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Do current and capital account liberalizations affect economic growth in the long run?

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Abstract
We investigate the long-run relationship between de jure trade openness and economic growth as well as between de jure financial openness and economic growth for a panel of 118 developed and developing countries in the period 1970–2017. We fit a cross-sectional autoregressive distributed lag model and unveil the positive association between both liberalizations and economic growth in the global panel and for the more developed countries. Conversely, only trade liberalization is linked with larger output growth in less developed nations. We complement these results with a long-run causality analysis, which reveals that for the whole sample and for the two subsamples both de jure trade and de jure financial openness jointly cause economic growth. These outcomes may be indicative that during the time span under scrutiny developing countries faced current account crises when they stuck to the early prescriptions of the Washington Consensus. Yet, they later adopted a more nuanced view of economic liberalism putting in place a number of capital controls, which protected them from sudden stops and capital reversals.

Keywords Trade liberalization · Capital liberalization · Capital controls · Economic growth · Washington Consensus

JEL classification F13 · F14 · F38 · F41 · F43

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1 Introduction

Do current and capital account liberalizations promote economic growth? This question has been pivotal in policy debates since the 1980s when a multi-year debt crisis sunk several Latin American countries into a "lost-decade" of economic slump and called for a change in the dominant development paradigm away from state dirigisme. Free trade and larger foreign direct investment (FDI) have been also advocated by the so-called Washington Consensus in 1989 when the fall of the Berlin Wall marked the burial of centrally planned economies (Williamson 1990, 2004). At first, opinions of economists were quite positive and earlier analyses supported the belief that trade openness is associated with better economic outcomes in terms of faster growth, higher productivity, larger investment and lower income inequality (Dollar 1992; Ben-David 1993; Sachs and Warner 1995; Edwards 1998; OECD 1998; Frankel and Romer 1999). However, in their in-depth critical review of the literature, Rodríguez and Rodrik (2000) question such findings and claim that "the strong results in this literature arise either from obvious misspecification or from the use of measures of openness that are proxies for other policy or institutional variables that have an independent detrimental effect on growth" (Rodríguez and Rodrik 2000, p. 315).

The profound institutional reforms implemented in several countries during the 1990s, also due to unexpected changes such as the collapse of the communist bloc, spurred renewed interest in this topic. Irwin (2019) surveys the field focusing on how changes in a country's own barriers affect its economic growth. According to this author the latest literature consistently shows that developing countries can largely benefit from more liberal trade policies supporting the prescriptions of the Washington Consensus even if the final outcome is quite heterogeneous across countries (Greenaway et al. 2002; Aksoy and Salinas 2006; Wacziarg and Welch 2008; Falvey et al. 2012, 2013; Estevadeordal and Taylor 2013; Feyrer and Irwin 2019; Grier and Grier 2020).

While trade liberalization is likely to be beneficial on standard comparative-advantage grounds, there exists a diversity of opinions regarding the consequences of capital account liberalization because short- and long-term flows may have different growth effects. Some authors do not find any strong correlation between capital account liberalization and growth (Grilli and Milesi-Ferretti 1995; Rodrik 1998; Eichengreen 2001; Edison et al. 2002; Andersen and Tarp 2003; Rodrik and Subramanian 2009; Adams and Klobodu 2018; Njikam 2017; Furceri et al. 2019b) while others assert that capital account liberalization can improve significantly the domestic growth rate (Levine and Zervos 1998; Chan-Lau and Chen 2001; Chinn and Ito 2002; Bekaert and al. 2005; Henry 2007; Klein and Olivei 2008; Hassan et al. 2011; Akinsola and Odhiambo 2017; Choi and Pyun 2019).

This paper enriches the empirical literature that addresses the long-run relationship between current and capital account liberalizations on the one hand, and economic growth, on the other hand. The contributions of the paper are many-fold. First, we provide evidence that both liberalizations accrue growth in the long run, using a large balanced panel dataset comprising 118 economies from 1970 to 2017. Second, we address how different stages of development may influence the growth-liberalization nexus. We show that both links are effective in more developed economies, while in low- and lower middle-income countries only trade liberalization positively impacts income. Third, we employ a robust econometric approach to tackle country heterogeneity and dependence of unobservables across countries, i.e., cross-sectional dependence (CSD), in a dynamic setting. Cross-sectional dependence is frequent in panel data like ours and may emerge from a number of reasons, such as omitted common effects, spatial effects, spillover effects, lack of dynamics or the erroneous pooling of units. We care about CSD because conventional panel estimators may lead to inconsistent estimates and thus distorted inference depending on the extent of CSD. Fourth, we shed new and methodologically robust evidence on the causal relationship between growth and liberalization for the above-mentioned panels using the long-run Granger-causality approach (Eberhardt and Teal 2013).

The remaining of the paper is organized as follows. The second section briefly discusses the relevant literature, while the third section illustrates the dataset and introduces the empirical methodology. The forth section describes the results, whereas the last one concludes and provides policy implications.

2 Related literature

Since the emergence of the Washington Consensus the economic discipline had to face the problem of devising the most accurate and appropriate tools to measure the extent of policies aimed at current and capital account liberalizations as well as the linkages between such policies and economic performance.

The policy levers governments adopted in the spirit of the Washington Consensus may not necessarily translate into the desired policy outcomes because the latter may be affected by non-policy factors, such as adverse exogenous macroeconomic shocks. Despite we acknowledge the literature takes stock of such diverse outcomes to assess the effectiveness of economic policy interventions (e.g., Easterly 2019), in this work we will only concentrate on how the change in current and capital account liberalization policies has a long-run impact on output growth. Our interest lies exclusively on how normative changes impact on economic growth, and not on other outcome measures.

The literature suggests that it is possible to look at current and capital account liberalization either via de facto or de jure measures (Gräbner et al. 2020). De facto measures concern the actual degree of a country's integration into world trade and capital flows, whereas de jure measures regard a country's intention to open its current and capital account relying on its regulatory environment. Abatement in tariffs rates or non-tariffs barriers are examples of de jure measures of current account liberalization. Lifting capital account restrictions and laxer capital controls are instead examples of de jure measures of capital account liberalization. Since the mid-1990s, the scholarship started to elaborate a number of proxies of both de jure trade and current account openness measures. Scholars used them to observe the relationship with economic growth and they came up with mixed evidence, which changed according to the set of countries under scrutiny, the time span, and the adopted proxy. Taking stock of such premises, in the next subsections we provide a selected review of the literature that addresses the relationship between trade or capital liberalization, on one side, and economic growth or standard of living, on the other side. A summary of these contributions is also reported in Table 1 from which we notice the variety of data sets and methodologies adopted.

2.1 Trade liberalization and economic growth

The list of contributions looking at the association between trade liberalization and economic growth dates back to Sachs and Warner (1995) who estimate a standard Barro's growth regression model adding a closed-economy dummy.[1] Their estimates show that open countries grow on average 2.45% more than closed ones and close countries have a lower per capita growth of about 1.2% per year. This result has been partially confirmed by Vamvakidis (2002), Wacziarg and Welch (2003) as well as Billmeier and Nannicini (2009) and Nannicini and Billmeier (2011). Wacziarg and Welch (2003) compare the performance of countries that liberalized with respect to those that did not and claim that countries that liberalized over the period 1950–1998 enjoyed annual GDP growth rates circa 1.5% higher than those achieved before liberalization. Nannicini and Billmeier (2011) find that trade liberalization (as represented by the updated Sachs–Warner indicator) had a positive impact on economic growth in four transition economies that liberalized their trade regime in the 1990s (Armenia, Azerbaijan, Georgia and Tajikistan), while Uzbekistan, that missed the opportunity to liberalize, paid a substantial cost in the medium-to-long run.

