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Persistence in China's household consumption level: implications for the new growth model

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Abstract

In this paper, we investigate the statistical features of China's household consumption level by calculating its degree of persistence and testing for the existence of trends in the data. Further, we contrast the results obtained in China with those corresponding to the G7 economies and in the other three countries that make up the fastest developing BRICs group, that is, Brazil, Russia and India. Our results show large degrees of heterogeneity depending on the country and the model used for the analysis though, in general, lower degrees of dependence and thus, more evidence of transitory shocks is observed in China and the rest of BRIC economies than in the G7 countries.

Keywords China · Consumption · Household · Persistence · Time trends

JEL Classification C25 · F64 · F21

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1 Introduction

China's government is trying to stimulate household consumption and foster a more balanced growth model which is less dependent on exports. In 2006, in the 11th Five-Year Plan, domestic market potential was qualified as "*enormous*" considering the "*accelerated upgrading of consumption structure of urban residents*", but "*uncoordinated investment and consumption relationship*" had already been identified as a new major problem (China Government 2006). In 2011, Premier Wen Jiabao delivered a report admitting "*a serious problem in that our development is not yet well balanced*, *coordinated or sustainable*", because of "*... imbalance between investment and consumption...*" among other reasons (Wen 2011). One decade later, in 2022, in a domestic context still conditioned by the Covid19 epidemic and the Covid-Zero policy, China's State Council issued a document with recommendations "*further unleashing consumption potential* ", and " *promoting continued recovery of consumption* " (State Council 2022).

China's economy reached extraordinary GDP growth rates over a remarkably long period of time —1978–2010—, but in the 2010s the reduced push of exports resulted in more moderate growth. The proposal to reduce dependence on exports and make domestic consumption the new engine growth is part of China's "*new normal*" approach to economic expansion, which also includes the objectives of better quality growth and more innovation to focus on higher value added activities. The challenge is huge and the People's Republic of China (PRC)'s efforts to rebalance its development model has attained limited success until now, with household consumption representing slightly less than 40% of GDP, which compares poorly with shares between 50% and 65% of GDP in the G7 and the rest of BRIC economies (see Fig. 1). Furthermore, the pandemic has led to a weakening of consumer confidence and household income expectations. According to a survey conducted by the People's Bank of China (PBOC) in the fourth quarter of 2022, only 22.8 per cent expected to increase consumption, whereas more than 60% preferred more savings deposits (People's Bank of China 2022).

The main objective of this paper is to analyse the statistical features of China's household consumption level in the long term. To this purpose, we calculate the degree of persistence of household's consumption in China, and we compare it with the results obtained in the G7 and in three BRIC economies (hereinafter referred to as BRI economies). The rest of the paper is structured as follows: Sect. 2 briefly analyses the previous literature on China's economic model challenges and solutions. Section 3 focuses on the stationarity of the consumption/income ratio. Section 4 describes the methodology and data used in the paper. In Sect. 5 we present the main empirical finding. Section 6 deals with the discussion of the results and, finally, in Sect. 7 we summarize the main conclusions and discuss policy implications.

