference

Regressor	Replications	Bootstrap mean	Bootstrap std. error	Bootstrap 95% CI
Lagged unemployment, u_t	1000	0.583418	0.221317	[-0.040910, 0.806169]
Expected inflation, $E_t[\pi_{t+1}]$	1000	-0.003432	0.014939	[-0.044692, 0.011280]
Real GDP growth, ΔY_{t+1} (YoY, %)	1000	-0.005933	0.012367	[-0.039119, 0.009329]

Eight-quarter circular blocks; 1,000 replications. The bootstrap interval for ρ is wide and includes zero, highlighting substantial small-sample uncertainty despite the strong HAC t statistic.

Model B explains considerably more variation than Model A: $R^2 = 0.541$, adjusted $R^2 = 0.469$, AIC=37.155, and BIC=49.801. CUSUM no longer rejects stability ($p=0.294$). The residual diagnostics are mixed rather than perfect. Durbin-Watson is 2.497; Breusch-Godfrey rejects no serial correlation at the 5% level ($p=0.040$), while Ljung-Box at lag four does not reject at 5% ($p=0.081$). Breusch-Pagan does not reject homoskedasticity ($p=0.484$), but Jarque-Bera rejects normality at 5% ($p=0.043$). These diagnostics justify retaining HAC and block-bootstrap uncertainty in the substantive interpretation.

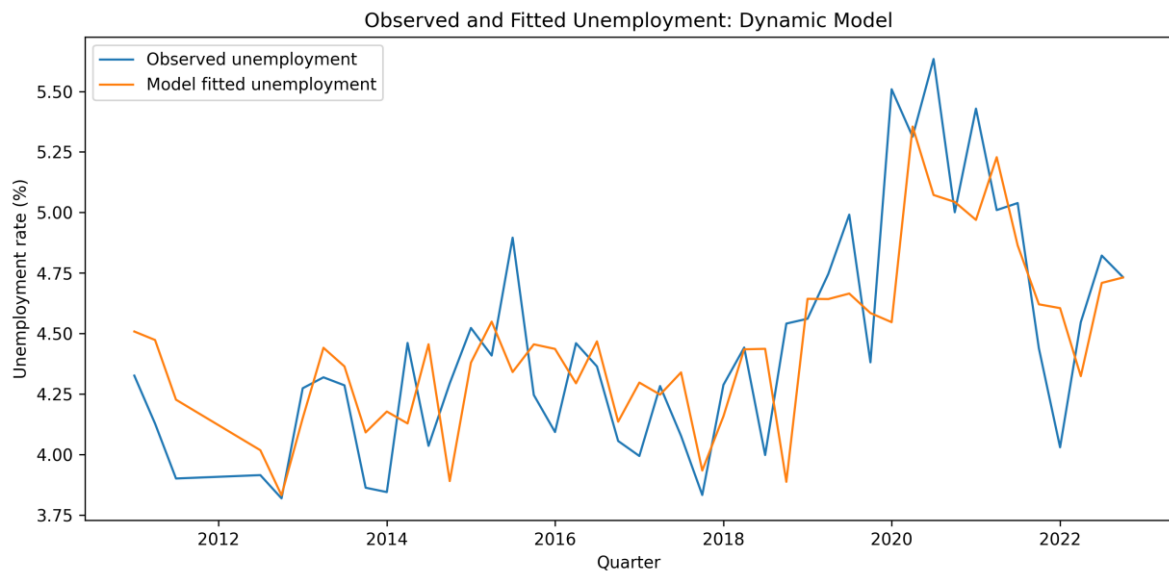


Figure 5 Observed and fitted unemployment in the dynamic Model B sample.