Another strand of literature addresses non-dichotomous measures of trade openness such as trade volumes or trade restrictions. These are de facto variables and encompass import duties (as a share of imports), average years of openness (as indicated by the Sachs and Warner index), as well as the difference between official and black market exchange rates. Measures of trade volumes indicate that there is a positive and significant association between trade openness and growth (Yanikkaya 2003) but small according to Lee et al. (2004) who take into account endogeneity in the data. Dejong and Ripoll (2006) claim that a 10% decrease in tariffs is associated with a 1.6% increase in annual GDP growth for richer nations, whereas the output change is negative, but often insignificant, among the world's poor countries. Yanikkaya (2003) finds results for trade barriers that contradict the conventional view. His specifications show a positive and significant relationship between trade barriers, such as the average import/export tariff rate or taxes on international trade, and growth. His results are essentially driven by less developed countries and thus consistent with the predictions of the theoretical growth literature that, under certain conditions, developing countries can actually benefit from trade restrictions. Mixed results are also obtained by Ulaşan (2015) who cannot provide support for a significant association between both de facto and de jure measures of trade openness and economic growth. However, Furceri et al. (2019a, 2020) estimate impulse response functions and show that tariff increases lead

[1] Sachs and Warner (1995) define a country closed when one of the following five conditions is met: (1) average tariff rates of 40% of more, (2) nontariff barriers covering 40% or more of trade, (3) a black market exchange rate at least 20% lower than the official exchange rate, (4) a state monopoly on major exports, (5) a socialist economic system.

Table 1 Summary of selected studies

Author	Countries	Years	Methodology
Trade liberalization and economic growth			
Sachs and Warner (1995)	79 countries	1970–1989	OLS, Logit
Vamvakidis (2002)	Between 54 and 89 countries	1870–1990	Extreme bounds analysis
Wacziarg and Welch (2003)	Between 27 and 136 countries	1950–1999	Difference in difference, SUR, Fixed effects
Yanikkaya (2003)	Between 52 and 114 countries	1970–1997	OLS
Lee et al. (2004)	About 100 countries	1961–2000	OLS, Identification through heteroskedasticity
Dejong and Ripoll (2006)	60 countries	1975–2000	OLS, SUR, Difference generalized Methods of moments, SYS-GMM
Billmeier and Nannicini (2009)	180 countries	1960–2000	Matching estimators, OLS
Ulaşan (2015)	130 countries	1960–2000	Difference generalized methods of moments, SYS-GMM
Furceri et al. (2019a)	151 countries	1963–2014	Impulse response, VAR, Instrumental variables
Furceri et al. (2020)	151 countries	1963–2014	Impulse response, VAR, Instrumental variables
Financial liberalization and economic growth			
Quinn (1997)	Between 58 and 64 countries	1960–1989	OLS
Arteta et al. (2001)	Between 47 and 141 countries	1973–1992	OLS, Weighted least squares, Instrumental variables
Honig (2008)	122 countries	1970–2005	OLS, Instrumental variables
Bussieré and Fratscher (2008)	45 countries	1980–2000	Difference generalized methods of moments, SYS-GMM
Quinn and Toyoda (2008)	94 countries	1950–2004	OLS, Instrumental variables, SYS-GMM
Misati and Nyamongo (2012)	34 countries	1983–2008	Pooled ordinary least squares, Fixed effects
Njikam (2017)	45 countries	1970–2010	SYS-GMM
Choi and Pyun (2019)	45 countries	1985–2012	OLS, Two-step SYS-GMM
Furceri et al. (2019b)	149 countries	1970–2015	OLS, Fixed and smooth transition autoregressive techniques

to economically and statistically significant declines in domestic output and productivity in the medium term. These effects tend to be magnified when tariffs rise during expansions, for advanced economies, and when tariffs go up, not down.

2.2 Financial liberalization and economic growth

Researchers not only have studied trade liberalization, but they also extensively investigated capital liberation relying on continuous, dummy or categorical variables. Quinn (1997) and Arteta et al. (2001) fit a standard growth model adding two variables drawn from IMF's Annual Report on Exchange Arrangements and Exchange Restrictions (AREAER) that measure the degree of financial openness and change in international financial regulation. Results point to a positive association between capital account liberalization and economic growth in the long run which is stronger in countries with a better rule of law. Honig (2008) adopts the Chinn-Ito index[2] to show that the de jure capital account liberalization has a positive impact on GDP growth, while there is little evidence that a better institutional quality strengthens this association. Klein and Olivei (2008) opt for the proportion of years in which the country had unrestricted capital mobility to show there is a positive association between capital account liberalization and financial depth which translates into economic growth only for highly industrialized economies.

Bussieré and Fratscher (2008) pinpoint that the lack of a robust evidence in the relationship between financial openness and economic growth lies on the different time impact that liberalization has on economic growth, which, they argue, leads to gains in the short-run and to losses in the long run. To shed light on this issue, they opt for the Schmuckler and Kaminsky (2003)'s dataset who classify countries on three degrees of liberalization, i.e., "partially liberalized," "fully liberalized" and "closed." Bussieré and Fratscher (2008) thus create a dummy which takes the value of 1 if the capital account is open for the majority of the five year periods, and zero otherwise. Their results show that countries tend to grow faster in the five years following the capital account liberalization, but afterward the path of the GDP is not statistically different from that of other times. Misati and Nyamongo (2012) embrace a financial liberalization dummy, which takes a value of 1 in the period subsequent a financial liberalization, and the Chinn-Ito measure of capital account liberalization. Results of pooled and fixed effects regressions show an ambiguous relationship, since the coefficient of the capital account liberalization is positive and significant, while the one associated with financial liberalization is not statistically different from zero. On the contrary, Quinn and Toyoda (2008) develop two categorical measure of financial liberalization. The first takes values between 0 and 4 where increasing values indicate a more open economy to capital flows. The second is "an indicator of how compliant a government is with its obligations under the IMF's Article VIII to free from government restriction the proceeds from international trade of goods and services" (Quinn

[2] This index aims at measuring the intensity of capital controls, and it is the first standardized principal component of AREAER binary variables on: (1) the existence of multiple exchange rates, (2) restrictions on current account, (3) capital account transactions, and (4) a variable indicating the requirement of the surrender of export proceeds (Chinn and Ito, 2006).

and Toyoda 2008, p. 1409) and takes values from 0 to 8 where 8 indicates complete compliance. Estimation results show that both measures of liberalization have a positive impact on economic growth. More recently, Njikam (2017) finds no significant effect of the Chinn-Ito liberalization dummy on growth, while domestic and external financial liberalization positively impacts economic growth only if coupled with improvements in complementary reform variables. Also Choi and Pyun (2019) show that capital controls alone do not have a significant impact on economic growth. Conversely, the presence of capital controls coupled with external savings and an increase in net exports lead to an increase in output in the manufacturing sector. Furceri et al. (2019b) assess the impact of capital account liberalization on income distribution as well as GDP growth relying on a much larger and longer panel then their previous peers. They opt for a dummy that is equal to one at the start of capital account liberalization episode, and zero otherwise. The episode occurs when the Chinn-Ito index exceeds by two standard deviations the average annual change over all observations. Interestingly, Furceri et al. (2019b) claim that liberalization episodes reduce the share of labor income but cannot find any significant relationship between them and changes in GDP growth.

2.3 A reappraisal on the association between de jure trade and financial liberalization and growth: the KOF measure of globalization

The last three decades saw a proliferation of measures of real and financial openness, both under the de jure and the de facto umbrella. This variety surely enriched the discipline, but also created a certain degree of confusion, since there is now a lack of a straightforward indication on which is the most appropriate proxy to employ. Gräbner et al. (2020) recently acknowledge the variety of measures of real and financial openness and provide a comprehensive review of the existing de jure and de facto proxies of economic openness. They analyze the correlation among all the measures under their scrutiny in levels and first differences for 216 countries over the period 1965–2019. They notice that de jure trade and de jure financial openness measures are more closely correlated than their de facto counterparts, and that the correlation between de facto and de jure measures is rather low. They claim that such a result is indicative that when countries decide to remove barriers on trade and financial flows they do so at the same time.