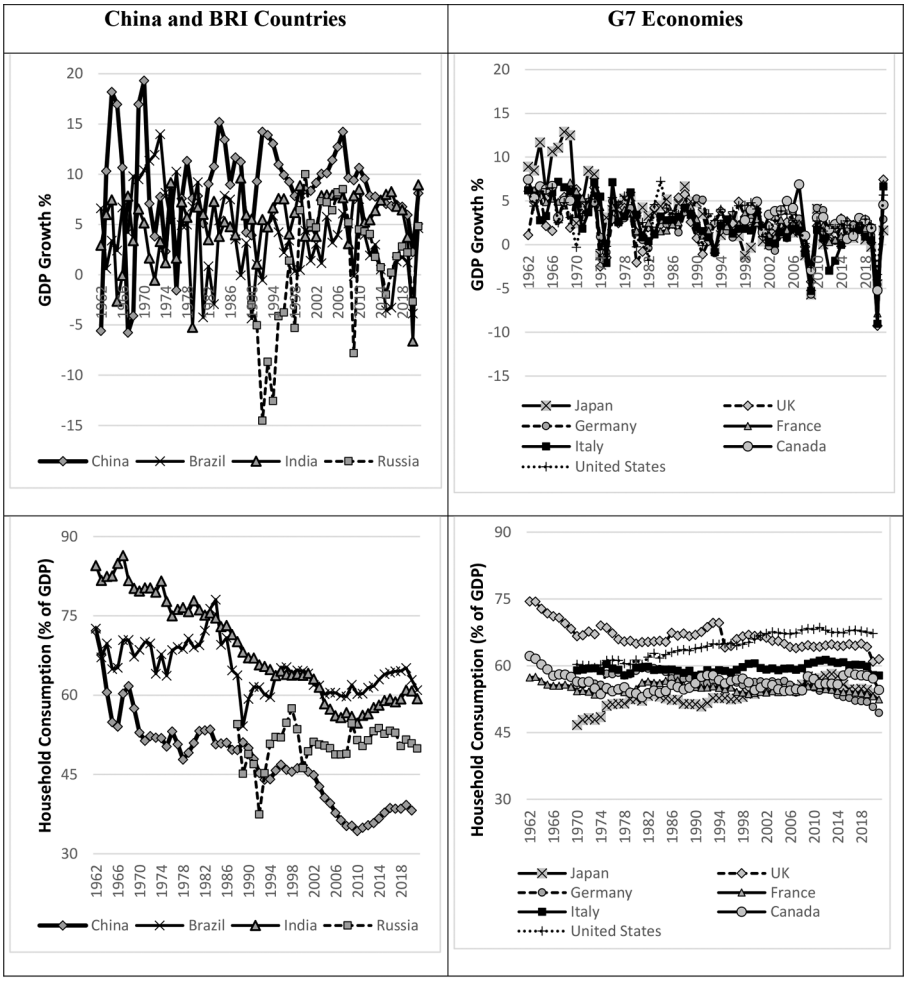

Fig. 1 GDP growth and share of household consumption. Source: World Bank. World Development Indicators

2 China economic model: the role of households' consumption in growth

There is a wide body of the literature studying the characteristics of Chinese economic growth and structural transformation. Several studies gravitate around the contribution to GDP of capital accumulation and residually estimated total factor productivity (TFP), most of them concluding that investment was a main contributor to growth in China (Borensztein and Ostry 1996; Bosworth and Collins 2008; Lin and Zhang 2015; etc.). Having become the '*workshop of the world*', the PRC continues to have a strong industrial exporting sector, compared to other more service-based and domestic dependent economies (Chen 2015), which goes some way to explaining China's more capital-intensive nature and its much higher share of savings and investment in

GDP (investment share is permanently above 40% in the PRC compared to slightly more than 20% in G7 and BRI economies).

There is also a vast literature documenting China's household consumption. A large share of it focuses on the analysis of the reasons that cause weak consumption patterns. Some researchers point to the rising of precautionary saving in parallel to market-oriented economy reforms and the disappearance of some public services — healthcare, housing, education, etc.—, earlier provided by State Owned Enterprises (Kuijs 2005; Blanchard and Giavazzi 2005; Chamon and Prasad 2010; etc.), while others focus on the role of weak financial markets and the share of income and wages in GDP (Aziz and Cui 2007; Dautović et al. 2019; etc.). Some authors hint that, due to lack of accuracy of consumption data, underestimation of household income and consumption could explain, at least to some point, the lower consumption levels in comparison with other economies (Wang and Woo 2011; Zhang and Zhu 2013; Huang et al. 2013; Guo 2015; Holz 2020; etc.).

Regarding the increase in China's domestic savings, a first set of scholars affirm that the underlying motives for saving are the same as elsewhere, such as income uncertainty (Chamon et al. 2013; Choi et al. 2017) or the increasing burden of housing, education and healthcare costs (Chamon and Prasad 2010). Instead, a second set of scholars consider that there are varying degrees of weight to variables of "*Chineseness*", such as habits, culture and even language (Wei and Zhang 2011; İmrohoroğlu and Zhao 2018; Lugauer et al. 2019; Yang and Wang 2021).