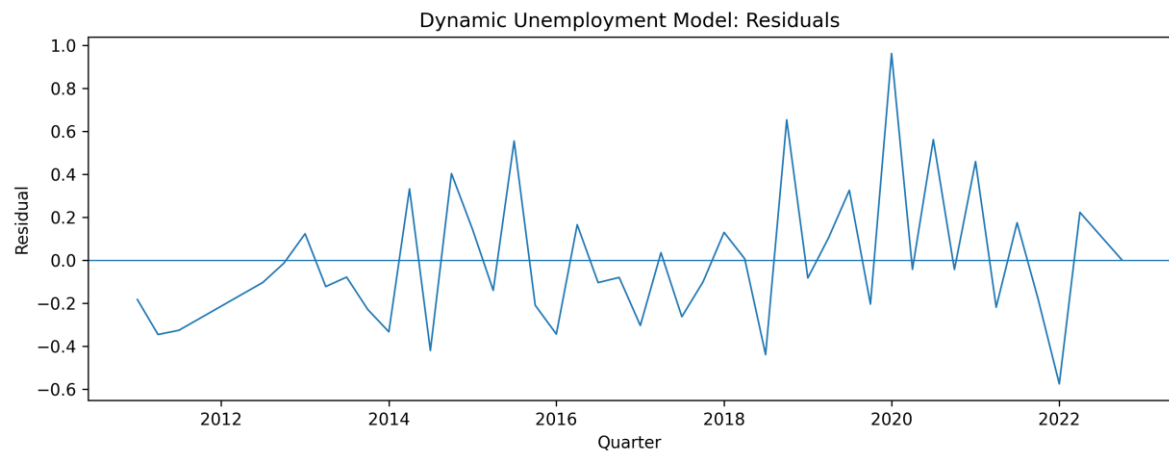


Figure 6 Residuals from the dynamic unemployment model.

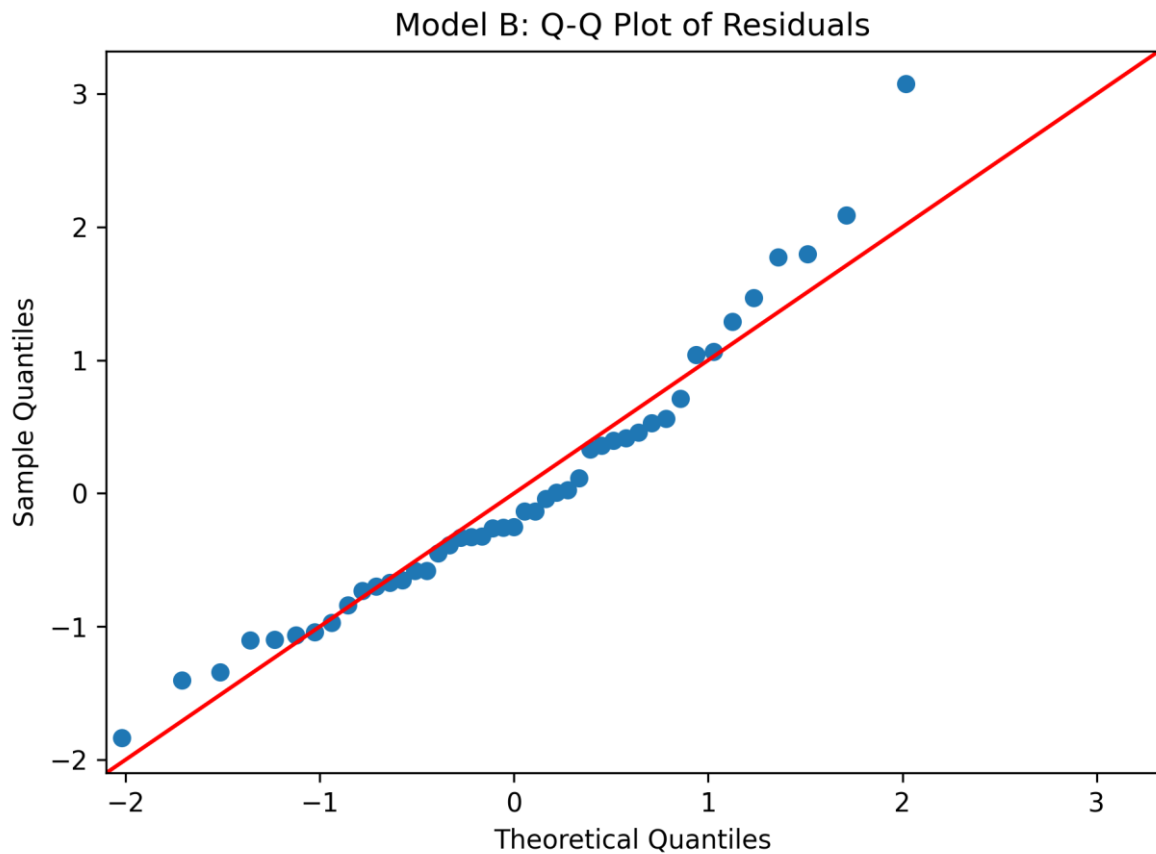


Figure 7 Model B residual Q-Q plot. The upper-tail deviation is consistent with the Jarque-Bera rejection at the 5% level.