Gräbner et al. (2020) use the Swiss Economic Institute (KOF) updated dataset about economic globalization which includes both trade and financial globalization in their de facto and the jure forms (Gygli et al. 2019). The de jure trade globalization index is a sub-dimension of trade globalization and refers to policies that foster trade between countries. It is measured using time-varying weighted values of variables on trade regulations, trade taxes, tariff rates and free trade agreements. The de jure financial globalization index is a sub-dimension of financial globalization and it aims at assessing the degree of a country's openness to international flows and investments. It is measured using time-varying weighted values of the Chinn-Ito index, of a variable quantifying investment restrictions based on the World Economic Forum Global Competitiveness Report, and of the number of international investment agreements.

The KOF economic globalization index and its sub-indexes take values between 0 and 100, with increasing numbers indicating higher degrees of globalization and cover 203 countries for a time span between 1970 and 2018.[3]

The overall KOF economic globalization index is correlated with most of the measures studied by Gräbner et al. (2020), an outcome supportive of KOF's capability to comprehensively depict economic openness. Furthermore, Gräbner et al. (2020) run a growth regression relying on data for 65 countries over the period 1995–2014, specifying it into two different versions. In one, all the variables (exception made for the initial level of GDP) are expressed as a five-year average, while in the other the variables are expressed in first differences. More specifically, the dependent variable is the growth rate of the GDP per capita while the independent variables include the initial level of GDP per capita and, in turn, each measure of openness reviewed in the paper, plus country-fixed effects and a set of control variables. As far as KOF is concerned, Gräbner et al. (2020) first employ the overall KOF measure of economic globalization, but also run a set of regressions in which they subsequently split the overall index into its components, i.e., trade and financial openness, then de facto and de jure openness, then trade openness de facto, trade openness de jure, financial openness de facto, financial openness de jure. The results of the last most disaggregated regression show that trade openness de jure has a positive impact on economic growth both when employed as a five-year average and in the first differences specification. Conversely, financial openness de jure is not statistically associated with GDP growth.

We agree with Gräbner et al. (2020) on the comprehensive nature of the KOF index thanks to its conceptual clarity, transparency and its unique modular structure that allows the researcher to investigate specific aspects of economic globalization through the use of sub-indices. Yet our paper differs from Gräbner et al. (2020) and from all the reviewed studies who allegedly analyze the long-run effect of trade and current account liberalization on growth. In fact, only the autoregressive distributed lag model (ARDL), with its error correction reparameterization, allows to describe an inherently long-run non-stationary process as growth typically is. To the best of our knowledge, this paper is the first work to make use of an error correction mechanism to describe the long-run association between the de jure KOF trade openness and economic growth as well as the long-run association between the de jure KOF financial openness and economic growth. Moreover, we tackle possible reverse causality using the long-run Granger-causality approach since better standard of living can demand more open economies.

3 Methodology

In order to assess the long-run impact of trade and capital account liberalization on economic growth, we make use of data about 118 countries over the period 1970–2017. We analyze the data for the sample as a whole and for two subsamples that is the low- and lower middle-income panel as well as the upper middle and high income panel

[3] For a detailed description on the process to define the weight for the sub-indexes, see Gygli et al. (2019, p. 559).

Table 2 Summary statistics

Variable	Obs	Mean	S.D.	Min	Max
All countries					
gdp	5664	10.921	1.999	5.943	16.728
tra	5664	− 0.917	0.588	− 2.998	− 0.022
fin	5664	− 0.793	0.532	− 3.755	− 0.017
pri	5664	− 1.038	0.637	− 3.765	0.756
Low- and lower middle-income countries					
gdp	2592	9.937	1.593	6.099	15.945
tra	2592	− 1.294	0.485	− 2.998	− 0.258
fin	2592	− 1.029	0.545	− 3.755	− 0.100
pri	2592	− 1.328	0.503	− 3.765	0.282
Upper middle- and high-income countries					
gdp	3072	11.751	1.929	5.943	16.728
tra	3072	− 0.598	0.467	− 2.278	− 0.022
fin	3072	− 0.594	0.431	− 2.451	− 0.017
pri	3072	− 0.794	0.637	− 3.193	0.756

All variables are in logs

following the classification provided by the World Bank.[4] The selected time span is the longest on which we can rely on to exploit at its best the data availability for the variables under scrutiny. These are the de jure KOF trade and financial globalization indices, and real GDP at chained PPP (2011 prices) available in the Penn World Tables version 9.1. (Feenstra et al. 2015). We control for one macroeconomic effect using the price level of output side of real GDP.[5] These data are also available in the Penn World Tables version 9.1 for the same time span and the same number of countries of the two main covariates (Feenstra et al. 2015). The inclusion of this explanatory variable (and of its change over time, i.e., inflation) in the model is consistent with selected contributions included in Table 1, which concentrated either on trade liberalization and growth (Lee et al. 2004) or financial liberalization and growth (Honig 2008; Bussieré and Fratscher 2008). Table 2 reports the summary statistics of (logs of) the real GDP (*gdp*), de jure KOF trade globalization (*tra*), de jure KOF financial globalization (*fin*), and the price level (*pri*).

We first check whether the variables under investigation are non-stationary and affected by CSD. To reach this aim, we make use of Pesaran's (2015) CD test together with the exponent of CSD (Bailey et al. 2016; Ditzen 2021). The null hypothesis of the CD test states the variables are weakly cross-sectional dependent. Bailey et al. (2016)

[4] We do not split the sample further to avoid an excessive reduction of the N dimension of the panels. That can hamper the accurateness of the results stemming from our econometric approach which builds upon large N and large T.

[5] The price level is equal to the PPP (ratio of nominal GDP to constant GDP) divided by the nominal exchange rate with the price level of USA output in 2011 equal to unity. These indices show how output price levels differ across countries (Feenstra et al. 2015).

expand the analysis as they show the exponent α measures the degree of CSD. Values of $\alpha < ½$ indicate different degrees of weak CSD, whereas values of $\alpha \geq ½$ suggest strong CSD. After proving that the series under scrutiny are affected by CSD, we adopt second generation unit root tests to define the order of integration of the covariates. We opt for the cross-sectional Im, Pesaran, and Shin test, called CIPS (Pesaran 2007; Pesaran et al. 2013). Moreover, we allow for individual lag structures according to two alternative criteria which are the Portmanteau test of white noise or the Wald test of a compound linear hypothesis on the model parameters (Burdisso and Sangiacomo 2016). This CIPS statistic has a nonstandard distribution and corresponding critical values are tabulated in Pesaran (2007). In this paper, we focus on the rate of change in liberalization indices and output, while controlling for inflation. This choice is due to two main reasons. First, Irwin (2019) observes that the positive cross-country relationship between the level of a liberalization indicator, say tariffs, and economic growth, could simply reflect the fact that less developed countries (LDC) have higher tariffs than richer countries and have also tended to grow faster than their more developed peers. LDC, despite the higher level of their tariffs, may have been reducing them quicker than high-income countries and growing faster as a result. Hence, by using the level of tariffs rather than the change in tariffs, these studies are not examining within-country growth as a result of liberalization policies. Second, we opt for the Common Correlated Effects Mean Group (CCEMG) estimator which can be applied to stationary heterogeneous panels with weakly exogenous regressors where the cross-sectional dimension (N) and the time series dimension (T) are sufficiently large (Chudik and Pesaran 2015). This approach considers an autoregressive distributed dynamic panel specification, ARDL(P, Q_c, Q_k, Q_p) with the following factor error structure[6]:

$$y_{i,t} = \alpha_i + \sum_{j=1}^{L} \gamma_{i,j} y_{i,t-j} + \sum_{j=0}^{Q_c} \beta_{i,j}^{c} c_{i,t-j} + \sum_{j=0}^{Q_k} \beta_{i,j}^{k} k_{i,t-j} + \sum_{j=0}^{Q_p} \beta_{i,j}^{p} p_{i,t-j} + u_{i,t} \tag{1}$$

$$u_{i,t} = \zeta_i f_t + \varepsilon_{i,t} \tag{2}$$

$$c_{i,t} = a_i^c + \psi_i^c f_t + v_{i,t}^c,\ \ k_{i,t} = a_i^k + \psi_i^k f_t + v_{i,t}^k,\ \ p_{i,t} = a_i^p + \psi_i^p f_t + v_{i,t}^p \tag{3}$$

where α_i are the country-specific fixed effects to control for country factors that are stable over time, $y_{i,t}$ is the growth rate of GDP in country i at time t, $c_{i,t}$ is the rate of change in the current account liberalization index in country i at time t, and $k_{i,t}$ is the rate of change in the capital account liberalization index in country i at time t, and $p_{i,t}$ is the rate of change in the price level in country i at time t. The error term $u_{i,t}$ encompasses unobservables which include m common factors f_t. Vectors ζ_i, ψ_i^c, ψ_i^k *and* ψ_i^p are factor loadings. ε, v^c, v^k, v^p are assumed to be uncorrelated idiosyncratic error terms. We can then reparametrize this model into the familiar error correction model (ECM):

[6] It is worth stressing that a model that neglects heterogeneity, common factors and dynamics may lead to unreliable estimates (Eberhardt and Teal, 2011; Eberhardt et al. 2013).

$$\begin{aligned}\Delta y_{i,t} = \delta_c + \varphi_i(y_{i,t-1} - \theta_i^c c_{i,t} - \theta_i^k k_{i,t} - \theta_i^p p_{i,t}) + \\ + \sum_{j=1}^{L-1} \lambda_{i,j} \Delta y_{i,t-j} + \sum_{j=0}^{Q_c-1} \mu_{i,j}^c \Delta c_{i,t-j} + \sum_{j=0}^{Q_k-1} \mu_{i,j}^k \Delta k_{i,t-j} \\ + \sum_{j=0}^{Q_p-1} \mu_{i,j}^p \Delta p_{i,t-j} + u_{i,t}\end{aligned} \tag{4}$$

where $\varphi_i = -\left(1 - \sum_{j=1}^{L} \gamma_{i,j}\right)$ is the error-correcting speed of adjustment (*ecm*) term and parameters of particular interest are $-\varphi_i\theta_i^j$, with $j = c,k,p$, i.e., long-run liberalization elasticities (Chudik et al. 2016). The *ecm* term is expected to be negative if the variables exhibit a return to the long-run equilibrium. The vector $\left[1\, \theta^c\, \theta^k\, \theta^p\right]$ defines the long-run relationship between output growth and changes in current and capital account liberalization, while $\lambda_{i,j}$ and $\mu_{i,j}$ capture the short-run dynamics between the covariates.

A simple approach to deal with CSD is to assume the common factor impact to be identical for all the units and to introduce year dummies to account for time-variant correlation across countries. Unfortunately, this method does not take into account country heterogeneity. To cope with these issues, Bai and Ng (2002) propose to estimate factors by principal component analysis, whereas Pesaran (2006) prefers to approximate them with the cross-sectional means of the variables in the long-run relationship. Furthermore, the dynamic CCEMG estimator requires to augment (4) with a sufficient number of lags of cross-sectional averages, where the number is the integer part of $T^{1/3}$. Yet, this requirement can largely reduce the degrees of freedom when the panel is not too long enough.[7] Hence, we rely on inflation as the only one macroeconomic control.

Finally, we employ the empirical methodology provided by Engle and Granger (1987) to check the directions of causality between growth and liberalization. According to this approach, current or capital account liberalization exerts a causal influence on growth if the former is a significant predictor of the current value of the latter, even when past values of output growth have been included in the model. In a similar fashion, liberalization is Granger-non-causal for growth when we are not able to predict the latter better with the past history of the former in the information set than when it is omitted. We rely on the two-step procedure based on the Granger representation theorem as suggested by Canning and Pedroni (2008). Granger's representation theorem states these series can be represented via a dynamic ECM:

$$\Delta y_{i,t} = \delta_{1i} + \varphi_{1i}\widehat{\varepsilon}_{i,t-1} + \sum_{j=1}^{K} \phi_{1i,j}^y \Delta y_{i,t-j} + \sum_{j=1}^{K} \phi_{1i,j}^c \Delta c_{i,t-j} + \sum_{j=1}^{K} \mu_{1i,j}^k \Delta k_{i,t-j}$$

[7] However, a smaller number of lags are sometimes sufficient to remove CSD in the residuals and lessen the issue of the degrees of freedom. Yet, results without cross-means are unreliable and differ from the ones shown in Table 5. They are available from the authors upon request.

$$+\sum_{j=1}^{K}\mu_{1i,j}^{p}\Delta p_{i,t-j}+u_{1i,t} \tag{5}$$

$$\Delta c_{i,t}=\delta_{2i}+\varphi_{2i}\widehat{\varepsilon}_{i,t-1}+\sum_{j=1}^{K}\phi_{2i,j}^{y}\Delta y_{i,t-j}+\sum_{j=1}^{K}\phi_{2i,j}^{c}\Delta c_{i,t-j}+\sum_{j=1}^{K}\mu_{2i,j}^{k}\Delta k_{i,t-j}$$

$$+\sum_{j=1}^{K}\mu_{2i,j}^{p}\Delta p_{i,t-j}+u_{2i,t} \tag{6}$$

$$\Delta k_{i,t}=\delta_{3i}+\varphi_{3i}\widehat{\varepsilon}_{i,t-1}+\sum_{j=1}^{K}\phi_{3i,j}^{y}\Delta y_{i,t-j}+\sum_{j=1}^{K}\phi_{3i,j}^{c}\Delta c_{i,t-j}+\sum_{j=1}^{K}\mu_{3i,j}^{k}\Delta k_{i,t-j}$$

$$+\sum_{j=1}^{K}\mu_{3i,j}^{p}\Delta p_{i,t-j}+u_{3i,t} \tag{7}$$

$$\Delta p_{i,t}=\delta_{4i}+\varphi_{4i}\widehat{\varepsilon}_{i,t-1}+\sum_{j=1}^{K}\phi_{4i,j}^{y}\Delta y_{i,t-j}+\sum_{j=1}^{K}\phi_{4i,j}^{c}\Delta c_{i,t-j}+\sum_{j=1}^{K}\mu_{4i,j}^{k}\Delta k_{i,t-j}$$

$$+\sum_{j=1}^{K}\mu_{4i,j}^{p}\Delta p_{i,t-j}+u_{4i,t} \tag{8}$$

where the disequilibrium term $\widehat{\varepsilon}$ is derived from the cointegrating relationship estimated via fully modified ordinary least squares (FMOLS). All the variables in this ECM system are stationary and replacing the error correction term with its estimate does not affect the standard properties due to the superconsistency of the estimator of the long-run relationship (Canning and Pedroni 2008). Furthermore, we follow the suggestion by Eberhardt and Teal (2013) and control for CSD augmenting the model by cross-sectional averages of all the variables needed to tackle the presence of common factors. Then, we carry out the usual tests on the parameter estimates. First, the Granger representation theorem implies that for a long-run relationship to exist at least one of the coefficient φ_i with $i = 1, 2, 3, 4$ must be different from zero. If a standard t-ratio rejects the null in a country, the set of covariates in that ECM specification has a causal impact on the dependent variable in that nation. Nonetheless, the sample size is relatively small, with less than 50 observations per country, and the reliability of these individual tests is limited. Therefore, we revert to the two panel tests recommended by Canning and Pedroni (2008). The first one is the Group Mean (GM) test that averages countries' t-statistics. This has an asymptotic normal distribution under the null of no long-run causal effect. The other one is a Fisher-Type (FT) test as it is constructed from the P-values of the t-ratios in each ECM regression. This lambda-Pearson statistic is equal to $-2\sum_{i=1}^{N} lnp_{\widehat{\varphi_i}}$ where $p_{\widehat{\varphi_i}}$ is the probability associated to the t-test of the error correction term in each nation. This statistic is distributed as a χ^2 with $2\,N$ degrees of freedom under the null of no long-run causation for the panel. The null and alternative hypotheses for both the GM and FT tests are the same under the assumption that coefficients are homogenous and equal to zero across countries, but they are dissimilar from nil for some non-negligible shares of the countries under the alternative.