Several authors have also analyzed the relationship between household consumption and other aspects such as economic policy, urbanization, size of cities or housing prices (Bielskis 2024; Broadway and Haisken-DeNew 2019; etc.). For example, Keely and Anderson (2015) analyze the influence of fiscal and monetary instruments on private household expenditures in China, concluding that the impact is smaller in the PRC than in other Asian countries. Yang et al. (2018) study the correlation between housing price and consumption, finding that there is a negative correlation between them. Ding and Li (2022) show that the size of the cities and the level of wages have an import influence on household consumption in China. Other papers focus on the relationship between urbanization and household consumption in the PRC (Molnar et al. 2017; etc.).

The time series properties of the consumption-income ratio have been analysed extensively since the second half of the 20th century, but it remains a controversial issue in macroeconomics. From a theoretical approach, the relative income hypothesis (Duesenberry 1952; Modigliani 1986), the permanent income hypothesis (Friedman 1957), and the life cycle hypothesis (Ando and Modigliani 1963) predict a stationary average propensity to consume (APC), so that the series should converge towards a constant in the long run and deviations should be temporary. On the contrary, the Keynesian absolute income hypothesis (Keynes 1936), and the involuntary savings theory (Deaton 1977; Campbell and Deaton 1989) suggest that the APC does not converge towards a constant in the long run.

The discussion about the statistical behaviour of consumption-income ratio has profound implications for policy makers: if APC is stationary, the implementation of fiscal and monetary policy to stimulate consumption in the short term would have no impact in the long term; on the contrary, if APC is non-stationary, policy shocks

could have a permanent effect on economy. The issue is that the results of empirical investigations are mixed, so that up to present both approaches have been validated in specific contexts.

Starting with some recent investigations that indicate that the consumption–income ratio is non-stationary, Tsionas and Christopoulos (2002) tested the stationarity property of this ratio applying three panel unit root procedures and an asymmetric unit root statistic to a sample of 14 European Union countries over the period 1960 − 1999. The authors concluded that, although the unit root procedures suggested stationarity, the results of the asymmetric unit root tests finding a stationary stochastic process was more plausible. Cook (2005), reconsidered recent research using the minimum LM unit root tests of Lee and Strazicich (2003), which allows for two breaks in both level and trend, obtaining that the APC is nonstationary for all 20 OECD economies examined over the period 1955 to 1994. Fallahi (2012) studied the stationarity of APC for 23 OECD countries over 1950–2007 period using three different bootstrapping techniques to construct the 90% confidence intervals, obtaining that the APC was nonstationary in most of the countries.

On the other hand, there is also a numerous bunch of papers supporting stationarity applying recent advances in panel and asymmetric unit root tests. Romero-Ávila (2009) examined the stochastic properties of the consumption–income ratio for a sample of 23 OECD countries for the period, 1960–2005. The results obtained from the application of the panel stationarity test, after allowing for multiple structural breaks as well as for general forms of cross-sectional dependence through bootstrap methods, supported the existence of regime-wise stationarity. However, the large width of the 90% confidence intervals associated with the half-life point estimates for most countries indicated a high degree of uncertainty and low precision in half-life estimation, as they were consistent with a wide range of degrees of persistence. Liao et al. (2011) examined the existence of the mean-reverting behaviour of the consumption-income ratio from a panel of 24 OECD countries through the application of the series-specific SURADF panel unit root test. The results showed mean-reverting behaviour in 22 OECD countries and the half-life of the estimation of consumption-income ratio for these 22 countries was between 0.28 and 3.48 years. Elmi and Ranjbar (2013) examined the integrated nature of the APC for 16 OECD countries from 1960 to 2009 utilizing a flexible Fourier stationary test to examine the mean reversion of consumption–income ratio in 16 OECD finding that 12 of the 16 consumption–income ratio series were stationary.