6 Consumption channel and long-memory assessment

Table 9 Consumption-income and long-memory results

Component	Parameter / statistic	Estimate	Inference	95% CI / note
Engle-Granger cointegration	Sample size (N)	61		
Engle-Granger cointegration	Engle-Granger statistic	-1.413	$p=0.793$	5% critical value = -3.440
Levels relation: $\ln(C_t)$ on $\ln(Y_t)$	Coefficient on $\ln(Y_t)$	0.817232	$p < 0.001$	95% CI [0.771131, 0.863333]
First-difference relation	$\Delta \ln(Y_t)$	0.727579	$p < 0.001$	95% CI [0.590994, 0.864163]
First-difference relation	Lagged residual, $\hat{e}_{t-1}$	-1.069336	$p < 0.001$	95% CI [-1.270408, -0.868264]
R/S long-memory diagnostic	Hurst exponent, H, on $\Delta \ln(C_t)$	0.395	N=60	Exploratory

Because the Engle-Granger test does not support cointegration, the significant levels coefficient and residual-based adjustment term are not interpreted as a validated long-run error-correction mechanism.

The Engle-Granger statistic is -1.413 with $p=0.793$, far above conventional rejection levels; the 5% critical value is -3.440. Thus the quarterly sample does not support a stable cointegrating relation between log real consumption and log real GDP. The levels regression nevertheless produces a highly precise positive coefficient of 0.8172 (HAC 95% CI 0.7711 to 0.8633). In the absence of cointegration, however, this coefficient cannot be presented as a stable long-run marginal propensity or elasticity; the strong trend correlation is insufficient.

The change specification shows economically meaningful short-run co-movement: $\Delta \log(\text{real GDP})$ has coefficient 0.7276 with HAC 95% CI 0.5910 to 0.8642 and $p < 0.001$. The lagged residual from the levels regression enters at -1.0693 (95% CI -1.2704 to -0.8683, $p < 0.001$). Because cointegration was not established, this term is reported only as a residual-based dynamic statistic rather than as a valid equilibrium error-correction speed. The safest empirical conclusion is that consumption and output growth move strongly together over the available quarterly sample, not that they share a proven stable long-run equilibrium.

The long-memory result does not support persistent fractional dynamics. The R/S estimate on $\Delta \log(\text{real consumption})$ is $H = 0.395$ using 60 observations. Since persistent long memory requires $H > 0.5$, the point

estimate instead lies in the anti-persistent region. Given the small sample and known sensitivity of R/S methods to short-run dynamics, this should not be overinterpreted as proof of anti-persistence; it is sufficient, however, to conclude that the present data do not support calibrating a persistent fractional Brownian consumption process.

7 Discussion

The results call for a conditional interpretation of the levels specification. The direct levels regression is not a convincing empirical confirmation of a Keynesian negative income-unemployment coefficient. Its income coefficient has the wrong sign relative to the theoretical prediction and loses significance under HAC and block-bootstrap inference. More fundamentally, the levels model inherits strong serial correlation and parameter instability, while its key regressors are highly persistent. The correct conclusion is not that higher real GDP raises unemployment, but that this static levels equation is not sufficiently identified to estimate the structural parameter g from the available quarterly overlap.

The dynamic model is more informative. It captures a substantial fraction of the short-run evolution of unemployment and passes the CUSUM stability check. The coefficient on lagged unemployment is large and precisely below one under HAC inference, which supports a concept of state dependence: labor-market conditions carry over from quarter to quarter. Yet the explicit $\rho = 1$ restriction is rejected. Thus the evidence is not consistent with the strongest version of hysteresis in which shocks have permanent one-for-one effects on unemployment. At the same time, the full-sample ADF/KPSS evidence for unemployment is nonstationary, and the block-bootstrap interval for ρ is wide. A balanced interpretation is therefore 'persistent and path dependent, but not robustly unit persistent' rather than either 'fixed natural rate' or 'permanent hysteresis.'