4 Findings

It is likely that cross-sectional units in our panel are interdependent because common shocks may impact trade liberalization, capital account liberalization, prices and income amongst countries. In order to distinguish between weak and strong dependence, we rely on the cross-sectional (CD) test we present in Table 3. The CD statistic always rejects the null hypothesis of weakly CSD at the 1% level of significance for all the variables both in levels and in first differences. This is also true by looking at the exponent of cross-sectional dependence (Bailey et al. 2016). Estimates are all close to unity, and the lower bound of the 90% interval band is above 0.5, exception made for the level of *gdp* and *tra* in the upper middle- and high-income countries.

Due to the presence of CSD, we perform CIPS tests for unit roots with individual dynamics according to the Wald and Portmanteau approaches (Burdisso and Sangiácomo 2016). Results reported in Table 4 are clear cut. No variable is I(2) at 1% significance level whatever the individual lag selection mechanism. Moreover, the liberalization indices and the price level are stationary as we can also reject the null hypothesis of a unit root in all country series in the global panel as well as in both sub-panels. On the contrary, real GDP is not stationary since we cannot reject the null of a unit root, with the partial exception of the global panel when we add a trend at the 10% significance level. Hence, we can safely assume that real GDP is I(1) and address the model in growth rates so that coefficients can be interpreted as elasticities.

According to the results we got from the above tests, our datasets, in its entirety and its subsamples, are characterized by some of the typical features of macro panels, i.e., CSD and heterogeneity.[8] To deal with these issues we employ an ARDL approach using the CS-ARDL model (Chudik et al. 2016; Ditzen 2018). Since we are working with moderately persistent growth rates, we choose a relatively small lag order. This holds for the whole sample, for which we specify an ARDL (2,0,2,1), for the sample of low- and lower middle-income countries, for which we chose an ARDL (1,0,2,3), as well as for the sample of upper middle and high income countries, for which we opt for an ARDL (3,3,2,3). Our focus is on the long-run elasticities which are reported in Table 5. The results indicate that, in the global sample, there is a positive association between changes in the de jure KOF measure of trade openness and economic growth as well as between changes in the de jure KOF measure of financial openness and economic growth.[9] The coefficient associated with trade liberalization is slightly larger. Yet, the difference in the magnitude is quite small and indicates that a 10% increase in

[8] We believe that pooling is a too strong assumption (and possibly a wrong one), since the impact of trade and financial liberalization (as well as that of inflation) on economic growth is likely to be different not only across but also within country groups. For the sake of completeness, results obtained imposing homogeneous long-run effects in the error correction model, which show insignificant coefficients, are available from the authors upon request.

[9] We also tested the association between the Chinn-Ito measure of financial openness and economic growth. The estimated results, available upon request, are similar to those obtained using the de jure KOF measure of financial openness. However, a comparison between the two results seems inappropriate because the

Table 3 Tests of cross-sectional dependence

	CD test	p-value	$\widehat{\alpha}$ and 90% confidence bands		
			$\widehat{\alpha}_{0.05}$	$\widehat{\alpha}$	$\widehat{\alpha}_{0.95}$
All countries					
gdp	500.7	0.00	0.578	1.002	1.426
d.gdp	41.8	0.00	0.604	0.662	0.720
tra	256.6	0.00	0.704	0.950	1.195
d.tra	34.9	0.00	0.671	0.729	0.787
fin	231.9	0.00	0.924	0.977	1.029
d.fin	55.3	0.00	0.781	0.835	0.890
pri	444.0	0.00	0.935	0.997	1.060
d.pri	145.3	0.00	0.909	0.935	0.961
Low- and lower middle-income countries					
gdp	216.7	0.00	0.810	0.998	1.185
d.gdp	17.5	0.00	0.651	0.705	0.759
tra	92.3	0.00	0.721	0.925	1.130
d.tra	24.5	0.00	0.604	0.667	0.730
fin	60.2	0.00	0.825	0.882	0.939
d.fin	20.7	0.00	0.739	0.819	0.899
pri	189.4	0.00	0.924	0.998	1.072
d.pri	50.0	0.00	0.881	0.914	0.948
Upper middle- and high-income countries					
gdp	285.9	0.00	0.253	1.003	1.752
d.gdp	45.3	0.00	0.640	0.703	0.767
tra	176.5	0.00	− 3.954	0.945	5.844
d.tra	25.9	0.00	0.506	0.573	0.641
fin	182.6	0.00	0.922	0.988	1.054
d.fin	43.4	0.00	0.767	0.828	0.889
pri	260.2	0.00	0.930	0.999	1.068
d.pri	101.3	0.00	0.926	0.953	0.980

The full sample includes 118 countries. The number of low- and lower middle-income countries is 54. The number of upper middle- and high-income countries is 64. The time span under scrutiny for the countries in the whole sample and in the two subsamples goes from 1970 to 2017, i.e., 47 years. Under the null of no CSD, the CD test is standard normal distributed. $\widehat{\boldsymbol{\alpha}}$ is the exponent of CSD; values of $\widehat{\boldsymbol{\alpha}}$ in the range of [0.5,1] depict different degrees of strong CSD

liberalization is associated with about a 2% increase in GDP growth. An extra 2% of growth each year may not sound like a lot but this gain cumulates to more than 20% higher income after a decade. There are very few credible prescriptions that might achieve such a goal (Estevadeordal and Taylor 2013).

Footnote 9 continued
Chinn-Ito measure of financial openness is continuously available for the time span of our interest only for 46 countries, whereas the de jure KOF measure of financial openness is available for 118 nations.

Table 4 Panel unit root CIPS tests with individual lags

Variable	Without trend		With trend	
	Levels	Differences	Levels	Differences
All countries				
gdp				
Wald	− 1.80	− 4.95***	− 2.51*	− 5.12***
Portmanteau	− 1.82	− 4.97***	− 2.57*	− 5.15***
tra				
Wald	− 2.22***	− 5.48***	− 2.99***	− 5.68***
Portmanteau	− 2.28***	− 5.90***	− 2.91***	− 6.04***
fin				
Wald	− 2.67***	− 5.91***	− 2.78***	− 6.11***
Portmanteau	− 2.55***	− 6.06***	− 2.71***	− 6.23***
pri				
Wald	− 2.37***	− 5.83***	− 2.65***	− 5.93***
Portmanteau	− 2.37***	− 5.76***	− 2.67***	− 5.87***
Low- and lower middle-income countries				
gdp				
Wald	− 2.00	− 5.27***	− 2.68**	− 5.49***
Portmanteau	− 1.97	− 5.32***	− 2.58*	− 5.47***
tra				
Wald	− 2.52***	− 5.60***	− 3.08***	− 5.63***
Portmanteau	− 2.45***	− 6.02***	− 3.03***	− 6.26***
fin				
Wald	− 2.59***	− 6.03***	− 2.76***	− 6.15***
Portmanteau	− 2.59***	− 6.06***	− 2.73***	− 6.25***
pri				
Wald	− 2.53***	− 5.78***	− 2.98***	− 6.02***
Portmanteau	− 2.53***	− 5.82***	− 3.02***	− 5.93***
Upper middle- and high-income countries				
gdp				
Wald	− 1.64	− 4.71***	− 2.18	− 4.85***
Portmanteau	− 1.72	− 4.63***	− 2.30	− 4.75***
tra				
Wald	− 2.13**	− 5.56***	− 2.98***	− 5.82***
Portmanteau	− 2.15**	− 5.80***	− 2.78***	− 5.96***
fin				
Wald	− 2.71***	− 5.87***	− 2.89***	− 6.19***