Fewer researchers have studied the stationarity of consumption-income ratio in developing countries, but again there are mix results. Gomes and Franchini (2009) examined the order of integration of the consumption-income ratio in 10 South American countries for the period 1951 to 2003 employing the individual ADF test, its panel versions and the Minimum LM unit root test with structural break(s). While the initial tests revealed more favourable evidence of an integrated process, upon factoring in structural breaks, only Uruguay appears to demonstrate integration. Consequently, overall, the consumption-income ratio was identified as a stationary process. Thus, in general, the consumption-income ratio was diagnosed as a stationary process. Baykara and Telatar (2012) analysed the stationarity properties of the consumption-income ratios for a sample of 14 transition economies by

taking account of nonlinearities and asymmetries together using the unit root tests based on the Threshold Autoregressive (TAR) models, providing evidence in favour of the stationary consumption-income ratios for all countries. Gozgor (2013) used heterogeneous panel unit root tests to account for cross-sectional dependence and the Modified Augmented Dickey-Fuller unit root test to investigate stochastic properties of the consumption-income ratios in 11 Central and Eastern European (CEE) countries over the period March 1997-September 2012, finding a strong mean-reversion in the consumption-income ratio for 9 of 11 CEE economies, and thus supporting the hypothesis that the consumption-income ratio converges towards a constant value. Solarin (2017) employed different types of nonstationarity tests showing that consumption–income ratio is non-stationary in most of the 10 ASEAN countries over the period 1970–2015.

The objectives of China's "*new normal*" development model are very ambitious. The model is based on the notion of better quality growth, and apart from the rebalancing from exports and investment towards domestic consumption, it emphasizes innovation as a means to foster higher value added activities, reducing inequalities and environmental sustainability (Green and Stern 2015). However, the transition to a more sustainable growth model is not going to be easy. Using cross country regression models, Pritchett and Summers (2014), Barro (2016) and Lee (2017) project a significant slowdown of China's growth rate. Perkins (2023) considers that there is little reason to think that the share of household consumption as a share of GDP will rise significantly in the next decade or two; therefore, in order to attain sustained economic growth, this author proposes the government to shift investment away from energy-intensive housing and infrastructure and toward investment in health care and education.

Several authors have signalled that more reforms are required in order to avoid the middle-income trap, including measures to foster the liberalization of the financial system, the improvement of education and research capabilities and a well designed public rental policy (Dorrucci et al. 2013; Yiping 2016; Thia 2019; Shao et al. 2022; etc.). Felipe and Lanzafame (2020) use the concept of the balance-of-payments equilibrium (BOPE) growth rate put forward by Thirlwall (1979), finding that China's long-run growth performance "will depend on the features of its switch from export-led growth to a new growth model, reflected in the ongoing structural change and rebalancing strategy. In particular, if the expected transition towards a consumption-driven economy occurs at a steady pace" (Felipe and Lanzafame 2020: 179). Xu (2010) considers that it is possible to stimulate domestic demand and service sector jobs thanks to a higher income to labor ratio if China "*... export less*, *but sell its export at higher prices...* " (Xu 2010: 32). Ma et al. (2017) nuance that the task is not going to be easy because of the knife-edged nature of Chinese rebalancing, concluding that for a rapid rebalancing, consumption needs to display exceptional resilience, and there must be a significant increase in the efficiency of investment.

In this paper we investigate Chinese households' consumption persistence and trends to determine whether there is a differential statistical behaviour compared to G7 and BRI economies. Our paper contributes to previous research firstly in the methodology used, which is based on fractional integration rather than other more standard methods that only consider integer degrees of differentiation. Note that in

this context a series may be nonstationary though mean reverting if the order of integration of the series is in the interval [0.5, 1). A second contribution is the number of countries examined, focussing on the comparison between China, G7 and BRI.

3 Methodology and data

We focus on fractional integration. This is a very flexible methodology that allows us to consider fractional orders of differentiation or I(d) behaviour in the series of interest. In fact, by allowing for such an approach we can consider a variety of modelling specifications, including the case of I(0) stationarity (or short memory) if $d=0$, and long memory models if $d>0$. Moreover, we can also consider nonstationary though mean reverting processes, if the order of integration d is in the interval [0.5, 1). Nonstationary, non-mean-reverting patterns are obtained if $d\geq 1$.

