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Current account determinants in a globalized world

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Abstract
This paper investigates the determinants of external imbalances for a group of 26 developed and emerging countries over the period 1972–2021. In addition to traditional factors, the model incorporates the impact of external imbalances in third countries. The empirical evidence highlights the importance of accounting for parameter instabilities in modeling external imbalances, with countries exhibiting heterogeneous behavior in terms of the estimated break dates. The results underscore the critical role of external drivers, such as oil shocks, and the growing influence of third-country imbalances in an increasingly globalized world. Additionally, demographic trends emerge as a significant long-run internal driver. Finally, the paper calculates regime-specific short-run multipliers.

Keywords Current account · Global factors · Structural break · Dynamic multipliers

JEL Classification F02 · F32 · C32

1 Motivation: external imbalances and globalization

Current account determinants have always been central to the international macroeconomics research agenda. This attention has been mainly due to the volume and persistence of the current account deficits in developed and developing countries.

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While historically, many external imbalances might have been linked to domestic factors, the increasing interconnectedness of economies in a globalized world has introduced additional complexities that can lead to external disequilibria. Relying solely on internal macroeconomic explanations to understand external imbalances may be increasingly misleading, as it overlooks the intricate web of global economic interactions and their influence on the international economic landscape (Ca' Zorzi et al. 2012). In this paper, we strive to solve this conundrum by proposing new variables embedding the external drivers of current account imbalances and implementing an empirical strategy based on cointegration techniques that may solve potential problems existing in previous empirical research (endogeneity, cross-sectional dependence, serial correlation or, simply, spurious relationships) within a full-fledged econometric setting.

Policymakers and economists need to consider the broader implications of globalization when analyzing and addressing external imbalances. Different arguments can give support to this conjecture. First, due to the increasing importance of global value chains, disruptions in one country have cascading effects on others (López-Villavicencio and Mignon 2021). Second, globalization has increased financial integration, and sudden shifts in factors influencing capital flows may lead to positions that may not align with a country's internal macroeconomic fundamentals (Lane and Milesi-Ferretti 2002). Third, trade imbalances can also be influenced by exchange rate fluctuations, trade policies and external demand shocks that are not solely determined by internal factors (Chinn 2017). Finally, policies of one country can have significant spillover effects on other countries due to trade, investment and financial linkages (Camarero et al. 2021). Against this backdrop, the importance of relying on an empirical approach to account for cross-sectional dependence across the countries considered in the sample is clear.

The gist of our investigation is to assess the relative importance of the role played in external disequilibria by internal—i.e., fiscal deficit and demography—versus external factors—i.e., real exchange rate and net foreign assets (NFA) position, among others. This is not a trivial question since it implies different corollaries concerning the optimal behavior of diverse groups of countries (debtors and creditors) together with dissimilar stress on the economic policy toolbox to implement the external adjustment.

This research aims to estimate the determinants of the external imbalances in a group of countries, including developed OECD members, together with some emerging countries, in a general specification as the one defined in Chinn and Ito (2007) and Chinn and Prasad (2003). The paper analyzes the period 1972–2021, which embeds the transformation of the world economy toward globalization through distinct stages of real and monetary integration. Together with the proper modelization of potential spillovers, another vital economic issue is distinguishing between short- and long-run effects from structural or cyclical drivers.

The contributions of the present paper are twofold. First, we add to an otherwise standard specification of the current account determinants, variables embedding the external disequilibria in third countries to capture their effects in a world increasingly globalized both on real and financial grounds—see Camarero et al. (2021). Second, from an econometric point of view, we address potential endogeneity, cross-sectional dependence—through a factor augmented model specification *à la* Pesaran—and parameter instability problems for the macroeconomic variables used as explanatory variables. The model specification is flexible enough to allow for the presence of one structural break affecting the parameters of the model. Moreover, we can calculate the short-run dynamic multipliers for each of the countries in the sample, taking into account the existence of structural breaks affecting the constant term and/or the long-run relationship. The short-run multipliers are obtained from an error correction mechanism model specification that is estimated by implementing a two-step estimation procedure *à la* Engle and Granger (1987). In the first stage, the long-run effects are computed from the efficient estimation of the cointegration relationship. In the second step, the short-run dynamic parameters are estimated, imposing the error correction term from the first step. Baek and Lee (2022) analyze the effects of misspecifying the number of lags (orders) of autoregressive and distributed lag (ARDL) models on the impulse response estimation of shocks. They outline that misleading conclusions about the long-run effects of shocks on the dependent variable can be obtained due to flattened impulse responses that cover long horizons. Thus, imposing the cointegrating vector from the first stage forces the multiplier analysis to satisfy the restriction that the long-run multipliers coincide with the (changing) cointegrating vector.

The remainder of the paper is organized as follows. Section 2 contains a revision of the previous empirical literature, proposing a group of the main determinants of the evolution of the current account. In Section 3, we present the theoretical framework that guides our empirical investigation and describe the econometric specification and the database. Section 4 reports and discusses the results. Finally, Section 5 concludes. Supplementary material is provided in the appendices.