Table 4 (continued)

Variable	Without trend		With trend	
	Levels	Differences	Levels	Differences
Portmanteau	− 2.61***	− 6.06***	− 2.78***	− 6.23***
pri				
Wald	− 2.39***	− 5.75***	− 2.56*	− 5.73***
Portmanteau	− 2.44***	− 5.72***	− 2.64**	− 5.79***

***, **, * indicate statistical significance at the 1%, 5%, and 10%. Critical values at 10%, 5%, 1%:
All countries: − 2.01, − 2.06, − 2.14 (without trend); − 2.50, − 2.54. − 2.62 (with trend)
Low- and lower middle-income countries: − 2.03, − 2.10, − 2.20 (without trend); − 2.53, − 2.58. − 2.69 (with trend)
Upper middle- and high-income countries: − 2.03, − 2.10, − 2.20 (without trend); − 2.53, − 2.58. − 2.68 (with trend)

It is interesting to notice that the trade elasticity is even larger for the subsample of low- and lower-income countries pinpointing an even stronger impact on cumulative output. Hence, this result may be indicative of the greater importance of easing trade regulations, reducing trade taxes and tariffs, as well as establishing trade agreements for the growth of poorer nations which mostly need to improve standard of living. Conversely, we notice that changes in capital account liberalization do not have a statistically significant impact on GPD changes. This outcome may seem at odds with the original receipt for growth devised by the Washington Consensus. Yet, it is consistent with a more nuanced approach put in place by the IMF itself during the last decade, when it recognized that capital account liberalization needs to be well planned and is less risky for countries that have achieved a certain degree of financial and institutional development. Therefore, full liberalization is likely not to be the most suitable option for all countries at all the times (IMF 2012). It is also in line with recent contributions which underline that developing countries may benefit from stronger growth if a certain degree of capital controls is put in place (Epstein et al. 2008; Schneider 2008; Rodrik 2008; Stiglitz 2016; Choi and Pyun 2019).

The results for the subsample of upper middle- and high-income countries show that both elasticities are positive and large. Moreover, the elasticity of capital account liberalization is significant at the 5% level and greater in magnitude than that of the full sample. This result may be indicative that richer countries can rely on an allegedly more robust financial systems and have been able to reap the benefits of capital account liberalization to foster economic growth, without the fears of sudden stops or capital reversals.

The magnitude of the speed of adjustment, i.e., the *ecm* coefficient, indicates that there is a fast return to equilibrium. In none of the models the change in the level of prices has a statistically significant impact on GDP changes. This superneutrality property is somehow consistent with the mixed evidence suggested by the literature as the effects of inflation on growth may largely depend on the countries investigated or

Table 5 Mean group long-run estimates

	Full sample	Low- and lower middle-income countries	Upper middle- and high-income countries
d.tra	0.207**	0.326**	0.561*
(S.E.)	(0.106)	(0.154)	(0.309)
d.fin	0.183**	− 0.224	0.202**
(S.E.)	(0.082)	(0.173)	(0.093)
d.pri	0.357	− 0.165	0.005
(S.E.)	(0.387)	(0.199)	(0.122)
ecm	− 0.915***	− 1.030***	− 1.053***
(S.E.)	(0.044)	(0.064)	(0.067)
constant	− 0.035*	− 0.046	− 0.030
(S.E.)	(0.020)	(0.035)	(0.026)
Observations	5192	2322	2752
RMSE	0.072	0.081	0.056
R^2	0.262	0.211	0.156
DFE	15	12	11
CD test	0.131	− 0.495	− 0.621
CD prob	0.896	0.621	0.535
$\hat{\alpha}$	0.565	0.577	0.561
(Bootstrapped S.E.)	(0.009)	(0.013)	(0.011)
(95% bootstrapped C.I.)	(0.548; 0.582)	(0.552; 0.601)	(0.540; 0.583)
BB—HR	0.691	0.288	0.807
BB—LM(1)	0.270	0.110	0.792
BB—LM(2)	0.778	0.193	0.714
BB—LM(3)	0.553	0.888	0.547
Stationarity	I(0)	I(0)	I(0)

***, **, * indicate statistical significance at the 1%, 5%, and 10%, respectively.

The full sample includes 118 countries. The number of low- and lower middle-income countries is 54. The number of upper middle- and high-income countries is 64. The time span under scrutiny for the countries in the whole sample and in the two subsamples goes from 1970 to 2017, i.e., 47 years. DFE: Degrees of freedom per group with cross-sectional averages. Residual Diagnostics: CD test, H_0: weak CSD. $\hat{\alpha}$ is the exponent of CSD with bootstrapped standard errors and confidence intervals (Bailey et al. 2019). BB-HR is the heteroskedasticity-robust test for first-order serial correlation due to Born and Breitung (2016). BB—LM is the Born and Breitung (2016) Lagrange multiplier test for serial correlation up to the third order. Values are probabilities associated with the null (for all these tests) of no serial correlation. Maddala–Wu and CIPS test results: I(0), stationary; I(1), non-stationary some inflation thresholds (Ericsson et al. 2001; Kremer et al. 2013; Ibarra and Trupkin 2016; Yilmazkuday 2022).[10]

The lower part of Table 5 reports the diagnostics. The CIPS and Maddala–Wu tests indicate that the residuals are stationary. The CD test strongly supports the null hypothesis of weakly cross-sectional dependent residuals. The absence of strong CSD is confirmed by the values taken by $\widehat{\boldsymbol{\alpha}}$, the exponent of CSD, in all three samples.[11] There is quite strong evidence that in the full sample errors are not correlated. Indeed, we cannot reject the null of no first order serial correlation via the heteroskedasticity-robust test statistic introduced by Born and Breitung (2016) and we also exclude the presence of serial correlation up to the third order via the Born and Breitung (2016) Lagrange multiplier tests.

For a robustness check, we also address the cross section augmented distributed lag (CS-DL) model which allows to directly estimate long-run elasticities as devised by Chudik et al. (2016). According to these authors, the CS-DL is complementary to the CS-ARDL due to its two drawbacks. First, it does not allow for feedback effects from the dependent variable onto the regressors. This can be a serious issue if output growth leads to economic reforms with trade or capital liberalization. Second, its small sample performance deteriorates when the roots of the AR polynomial in the ARDL representation are close to the unit circle. On the contrary, the CS-DL model has better small sample performance for moderate values of T and it is robust to residual serial correlation and breaks in the error processes. Table 6 reports results which corroborate previous findings.