The estimation of the differencing parameter d is carried out by using the Whittle function in the frequency domain as proposed in Dahlhaus (1989) and presented in a testing-procedure form in Robinson (1994). Robinson's (1994) tests are very general, including cases with multiple orders of integration at different frequencies in the series. In this work we use a version of his tests, widely employed in empirical applications, and which functional form can be found, for example, in Gil-Alana and Robinson (1997). One of the advantages of this method is that is not restricted to be in the stationary range (i.e., $d<0.5$) and thus we do not need preliminary differentiation if we believe the series is nonstationary. This is in fact one of the distinguishing features of this approach compared with other methods like Sowell's (1992) maximum likelihood approach or the classical semiparameteric method of Geweke and Porter-Hudak (1983). Note that in these cases, if we believe the series is nonstationary, preliminary differentiation is required, adding then the value 1 to the estimated number to get the proper order of integration of the series. In Robinson (1994) the method is conducted for any real value d, including thus those outside the stationary range ($-0.5\leq d<0.5$).

The model examined in the empirical section is the following one,

$$y_t = \beta + \gamma t + x_t, \quad t = 1, 2, \ldots \tag{1}$$

$$(1 - L)^d x_t = u_t, \quad t = 1, 2, \ldots \tag{2}$$

where y_t is each of our observable series of interest; β and γ are unknown parameters to be estimated and corresponding to a constant and a (linear) time trend; d is the order of integration of the series, which may be any real value and thus potentially fractional, and u_t is the I(0) error term that will be modelled in terms of both white noise and autocorrelation, in the latter case, by means of the exponential spectral model of Bloomfield (1973). The use of the latter approach is quite convenient in the sense that approximates the behaviour of stationary AutoRegressive (AR) structures with very few parameters. It is in fact a non-parametric method which is described in terms of its spectral density function, which logged form is very similar to that

produced by an AR model.[1] In the context of exogenous shocks the crucial value is 1 since order of integration strictly smaller than 1 will imply mean reversion, while this property will not be satisfied if d is equal to or higher than 1.

For the purpose of this analysis, we use yearly data between 1960 and 2021 of "*Households and NPISHs*[2]*final consumption expenditure (% of GDP)*" indicator included in the World Development Indicators database, which is obtained from World Bank national accounts data, and OECD National Accounts data files. The World Bank defines this indicator as "*the market value of all goods and services, including durable products (such as cars, washing machines, and home computers), purchased by households…*", and, besides, "*… includes the expenditures of non-profit institutions serving households*". This database is probably the best one to provide an estimate of household consumption over a long period of time for all the countries included in our analysis: China, G7 economies (Canada, Germany, France, United Kingdom, Italy, Japan and the United States), and the three BRI economies (Brazil, Russia and India). Nevertheless, the World Bank admits relevant limitations that affect the consistency and completeness of the series: many countries generate their primary estimate of GDP using the production approach and household final consumption expenditure is often estimated as a residual, by subtracting all other known expenditures from GDP.

4 Empirical results

Prior to the analysis of fractional integration we first conducted classical unit root tests on the series. In particular, we performed Dickey and Fuller (ADF, Dickey and Fuller 1979), Phillips and Perron (PP, Phillips and Perron 1988) and Elliot et al. (ERS, Elliot et al. 1996) tests and the results, though not reported, supported the unit root case in all series examined. Nevertheless, it should be taken into account that these tests have very low power if the true data generating process is in fact fractionally integrated with an order of integration different from 1. In other words, even if the tests support the unit root hypothesis, this conclusion might be invalid if $d \neq 1$ (Diebold and Rudebusch 1991; Hassler and Wolters 1994; Lee and Schmidt 1996).

Based on the above, we consider the model given Eqs. (1) and (2), firstly, assuming that the error term u_t in (2) is uncorrelated, i.e., following a white noise process, and then imposing autocorrelation throughout the model of Bloomfield (1973). In the two cases we examine three models: (i) including no deterministic terms, i.e., imposing that $\beta = \gamma = 0$ a priori in Eq. (1); (ii) including an intercept, i.e., with $\gamma = 0$ a priori; and (iii) with a constant and a linear time trend. Table 1 displays the estimates of d in the three models under the assumption of white noise errors, marking in bold the selected model according to the deterministic terms; Table 2 focussed on the estimated coefficients of the selected model for each series; Tables 3 and 4 are similar to Tables 1 and 2 but imposing autocorrelated (Bloomfield) disturbances.