The output results are similarly conditional. The negative sign on GDP growth in Model B is consistent with Keynesian demand transmission and Okun's law, and the rank correlation between unemployment and GDP growth is significantly negative. But the regression coefficient is statistically imprecise. This may reflect the small joint sample, the extraordinary 2015 break in the real-GDP series, the 2019 Easter attacks, the COVID-19 shock, and the 2022 balance-of-payments crisis. A richer open-economy model that incorporates exchange rates, real wages, tourism, external demand, and policy-regime changes is therefore likely to be more informative than an attempt to force a single stable output coefficient through the full crisis sample.

Expected inflation is the one variable that repeatedly enters the levels regressions with a positive HAC coefficient, and the recursive AR(4) forecast tracks subsequent realized inflation closely. Nevertheless, the block-bootstrap interval in Model A includes zero, and the expected-inflation term is only marginally significant in the dynamic model. The evidence therefore supports retaining expectations in the theoretical framework, but not assigning them a quantitatively stable structural coefficient on the basis of this sample.

The consumption results point to short-run rather than long-run transmission. The very high contemporaneous correlation between real GDP and real consumption does not translate into cointegration, and the Hurst statistic does not indicate persistent long memory. The robust result is short-run co-movement in growth rates. From a Keynesian perspective, this still matters: changes in activity and household expenditure move together strongly. What the data do not justify is the stronger claim that consumption follows a persistent fractional process through which policy shocks necessarily propagate indefinitely.

Taken together, these findings sharpen the empirical distinction among the natural-rate hypothesis, hysteresis, and Keynesian demand persistence. They are not binary alternatives that can be settled by one coefficient. Sri Lanka's quarterly data point toward an intermediate case in which unemployment is highly persistent, dynamic specifications outperform static levels equations, and macroeconomic shocks matter, but the exact long-run policy multiplier is not identified. That conclusion is consistent with the broader literature emphasizing time-varying unemployment benchmarks, institutional interactions, and uncertainty around NAIRU estimation [5-13].

8 Conclusion

This paper analyzes Sri Lankan unemployment using observed quarterly data and robust time-series inference. All series enter at their observed quarterly frequency; no interpolation, forward fill, backward fill, or synthetic frequency conversion is used. The expectation term is generated recursively without using future information, and the empirical design separates the theoretical levels equation from a dynamic persistence model.

Three substantive conclusions follow. First, the direct levels specification does not provide stable evidence for a negative real-income coefficient. The model has low explanatory power, strong serial correlation, parameter instability, and coefficient sensitivity across covariance estimators and the block bootstrap. Second, the dynamic model provides clear evidence that unemployment is state dependent, with ρ approximately 0.705 under HAC inference, but the restriction $\rho = 1$ is rejected. The most defensible characterization is therefore substantial persistence with mean reversion rather than full hysteresis. Third, the consumption channel is supported in short-run growth rates but not as a proven long-run cointegrating relation, and the exploratory Hurst estimate of 0.395 does not support persistent fractional long memory.

The policy implication is correspondingly modest. The results do not show that demand management is powerless, but neither do they identify a causal policy rule capable of permanently eliminating a natural rate. In a small open economy exposed to external financing, supply, exchange-rate, and institutional shocks, the unemployment response to expansion is likely to be regime dependent. Future work should extend the joint quarterly sample using official historical national-accounts data, incorporate exchange rates, real wages, fiscal and monetary instruments, tourism receipts, and structural-break or regime-switching methods, and use identification strategies that separate policy shocks from endogenous responses to unemployment. Such extensions would provide a stronger test of whether policy can shift the medium-run unemployment benchmark rather than merely accelerate cyclical recovery.

Declarations

Conflict of interest: The author declares no conflict of interest.

Funding: No specific funding was received for this research.

Author contributions: Houssam Boughabi is the sole author and was responsible for conceptualization, methodology, software implementation, formal analysis, interpretation, and manuscript preparation.

Ethics approval: Not applicable. The study uses aggregate publicly available macroeconomic data and does not involve human participants or animals.