The final step in our analysis concerns causality since results in Table 5 and 6 only show that there is a long-run relationship between output growth and changes in liberalization indices and inflation. This is an important issue since the CS-DL estimation of the long-run effects is consistent only in the case when the feedback effects (or reverse causality) are not present. We present causality analysis in Table 7. First, we report the results for a test if both liberalization and price changes have a causal impact on output growth, with the null hypothesis of 'no causal impact' when we opt for an ECM with only one lag. In analogy, in the subsequent rows, we test whether other three variables have a causal impact on the remaining one. Looking at columns, starting from the left we first report the Group Mean t-test (GM), i.e., the averaged t-ratio and its probability value, then the Fisher-Type statistic (FT), that is derived from the P-values of the t-ratios in each ECM regression and its probability value, and finally the mean group estimate of the coefficient of disequilibrium term $\widehat{\varepsilon}$ in (5)-(8) with the associated probability. In all the tests the null is that the corresponding value is equal to zero. Results in the first row clearly show that, in the global panel,

[10] A standard panel estimation with country and fixed effects produces results different from ours because fixed effects assume the presence of an (unlikely) strict exogeneity between the covariates and the time varying component of the error term in non-dynamic specifications. Moreover, in the presence of a lagged dependent variable fixed effect estimators are generally inconsistent even in the presence of strict exogeneity.

[11] Bailey et al. (2019) argue this is a suitable test also in the absence of a common factor representation structure and it is able to capture moderate to strong CSD. This is an improved version of the CSD estimator developed by Bailey et al. (2016) who provide theoretical justification only in the case of demeaned observations and do not consider residuals from panel regression. The estimator proposed by Bailey et al. (2019) performs well in small samples and it is scaled to lie in the interval (1/2,1]. Values of $\widehat{\boldsymbol{\alpha}}$ close to ½, as in our case, indicate that errors are cross-sectionally weakly correlated.

Table 6 Mean group long-run estimates for CS-DL models

	Full sample	Low- and lower middle-income countries	Upper middle- and high-income countries
d.tra	0.198*	0.342**	0.300**
(S.E.)	(0.118)	(0.141)	(0.132)
d.fin	0.163*	0.168	0.240*
(S.E.)	(0.095)	(0.262)	(0.142)
d.pri	− 0.039	− 0.096	0.016
(S.E.)	(0.047)	(0.075)	(0.095)
constant	− 0.008	− 0.007	− 0.004
(S.E.)	(0.008)	(0.018)	(0.012)
Observations	4956	2160	2688
RMSE	0.068	0.036	0.054
R^2	0.364	0.750	0.241
DFE	16	10	15
CD test	0.484	− 1.228	0.758
CD prob	0.629	0.219	0.448
$\hat{\boldsymbol{\alpha}}$	0.545	0.560	0.562
(Bootstrapped S.E.)	(0.007)	(0.013)	(0.011)
(95% bootstrapped C.I.)	(0.531; 0.558)	(0.535; 0.585)	(0.540; 0.584)
BB—HR	0.303	1.00	0.046
BB—LM(1	0.997	0.582	0.117
BB—LM(2)	0.095	0.218	0.334
BB—LM(3)	0.084	0.645	0.003
Stationarity	I(0)	I(0)	I(0)

***, **, * indicate statistical significance at the 1%, 5%, and 10%, respectively.

The full sample includes 118 countries. The number of low- and lower middle-income countries is 54. The number of upper middle- and high-income countries is 64. The time span under scrutiny for the countries in the whole sample and in the two subsamples goes from 1970 to 2017, i.e., 47 years. DFE: Degrees of freedom per group with cross-sectional averages. Residual Diagnostics: CD test, H_0: weak CSD. $\hat{\boldsymbol{\alpha}}$ is the exponent of CSD with bootstrapped standard errors and confidence intervals (Bailey et al. 2019). BB-HR is the heteroskedasticity-robust test for first order serial correlation due to Born and Breitung (2016). BB-LM is the Born and Breitung (2016) Lagrange multiplier test for serial correlation up to the third order. Values are probabilities associated with the null (for all these tests) of no serial correlation. Maddala–Wu and CIPS test results: I(0), stationary; I(1), non-stationary

we can safely reject the null and claim that inflation and, notably, a change in trade and capital account liberalization together cause a change in output. This change is positive according to long-run estimates given in Table 5 and 6 while is not significant for inflation. On the contrary, the three subsequent rows indicate that there is not such a causal effect on changes in either trade or capital account liberalization and in inflation. The same results hold when we consider an ECM with more lags. Moreover, this finding is quite robust as it applies in less and more developed countries too.

Table 7 Long-run causality tests

	GM	(*p*)	*FT*	(*p*)	*Mean* $\widehat{\varphi}_i$	(*p*)
All countries						
ECM with 1 lag						
$gdp \leftarrow tra, fin, pri$	– 2.932	0.00	1471.7	0.00	– 0.822	0.00
$tra \leftarrow gdp, fin, pri$	0.055	0.96	193.9	0.98	0.019	0.50
$fin \leftarrow gdp, tra, pri$	– 0.024	0.98	164.3	1.00	– 0.029	0.98
$pri \leftarrow gdp, tra, fin$	0.248	0.80	245.0	0.33	0.017	0.77
ECM with 2 lags						
$gdp \leftarrow tra, fin, pri$	– 2.218	0.03	980.7	0.00	– 0.811	0.00
$tra \leftarrow gdp, fin, pri$	0.048	0.96	177.1	1.00	0.035	0.43
$fin \leftarrow gdp, tra, pri$	– 0.128	0.90	193.0	0.99	– 0.078	0.24
$pri \leftarrow gdp, tra, fin$	0.005	1.00	208.2	0.90	– 0.052	0.48
ECM with 3 lags						
$gdp \leftarrow tra, fin, pri$	– 1.796	0.07	736.8	0.00	– 0.828	0.00
$tra \leftarrow gdp, fin, pri$	0.011	0.92	229.1	0.61	0.052	0.43
$fin \leftarrow gdp, tra, pri$	– 0.021	0.98	193.7	0.98	– 0.001	0.99
$pri \leftarrow gdp, tra, fin$	– 0.068	0.95	237.9	0.45	– 0.111	0.23
Low- and lower middle-income countries						
ECM with 1 lag						
$gdp \leftarrow tra, fin, pri$	– 3.244	0.00	801.1	0.00	– 0.951	0.00
$tra \leftarrow gdp, fin, pri$	– 0.077	0.94	104.7	0.57	– 0.023	0.77
$fin \leftarrow gdp, tra, pri$	– 0.079	0.94	75.3	0.99	– 0.058	0.52
$pri \leftarrow gdp, tra, fin$	0.354	0.72	109.7	0.44	0.004	0.96
ECM with 2 lags						
$gdp \leftarrow tra, fin, pri$	– 2.275	0.02	464.7	0.00	– 0.887	0.00
$tra \leftarrow gdp, fin, pri$	– 0.190	0.85	89.4	0.90	– 0.941	0.19
$fin \leftarrow gdp, tra, pri$	– 0.044	0.96	92.9	0.85	– 0.032	0.78
$pri \leftarrow gdp, tra, fin$	0.031	0.98	75.6	0.99	– 0.102	0.30
ECM with 3 lags						
$gdp \leftarrow tra, fin, pri$	– 1.678	0.09	297.6	0.00	– 0.869	0.00
$tra \leftarrow gdp, fin, pri$	– 0.034	0.97	99.0	0.72	– 0.040	0.69
$fin \leftarrow gdp, tra, pri$	0.055	0.96	84.4	0.95	0.048	0.65
$pri \leftarrow gdp, tra, fin$	0.005	1.00	81.7	0.97	– 0.152	0.23
Upper middle- and high-income countries						
ECM with 1 lag						
$gdp \leftarrow tra, fin, pri$	– 2.816	0.00	726.7	0.00	– 0.736	0.00
$tra \leftarrow gdp, fin, pri$	– 0.076	0.94	133.6	0.35	0.019	0.75
$fin \leftarrow gdp, tra, pri$	– 0.041	0.97	81.5	1.00	– 0.007	0.91

Table 7 (continued)