[1] See Gil-Alana (2004) for the convenience of the model of Bloomfield (1973) for autocorrelation in the context of fractional integration.

[2] Non-profit institutions serving households.

Table 1 Estimates of the order of integration. White noise errors

Series	No terms	An intercept	An intercept with a linear time trend
BRAZIL	0.93 (0.79, 1.11)	**0.62 (0.45, 0.86)**	0.62 (0.41, 0.86)
CANADA	0.88 (0.72, 1.11)	**1.10 (0.83, 1.46)**	1.09 (0.89, 1.40)
CHINA	1.19 (0.98, 1.53)	**1.20 (0.67, 2.08)**	1.21 (0.30, 2.09)
FRANCE	0.93 (0.76, 1.16)	**1.05 (0.83, 1.32)**	1.05 (0.87, 1.31)
GERMANY	0.92 (0.75, 1.16)	**0.90 (0.77, 1.11)**	0.90 (0.74, 1.12)
INDIA	0.94 (0.79, 1.15)	0.96 (0.79, 1.27)	**0.96 (0.79, 1.25)**
ITALY	0.91 (0.73, 1.18)	**0.79 (0.53, 1.10)**	0.79 (0.52, 1.10)
JAPAN	0.93 (0.75, 1.19)	**1.09 (0.86, 1.40)**	1.08 (0.87, 1.38)
RUSSIA	0.68 (0.39, 1.01)	**0.38 (0.09, 1.01)**	0.28 (-0.18, 1.02)
UK	0.92 (0.76, 1.16)	0.85 (0.51, 1.19)	**0.89 (0.69, 1.18)**
US	0.90 (0.72, 1.17)	0.88 (0.76, 1.13)	**0.84 (0.65, 1.12)**

The values in bold refer to the selected specification for each series in relation with the deterministic terms. Values in parenthesis are 95% confidence bands

Table 2 Estimated coefficients of the selected models. White noise errors

Series	d	Intercept (t-value)	Time trend (t-value)
BRAZIL	0.62 (0.45, 0.86)*	62.194 (23.11)	---
CANADA	1.10 (0.83, 1.46)	63.621 (67.44)	---
CHINA	1.20 (0.67, 2.08)	47.771 (16.29)	---
FRANCE	1.05 (0.83, 1.32)	57.241 (126.74)	---
GERMANY	0.90 (0.77, 1.11)	55.632 (73.24)	---
INDIA	0.96 (0.79, 1.25)	87.779 (63.05)	-0.4618 (-3.02)
ITALY	0.79 (0.53, 1.10)	59.007 (99.99)	---
JAPAN	1.09 (0.86, 1.40)	46.519 (64.54)	---
RUSSIA	0.38 (0.09, 1.01)	50.441 (38.77)	---
UK	0.89 (0.69, 1.18)	74.222 (68.18)	-0.1963 (-2.11)
US	0.84 (0.65, 1.12)	60.042 (120.77)	0.1521 (3.74)

*: Evidence of mean reversion at the 95% level

We start by displaying the results under the assumption of white noise errors. We notice that the time trend is required in the cases of India, the UK and the US. For the remaining series, the intercept is sufficient to describe the deterministic terms. Nevertheless, these trends are irrelevant from a statistical viewpoint based on the

Table 3 Estimates of the order of integration. Autocorrelated errors

Series	No terms	An intercept	An intercept with a linear time trend
BRAZIL	1.06 (0.77, 1.50)	**0.58 (0.25, 1.13)**	0.49 (-0.17, 1.13)
CANADA	0.76 (0.46, 1.17)	**0.90 (0.26, 1.41)**	0.93 (0.63, 1.33)
CHINA	0.79 (0.51, 1.22)	0.29 (0.11, 0.54	**-0.33 (-0.84, 0.21)**
FRANCE	0.82 (0.48, 1.20)	0.31 (-0.12, 1.37)	**0.81 (-0.03, 1.32)**
GERMANY	0.81 (0.43, 1.24)	0.74 (0.46, 1.07)	**0.68 (0.27, 1.07)**
INDIA	0.87 (0.62, 1.25)	0.77 (0.64, 1.06)	**0.67 (0.33, 1.05)**
ITALY	0.74 (0.34, 1.19)	0.31 (-0.02, 1.06)	**-0.08 (-0.47, 1.06)**
JAPAN	0.78 (0.32, 1.22)	0.58 (0.35, 1.06)	**0.46 (-0.24, 1.03)**
RUSSIA	-0.06 (-0.37, 1.17)	-0.40 (-0.94, 0.44)	**-1.01 (-1.24, 0.11)**
UK	0.77 (0.46, 1.18)	0.24 (0.00, 0.51)	**0.43 (-0.03, 0.92)**
US	0.75 (0.06, 1.23)	0.82 (0.63, 1.24)	**0.68 (0.17, 1.19)**