Data availability: The empirical series are publicly available from the IMF International Financial Statistics database and were retrieved through DBnomics. The executed run uses Q.LK.LUR_PT (unemployment), Q.LK.NGDP_R_NSA_XDC (real GDP), Q.LK.NCP_R_NSA_XDC (real private consumption), and Q.LK.PCPI_IX (CPI).

Code availability: The Python code and generated replication outputs underlying the reported tables and figures are retained in the accompanying empirical replication package.

Appendix A. Full preliminary statistics

A.1 Full descriptive statistics

Table A1 Full descriptive statistics from the compiled output

Series / notation	N	Mean	Std. dev.	Min	25th pct.	Median	75th pct.	Max	Skewness	Excess kurtosis
u_t (%)	104	7.189	3.777	3.818	4.402	5.389	8.279	16.586	1.224	0.132
Y_t	61	2.717×10^6	6.572×10^5	1.415×10^6	1.967×10^6	2.959×10^6	3.290×10^6	3.563×10^6	-0.558	-1.215
ΔY_t (%)	57	5.448	11.728	-16.975	-0.087	4.367	6.760	52.060	2.097	6.455
P_t	142	98.706	81.246	13.667	33.298	77.469	136.757	322.468	1.255	1.022
π_t (%)	138	9.513	10.179	-9.867	4.415	7.593	12.198	69.329	3.361	15.746
C_t	61	1.731×10^6	3.692×10^5	1.059×10^6	1.352×10^6	1.846×10^6	1.999×10^6	2.313×10^6	-0.440	-1.183
ΔC_t (%)	57	5.068	10.350	-19.283	-0.214	3.679	7.610	45.009	1.642	5.203

Mean 95% confidence intervals are reported in Table 2 of the main text.

A.2 Full pairwise Pearson and Spearman correlations

Table A2 Full pairwise correlation matrix in long form

Variable X	Variable Y	N	Pearson r	p-value	Spearman ρ	p-value
u_t	Y_t	50	0.255	0.074	0.202	0.159
u_t	ΔY_t	46	-0.151	0.317	-0.422	0.003
u_t	P_t	104	-0.759	<0.001	-0.818	<0.001

Variable X	Variable Y	N	Pearson r	p-value	Spearman ρ	p-value
u_t	π_t	100	0.059	0.558	0.384	<0.001
u_t	C_t	50	0.312	0.027	0.311	0.028
u_t	ΔC_t	46	-0.127	0.400	-0.145	0.335
Y_t	ΔY_t	57	-0.115	0.395	-0.296	0.025
Y_t	P_t	61	0.495	<0.001	0.710	<0.001
Y_t	π_t	61	0.028	0.833	-0.217	0.093
Y_t	C_t	61	0.944	<0.001	0.853	<0.001
Y_t	ΔC_t	57	-0.120	0.373	-0.266	0.045
ΔY_t	P_t	57	-0.353	0.007	-0.548	<0.001
ΔY_t	π_t	57	-0.379	0.004	-0.208	0.121
ΔY_t	C_t	57	-0.113	0.401	-0.337	0.010
ΔY_t	ΔC_t	57	0.856	<0.001	0.538	<0.001
P_t	π_t	138	0.114	0.182	-0.372	<0.001
P_t	C_t	61	0.578	<0.001	0.752	<0.001
P_t	ΔC_t	57	-0.256	0.055	-0.337	0.010
π_t	C_t	61	0.087	0.504	-0.265	0.039
π_t	ΔC_t	57	-0.246	0.066	-0.351	0.007
C_t	ΔC_t	57	-0.037	0.786	-0.090	0.504

A.3 Recursive inflation-expectation validation summary

The full forecast path contains 114 validation observations and is visualized in Figure 3. The code-generated summary is reproduced exactly in Table 4. The observation-level file 05a_expected_inflation_validation.csv is part of the replication package; the manuscript does not print all 114 rows because they are already represented graphically and are not independent test statistics.