	GM	(*p*)	*FT*	(*p*)	*Mean* $\widehat{\varphi}_i$	(*p*)
pri ← *gdp, tra, fin*	0.054	0.96	126.3	0.53	0.028	0.71
ECM with 2 lags						
gdp ← *tra, fin, pri*	− 2.249	0.02	538.8	0.00	− 0.732	0.00
tra ← *gdp, fin, pri*	0.040	0.97	125.1	0.55	0.052	0.38
fin ← *gdp, tra, pri*	− 0.166	0.87	93.6	0.99	− 0.116	0.21
pri ← *gdp, tra, fin*	0.108	0.91	134.6	0.33	− 0.032	0.71
ECM with 3 lags						
gdp ← *tra, fin, pri*	− 2.046	0.04	467.6	0.00	− 0.822	0.00
tra ← *gdp, fin, pri*	0.201	0.84	143.3	0.18	0.145	0.12
fin ← *gdp, tra, pri*	− 0.119	0.91	113.4	0.82	− 0.122	0.30
pri ← *gdp, tra, fin*	− 0.189	0.85	135.4	0.31	− 0.059	0.54

GM gives the group-mean average of country-specific t-ratios for the coefficient on the disequilibrium term ($\widehat{\varphi}_i$) which is distributed $N(0,1)$. Fisher (*FT*) gives $-2\sum_i \log\pi_i$ where π_i is the probability value of the country-specific *t*-ratio on the disequilibrium term. The Fisher statistic is distributed $\chi^2(2\,N)$

Hence, we can conclude that liberalization policies appear to cause real output growth in all the panels under scrutiny and there are not feedback effects.

5 Conclusions

Academics have extensively analyzed the association between current and capital account liberalization and economic growth for decades, often recording mixed evidence. A very narrow and recent scholarship reassesses the positive effects on economic growth stemming from liberalization policies, as prescribed by the Washington Consensus, taking stock that successful development strategies and supporting policies are always time- and context-specific (Spence 2021). Easterly (2019) observes that bad policy outcomes were common during the 1980s and 1990s, but disappeared afterward, and asks himself whether the recommendations of the Washington Consensus may have had a delayed effect. It is the case of Grier and Grier (2020) who find that policy reforms in line with the Washington Consensus measured as jumps in the Fraser Institute's Economic Freedom of the World led to an increase in per capita income.

Mixed results may not merely depend from the variety among nations and time spans under scrutiny, but they may also stem from the variety of proxies used to measure the opening up to trade and the extent of current or capital account liberalization (Gräbner et al. 2020). A further possible cause of conceptual bias lies on looking at policy outcomes—rather than on policy levers—as means to assess the effectiveness of normative efforts toward trade and financial openness in achieving economic growth (Easterly 2019).

We take stock of these weaknesses in the extant literature and resort to KOF database, the most applied index of economic openness (Potrafke 2015), to analyze the relationship between de jure trade openness and economic growth and between de jure financial openness and economic growth for a panel of 118 countries in the period 1970–2017. We make use of an ARDL approach and on its error correction formulation to show that, for the entire sample and for upper middle and high income countries, there is a positive long-run impact of both changes in trade and capital openness, on the one hand, and economic growth, on the other hand. This result fits in a well-established literature since it underlines that the *laissez-faire* approach has led globally to economic growth in the last 50 years. The picture is, yet, slightly different when we concentrate on low- and lower middle-income countries, since only changes in the de jure trade liberalization have a positive and statistical impact on economic development in the long run. Conversely, changes in the de jure financial openness do not seem to affect output growth. We claim that this is possibly because during the time span of interest, developing countries faced current account crises when they stuck to the early bold prescriptions of the Washington Consensus, but they later embraced a more nuanced view of economic liberalism putting in place a number of capital controls which shielded them from sudden stops and capital reversals. We complement these results with a Granger-causality approach, which unveils that for the whole sample and for the two subsamples changes in both de jure trade and de jure financial openness jointly cause economic growth, while controlling for inflation. There is never evidence of a feedback effect, providing support that the two forms of liberalization do positively impact economic growth. Summing up, our results are robust to potential bias stemming from reverse causality, CSD and serial correlation.

Despite simple in its formulation and on how it unfolds, this paper brings evidence that is relevant for both scholars and policy makers.

From an academic perspective we believe that the use of KOF sub-indexes to measure de jure trade and de jure financial openness is the most appropriate option now available given its comprehensiveness, clarity and time coverage (Gräbner et al. 2020). The use of the ECM unveils the real nature and magnitude of the relationship between the above-mentioned measures of openness and economic growth, possibly solving a long-standing controversy the discipline often tackled, maybe inappropriately, via simple growth models (Spence 2021). The validity of our outcomes is reinforced by the strong evidence that the direction of causation goes from our covariates to output growth.

From a policy perspective, the results support the removal of trade taxes, tariffs and regulations and encourage the establishment of trade agreements, as well as an ease of capital account restrictions. Yet, the latter requires caution in the case of countries characterized by lower income levels which are often developing or emerging nations. In these circumstances, the policy maker is advised that the country will likely benefit from some forms of capital controls. This may be attained from the adoption of a gradualist approach in undertaking reforms toward freer international capital flows, since these proved more effective in promoting growth when "markets and institution are not at their infancy" (Prati et al. 2013, p. 967), and it can also be achieved via a precautionary and mercantilism approach with reserve accumulations (Choi and Taylor 2017).

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Data availability The data that support the findings of this study are available from the corresponding author upon request.

Declarations

Conflict of interest All authors declare that they have no conflict of interest.

Appendix

See Table A.

Table A Country list

N	Countries	N	Countries	N	Countries
1	Albania	41	Germany	81	Niger
2	Algeria	42	Ghana	82	Nigeria
3	Argentina	43	Greece	83	Norway
4	Australia	44	Guatemala	84	Oman
5	Austria	45	Guinea	85	Pakistan
6	Barbados	46	Haiti	86	Panama
7	Belgium	47	Honduras	87	Paraguay
8	Benin	48	Hungary	88	Peru
9	Bhutan	49	Iceland	89	Philippines
10	Bolivia	50	India	90	Poland
11	Botswana	51	Indonesia	91	Portugal
12	Brazil	52	Iraq	92	Republic of Korea
13	Bulgaria	53	Ireland	93	Rwanda
14	Burkina Faso	54	Israel	94	Saudi Arabia
15	Burundi	55	Italy	95	Senegal
16	Cambodia	56	Jamaica	96	Seychelles
17	Cameroon	57	Japan	97	Sierra Leone
18	Canada	58	Jordan	98	Singapore
19	Central African Republic	59	Kenya	99	South Africa

Table A (continued)

N	Countries	N	Countries	N	Countries
20	Chad	60	Kuwait	100	Spain
21	Chile	61	Lao People's Dem. Republic	101	Sri Lanka
22	China	62	Lebanon	102	Sudan
23	China, Hong Kong SAR	63	Lesotho	103	Sweden
24	Colombia	64	Liberia	104	Switzerland
25	Congo	65	Luxembourg	105	Syrian Arab Republic
26	Costa Rica	66	Madagascar	106	Thailand
27	Cyprus	67	Malawi	107	Togo
28	Côte d'Ivoire	68	Malaysia	108	Trinidad and Tobago
29	Denmark	69	Mali	109	Tunisia
30	Dominican Republic	70	Malta	110	Turkey
31	Ecuador	71	Mauritania	111	U.R. of Tanzania: Mainland
32	Egypt	72	Mauritius	112	Uganda
33	El Salvador	73	Mexico	113	United Kingdom
34	Eswatini	74	Mongolia	114	United States
35	Ethiopia	75	Morocco	115	Uruguay
36	Fiji	76	Myanmar	116	Venezuela
37	Finland	77	Nepal	117	Viet Nam
38	France	78	Netherlands	118	Zambia
39	Gabon	79	New Zealand		
40	Gambia	80	Nicaragua		

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