The values in bold refer to the selected specification for each series in relation with the deterministic terms. Values in parenthesis are 95% confidence bands

large degree of persistence of the data. [3] Looking at the differencing parameter for the selected models we first underline that the PRC displays a value above 1 [1.20 (0.67, 2.08)]). Similarly, in most of the remaining countries, the values are around 1. Thus, for China, we observe a large degree of integration; in fact, it is the largest value for all countries examined in the sample, and also the confidence interval is extremely wide, including also the value of 2 and thus the I(2) hypothesis cannot be rejected. Thus, the series of China displays the highest degree of persistence, being unable to reject the I(1) and the I(2) hypotheses.

Mean reversion (i.e., statistical evidence of $d<1$) is only detected in the case of Brazil. For this country, the estimated value of d is equal to 0.62 and all the values in the confidence intervals are smaller than 1, supporting thus the hypothesis of reversion to the mean. Thus, shocks in this series are expected to be transitory, disappearing in the long run. There is another country that also displays a significant value below 1, Russia (0.38); however, for this country, the confidence interval is also too wide and the I(1) hypothesis cannot be rejected. Note that the time trend coefficients are negative in the cases of India and the UK but positive for the US. Explaining the reasons for the value of each time trend coefficient may be complex and goes far beyond the objectives of this paper, but it is worth mentioning that the evolution of factors such as access to credit, income level, (de)-industrialization, etc. might be behind these coefficients (Kim et al. 2015; Mishra 2011; etc.). Nevertheless, remem-

[3] In fact, with $d=1$ for example, the time trend disappears in (1) and (2) and, for $t>1$, the model becomes a random walk with a drift.

Table 4 Estimated coefficients of the selected models. White noise errors

Series	d	Intercept (t-value)	Time trend (t-value)
BRAZIL	0.58 (0.25, 1.13)	63.560 (30.11)	---
CANADA	0.90 (0.26, 1.41)	63.181 (67.20)	---
CHINA	-0.33 (-0.84, 0.21)*	61.184 (114.22)	-0.4405 (-25.64)
FRANCE	0.81 (-0.03, 1.32)	57.256 (129.13)	-0.0677 (-2.34)
GERMANY	0.68 (0.27, 1.07)	56.440 (76.77)	-0.1021 (-2.71)
INDIA	0.67 (0.33, 1.05)	87.041 (62.26)	-0.4876 (-8.05)
ITALY	-0.08 (-0.47, 1.06)	87.851 (409.20)	0.0226 (4.86)
JAPAN	0.46 (-0.24, 1.03)	48.266 (71.82)	0.1557 (6.52)
RUSSIA	-1.01 (-1.24, 0.11)*	48.289 (420.05)	0.1399 (16.79)
UK	0.43 (-0.03, 0.92)*	72.587 (80.86)	-0.1624 (-6.30)
US	0.68 (0.17, 1.19)	59.961 (122.49)	0.1622 (6.37)

*: Evidence of mean reversion at the 95% level

ber that these coefficients tend to zero and therefore lose its importance if the order of integration of the series is equal to or higher than 1.