Appendix B. Full Model A output

Table B1 Model A: OLS, HC3, and HAC(4) coefficient output

Inference method	Regressor	Coefficient	Std. error	t / z	p-value	95% CI	N
OLS (classical)	Intercept (α)	3.993525	0.275694	14.485	<0.001	[3.437900, 4.549149]	50
OLS (classical)	$E[\pi_{t+1}]$	0.005078	0.005146	0.987	0.329	[-0.005293, 0.015448]	50
OLS (classical)	$Y_{t+1} / 10^{12}$	2.174×10^5	99266.980767	2.190	0.034	[17333.822020, 4.175×10^5]	50
OLS (classical)	Quarter 2 indicator	-0.121341	0.182483	-0.665	0.510	[-0.489112, 0.246430]	50
OLS (classical)	Quarter 3 indicator	-0.380034	0.190108	-1.999	0.052	[-0.763172, 0.003104]	50
OLS (classical)	Quarter 4 indicator	-0.160060	0.186921	-0.856	0.396	[-0.536775, 0.216655]	50
OLS (HC3 robust)	Intercept (α)	3.993525	0.337319	11.839	<0.001	[3.313702, 4.673347]	50
OLS (HC3 robust)	$E[\pi_{t+1}]$	0.005078	0.002330	2.179	0.035	[0.000382, 0.009774]	50
OLS (HC3 robust)	$Y_{t+1} / 10^{12}$	2.174×10^5	1.156×10^5	1.881	0.067	[-15557.514852, 4.503×10^5]	50
OLS (HC3 robust)	Quarter 2 indicator	-0.121341	0.187300	-0.648	0.520	[-0.498818, 0.256137]	50
OLS (HC3 robust)	Quarter 3 indicator	-0.380034	0.161624	-2.351	0.023	[-0.705767, -0.054301]	50
OLS (HC3 robust)	Quarter 4 indicator	-0.160060	0.200831	-0.797	0.430	[-0.564809, 0.244689]	50
OLS (Newey-West HAC(4))	Intercept (α)	3.993525	0.462602	8.633	<0.001	[3.061211, 4.925838]	50
OLS (Newey-West HAC(4))	$E[\pi_{t+1}]$	0.005078	0.001892	2.684	0.010	[0.001265, 0.008891]	50
OLS (Newey-West HAC(4))	$Y_{t+1} / 10^{12}$	2.174×10^5	1.803×10^5	1.206	0.234	[-1.460 $\times 10^5$, 5.807×10^5]	50
OLS (Newey-West HAC(4))	Quarter 2 indicator	-0.121341	0.095152	-1.275	0.209	[-0.313108, 0.070426]	50
OLS (Newey-West HAC(4))	Quarter 3 indicator	-0.380034	0.110050	-3.453	0.001	[-0.601826, -0.158242]	50
OLS (Newey-West HAC(4))	Quarter 4 indicator	-0.160060	0.111240	-1.439	0.157	[-0.384250, 0.064131]	50

Table B2 Model A circular block-bootstrap output

Regressor	Successful replications	Bootstrap mean	Bootstrap std. error	95% CI lower	95% CI upper
$E_t[\pi_{t+1}]$	1000	0.002242	0.016069	-0.035225	0.041483
$Y_{t+1} / 10^{12}$	1000	2.025×10^5	2.270×10^5	-1.766×10^5	6.336×10^5

Table B3 Model A diagnostics from the compiled output

N	R ²	Adj. R ²	AIC	BIC	DW	BG LM(4)	BG p-value	LB Q(4)	LB p-value	BP p-value	JB p-value	CUSUM p-value	RESET p-value
50	0.155	0.059	69.664	81.136	0.605	27.195	<0.001	52.265	<0.001	0.509	0.166	0.025	n.a.