Tables 3 and 4 refer to the model with autocorrelated errors. We first observe that the time trend coefficient is now significant and negative for China. It is also significant and negative for France, Germany, India and the UK, and positive for Italy, Japan, Russia and the US. However, it is not significant for Canada or Brazil. More importantly, and focussing on d, mean reversion is found to be statistically significant (i.e. values of d below unity in the confidence intervals) in the PRC (d = -0.33), which supports the hypothesis of reversion to the mean, and it is also significant in Russia (-1.01) and the UK (0.43). For the rest of the countries, the unit root null hypothesis (d = 1) cannot be rejected, and the confidence intervals are so wide in some countries that neither hypothesis (d = 0 and d = 1) can be rejected (France, Italy and Japan). These wide confidence intervals are clearly related to the small number of observations used in the application. Thus, China presents some of the most striking results since it displays the highest degree of integration under the assumption of white noise errors, but the lowest one (along with Russia) if autocorrelation is permitted in the error term.

5 Discussion of results

The results are heterogeneous across countries, observing orders of integration statistically below 1 only in a few countries. In the case of China, we find that the hypothesis of reversion to the mean is supported in the model with autocorrelated errors, which is more plausible since we allow for a further structure in the model. Accord-

ing to these results, shocks in the series of household consumption as a share of GDP are expected to be transitory, disappearing in the long run. This result has significant consequences, since it means that Chinese authorities need to implement a collection of particularly strong public measures in order to increase the relevance of domestic consumption in the long run and obtain significant progress in the process towards China's "*new normal*" development model.

When autocorrelated errors are permitted, we find similar results only in Russia and the UK. In the remaining countries, the values are around or statistically significantly above 1. Under the assumption of white noise errors, mean reversion is detected only in Brazil. Thus, considering all the models and hypothesis, mean to reversion is found in China and two of the other BRIC economies (Russia and Brazil), whereas only in one of the seven G7 economies (the UK) do we detect this result. Therefore, according to this analysis based on fractional integration, in the long run, household consumption seems to be less persistent in the BRIC economies than in the G7 countries.

These results have important policy implications. We have evidenced that the Chinese authorities will have to adopt a set of enhanced measures to increase the share of household consumption in GDP. Initiatives to increase the spending on basic public services (Xiong et al. 2022), reduce income inequality (Anand and Chhikara 2013), and promote digital inclusive finance (Li et al. 2020) should be implemented to this end, particularly in a context in which the population needs to regain confidence after the severe measures adopted during the COVID-19 pandemic.

6 Conclusions

In this article we have examined the stochastic structure of China's household consumption level by looking at its degree of integration and comparing it with the G7 and BRI countries. The overall results indicate high levels of persistence though they are very heterogeneous across countries and depending on the model used. In general, evidence of mean reversion, and thus transitory shocks with temporal effects are obtained in China along with two BRI countries (Brazil and Russia), while this hypothesis is only supported for UK in the G7 economies. This statistical pattern means that it is more complex to permanently increase the share of domestic consumption in China than in the G7 economies and, in consequence, that the PRC will have to implement a sound package of initiatives to boost domestic consumption and obtain significant progress in its shift towards the inspirational "*new growth model*".

Admittedly, China's national accounts could be underestimating consumption data, thus magnifying the problem of insufficient household consumption. Nevertheless, even in this case, if underestimating errors were permanent and of the same magnitude, our results would be relevant, considering that it is certain that Chinese authorities are fully engaged with the objective of increasing domestic consumption from its current level.

This paper has some important limitations. First, the sample size uses annual data starting at 1960 and thus comprises only 61 observations per series. On the other hand, being historical data, the presence of structural breaks is an issue that should

be analysed, though once again a limitation occurs here due to the small sample sizes of the subsamples. Future studies may analyze the degree of persistence of households' consumption using alternative approaches, based for example on non-linear trends (e.g., Chebyshev polynomials in time, Cuestas and Gil-Alana 2016; Fourier functions, Gil-Alana and Yaya 2021; or even neural networks, Yaya et al. 2021) implemented on these or alternative data sets. Additionally, it would be interesting to examine in a more detailed way whether the pattern of households' consumption in fast growing developing economies present special and distinctive persistent features. Work in this direction is now in progress.

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Data availability We use yearly data between 1960 and 2021 of "Households and NPISHs final consumption expenditure (% of GDP)" indicator included in the World Development Indicators database, which is obtained from World Bank national accounts data, and OECD National Accounts data files. The data are publicly available in https://databank.worldbank.org/source/world-development-indicators.

Declarations

Conflict of interest The authors report there are no competing interests to declare.

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