Appendix C. Adaptive-expectations robustness output

Table C1 Model A2: OLS, HC3, and HAC(4) coefficient output

Inference method	Regressor	Coefficient	Std. error	t / z	p-value	95% CI	N
OLS (classical)	Intercept (α)	3.999630	0.275994	14.492	<0.001	[3.443401, 4.555859]	50
OLS (classical)	$E_t[\pi_{t+1}] = \pi_t$	0.005326	0.006059	0.879	0.384	[-0.006884, 0.017537]	50
OLS (classical)	$Y_{t+1} / 10^{12}$	2.166×10^5	99495.052427	2.177	0.035	[16114.737170, 4.172×10^5]	50
OLS (classical)	Quarter 2 indicator	-0.119057	0.182823	-0.651	0.518	[-0.487513, 0.249398]	50
OLS (classical)	Quarter 3 indicator	-0.376680	0.190437	-1.978	0.054	[-0.760480, 0.007119]	50
OLS (classical)	Quarter 4 indicator	-0.159116	0.187373	-0.849	0.400	[-0.536741, 0.218509]	50
OLS (HC3 robust)	Intercept (α)	3.999630	0.341858	11.700	<0.001	[3.310660, 4.688599]	50
OLS (HC3 robust)	$E_t[\pi_{t+1}] = \pi_t$	0.005326	0.003718	1.433	0.159	[-0.002166, 0.012819]	50
OLS (HC3 robust)	$Y_{t+1} / 10^{12}$	2.166×10^5	1.161×10^5	1.865	0.069	[-17443.672900, 4.507×10^5]	50
OLS (HC3 robust)	Quarter 2 indicator	-0.119057	0.187154	-0.636	0.528	[-0.496242, 0.258128]	50
OLS (HC3 robust)	Quarter 3 indicator	-0.376680	0.161671	-2.330	0.024	[-0.702507, -0.050854]	50
OLS (HC3 robust)	Quarter 4 indicator	-0.159116	0.201747	-0.789	0.435	[-0.565711, 0.247479]	50
OLS (Newey-West HAC(4))	Intercept (α)	3.999630	0.468468	8.538	<0.001	[3.055494, 4.943765]	50
OLS (Newey-West HAC(4))	$E_t[\pi_{t+1}] = \pi_t$	0.005326	0.002193	2.429	0.019	[0.000907, 0.009746]	50
OLS (Newey-West HAC(4))	$Y_{t+1} / 10^{12}$	2.166×10^5	1.811×10^5	1.196	0.238	[-1.484 $\times 10^5$, 5.817×10^5]	50
OLS (Newey-West HAC(4))	Quarter 2 indicator	-0.119057	0.094879	-1.255	0.216	[-0.310274, 0.072159]	50
OLS (Newey-West HAC(4))	Quarter 3 indicator	-0.376680	0.110397	-3.412	0.001	[-0.599171, -0.154190]	50
OLS (Newey-West HAC(4))	Quarter 4 indicator	-0.159116	0.111311	-1.429	0.160	[-0.383448, 0.065216]	50

Table C2 Model A2 diagnostics

N	R ²	Adj. R ²	AIC	BIC	DW	BG LM(4)	BG p-value	LB Q(4)	LB p-value	BP p-value	JB p-value	CUSUM p-value	RESET p-value
50	0.151	0.055	69.888	81.360	0.603	26.930	<0.001	51.965	<0.001	0.523	0.174	0.023	n.a.

Appendix D. Full dynamic-model output

Table D1 Model B: OLS, HC3, and HAC(4) coefficient output

Inference method	Regressor	Coefficient	Std. error	t / z	p-value	95% CI	N
OLS (classical)	Intercept (α)	1.409559	0.518677	2.718	0.010	[0.359554, 2.459565]	45
OLS (classical)	u_t	0.705276	0.113576	6.210	<0.001	[0.475353, 0.935198]	45
OLS (classical)	$E_t[\pi_{t+1}]$	0.003150	0.004056	0.777	0.442	[-0.005060, 0.011360]	45
OLS (classical)	$\Delta \cdot Y_{t+1} (Y_0 Y_t \%)$	-0.001916	0.004285	-0.447	0.657	[-0.010591, 0.006760]	45
OLS (classical)	Quarter 2 indicator	-0.104914	0.142829	-0.735	0.467	[-0.394055, 0.184228]	45
OLS (classical)	Quarter 3 indicator	-0.360794	0.147326	-2.449	0.019	[-0.659039, -0.062549]	45
OLS (classical)	Quarter 4 indicator	0.025176	0.147004	0.171	0.865	[-0.272418, 0.322770]	45
OLS (HC3 robust)	Intercept (α)	1.409559	0.513372	2.746	0.009	[0.370291, 2.448828]	45

Inference method	Regressor	Coefficient	Std. error	t / z	p-value	95% CI	N
OLS (HC3 robust)	u_t	0.705276	0.111722	6.313	<0.001	[0.479106, 0.931445]	45
OLS (HC3 robust)	$E_t[\pi_{t+1}]$	0.003150	0.002324	1.355	0.183	[-0.001555, 0.007856]	45
OLS (HC3 robust)	ΔY_{t+1} (YoY, %)	-0.001916	0.005519	-0.347	0.730	[-0.013088, 0.009256]	45
OLS (HC3 robust)	Quarter 2 indicator	-0.104914	0.134117	-0.782	0.439	[-0.376418, 0.166591]	45
OLS (HC3 robust)	Quarter 3 indicator	-0.360794	0.125182	-2.882	0.006	[-0.614212, -0.107375]	45
OLS (HC3 robust)	Quarter 4 indicator	0.025176	0.165101	0.152	0.880	[-0.309054, 0.359406]	45
OLS (Newey-West HAC(4))	Intercept (α)	1.409559	0.453844	3.106	0.004	[0.490801, 2.328318]	45
OLS (Newey-West HAC(4))	u_t	0.705276	0.097545	7.230	<0.001	[0.507806, 0.902745]	45
OLS (Newey-West HAC(4))	$E_t[\pi_{t+1}]$	0.003150	0.001725	1.826	0.076	[-0.000342, 0.006642]	45
OLS (Newey-West HAC(4))	ΔY_{t+1} (YoY, %)	-0.001916	0.003836	-0.499	0.620	[-0.009681, 0.005850]	45
OLS (Newey-West HAC(4))	Quarter 2 indicator	-0.104914	0.127665	-0.822	0.416	[-0.363358, 0.153531]	45
OLS (Newey-West HAC(4))	Quarter 3 indicator	-0.360794	0.099982	-3.609	<0.001	[-0.563196, -0.158391]	45
OLS (Newey-West HAC(4))	Quarter 4 indicator	0.025176	0.178419	0.141	0.889	[-0.336014, 0.386367]	45

Table D2 Model B circular block-bootstrap output

Regressor	Successful replications	Bootstrap mean	Bootstrap std. error	95% CI lower	95% CI upper
u_t	1000	0.583418	0.221317	-0.040910	0.806169
$E_t[\pi_{t+1}]$	1000	-0.003432	0.014939	-0.044692	0.011280
ΔY_{t+1} (YoY, %)	1000	-0.005933	0.012367	-0.039119	0.009329

Table D3 Model B diagnostics

N	R ²	Adj. R ²	AIC	BIC	DW	BG LM(4)	BG p-value	LB Q(4)	LB p-value	BP p-value	JB p-value	CUSUM p-value	RESET p-value
45	0.541	0.469	37.155	49.801	2.497	10.029	0.040	8.297	0.081	0.484	0.043	0.294	n.a.

Table D4 Persistence restriction test

Restriction	Estimate of ρ	Test statistic	p-value
$H_0: \rho = 1$	0.705276	-3.021	0.004

Appendix E. Full consumption-channel output

Table E1 Consumption-income cointegration test

N	Engle-Granger statistic	p-value	1% critical value	5% critical value	10% critical value
61	-1.413425	0.793	-4.088285	-3.439860	-3.115892

Table E2 Log consumption-log GDP levels regression with HAC inference

Regressor	Coefficient	HAC std. error	p-value	95% CI lower	95% CI upper
Intercept (α)	2.258991	0.337323	<0.001	1.584009	2.933974
$\ln(Y_t)$	0.817232	0.023039	<0.001	0.771131	0.863333

Table E3 Residual-based dynamic consumption regression with HAC inference

Regressor	Coefficient	HAC std. error	p-value	95% CI lower	95% CI upper
Intercept (α)	0.000343	0.008627	0.968	-0.016932	0.017619
$\Delta \ln(Y_t)$	0.727579	0.068208	<0.001	0.590994	0.864163
Lagged residual, $\hat{e}_{t-1}$	-1.069336	0.100412	<0.001	-1.270408	-0.868264

Table E4 Exploratory consumption long-memory diagnostic

Series / transformation	N	Hurst exponent (R/S), H	Interpretation
$\Delta \ln(C_t)$	60	0.394531	Exploratory only. $H > 0.5$ suggests persistence, but Hurst estimation alone does not establish fractional Brownian motion.

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