2 Potential internal and external drivers shaping current account balances

It is widely recognized the existence of two perspectives, not mutually exclusive, of potential adjustment channels of the current account, namely the domestic and the international one. According to these two perspectives, we can single out the following drivers:

First, the ‘internal absorption’ is based on the interaction between the current account and the fiscal policy (twin deficits) or the difference between private savings and investment (Barro–Ricardian equivalence). According to the ‘twin deficits hypothesis’ (Feldstein 1985), we should observe a positive relationship between the government fiscal balance and the current account. In contrast, in inter-temporal current account models with Ricardian agents, an increase in the deficit will be offset by a rise in private savings so that the current account position remains unaltered (Obstfeld and Rogoff 1995).[1] Within this framework, the impact of the fiscal deficit on the current account is positive and will depend on the share of non-Ricardian households in the economy. Evidence by Abbas et al. (2011) showed that sizeable current account dete-

[1] More recently, some inter-temporal models have assumed that there are two kinds of agents: ‘spenders’, which spend their disposable income at each period, and ‘savers’, which consume according to their permanent income and smooth their resources inter-temporally (Bussière et al. 2010).

rioration was associated with changes in the private sector saving–investment balance rather than the public one.

Second, another driver of the current account evolution is 'income.' Higher average output growth or productivity may affect the current account position. The effect critically depends on whether they signal a temporary or permanent increase in income. If temporary, saving would rise in the short run, leading to a decline in the current account deficit, while in the long run, changes in investment could match that of saving (Kraay and Ventura 2000). If the increase in income is permanent, consumption and investment will rise, and the deficit increases.

Third, another group of drivers is related to competitiveness and, more specifically, to the 'terms of trade' (ToT) channel. The Haberger–Laursen–Metzler effect[2] states that a temporary increase in terms of trade leads to a temporary rise in the level of real income above its permanent level. As a result, consumption is smoothed by increasing short-term savings (if one assumes that the propensity to consume is less than one, consumption increases less than income), leading to a current account surplus.

Fourth, an increasingly influential driver group is related to 'financial factors.' The inefficient financial markets of emerging economies discourage domestic investment and encourage savings. This excess saving must then be invested in countries where the financial markets are supposed to be more efficient, specifically in the USA (saving glut hypothesis), explaining the Lucas Paradox.[3] Moreover, the net effect of financial depth on the current account is conceptually unclear. It could lead to higher financial savings and significantly boost consumption and investment through looser borrowing constraints (Chinn and Prasad 2003).

As pointed out by Gossé and Serranito (2014), in the empirical literature, financial development can be measured either by asset price variables—i.e., real interest rate differentials—or by quantity indexes—among others, the ratio of private credit to GDP or the NFA position to GDP.

While the real interest rate differentials reflect straight differences in asset preferences and/or risk premiums, the NFA position influence on the current account has a more complex explanation. On the one hand, a high NFA position is associated with positive investment income flows, which improve the current account. On the other hand, a highly indebted country may have to improve its current account position to preserve solvency eventually. Hence, the theoretically expected sign is ambiguous. However, most empirical studies find a positive link (deficits rise with more negative NFA). Finally, having a reserve currency (proxied by the share of a country's currency in world international reserves) is associated with higher consumption and a more significant deficit (Phillips et al. 2013).

Fifth, drivers linked to 'demographic factors' influence mainly the saving behavior of an economy. The life cycle hypothesis[4] suggests that savings are accumulated during the working age while younger and older age cohorts generally dissave. Thus, a country with high old and/or young age dependency ratios could generally be expected to save

[2] See Laursen and Metzler (1950).

[3] While economic theory suggests that capital should flow from rich to developing countries, capital flows actually go in the opposite direction (Lucas 1990).

[4] See Ando and Modigliani (1963).

relatively less. The higher the proportion of the 'dependent' population,[5] the lower the level of national savings should be, and the lower the current account balance is.

Sixth, another important factor is the "oil dependency. This driver is closely related to ToT, but it has been treated as a separate element due to its relevance. Thus, we expect either a positive or nonsignificant relationship, depending on whether the oil shock is offset by the non-oil trade balance (Kilian et al. 2009). Higher oil prices improve the current account balance of oil exporters while reducing the balance of oil importers.

Seventh, variables included in the group of 'institutional and regulatory quality' have gained momentum in the current account analysis. Improving the quality of the legal and regulatory system should generally reduce uncertainty and promote economic growth through a boost in investment that would reduce the current account balance. However, both forces would increase investment but have an ambiguous impact on saving (lower uncertainty reduces saving but higher growth increases it).

Eight, some 'dummy variables' are usually added to grasp the effects of different types of shocks. For example, a financial center dummy is often included, as economies that serve as hubs for international financial flows have tended to run substantial current account surpluses and net creditor positions.

3 Model specification and data

The analysis includes twenty-six countries. The majority are developed OECD economies, among them eleven countries of the eurozone (Austria, Belgium, Finland, France, Germany, Greece, Ireland, Italy, the Netherlands, Portugal and Spain); two EU countries non-eurozone members (Denmark and Sweden); ten developed non-European Union countries (Australia, Canada, China, Japan, Korea, New Zealand, Russia, Switzerland, the UK and the USA); and, finally, three emerging countries: Brazil, India and Mexico. The combination of different data sources has allowed us to define an annual database ranging from 1972 to 2021 ($T = 50$).[6] The choice of the countries relies not only on the availability of data necessary for robust dynamic analysis but also on their role in the world economy. We are interested in the behavior of the main economic actors and, in particular, those that exhibit the most prominent imbalances, both developed and emerging.

The extant economic literature has suggested different groups of drivers that can be used to model the evolution of the current account. Therefore, we choose one representative variable from the main groups of drivers present in Sect. 2 as the determinants that will enter the covariates considered in our econometric specification.[7] Finally, we

[5] The dependent population consists of people who are either too young or too old to work.

[6] Data availability shortens the analyzed period for China (1981–2021) and Russia (1994–2021). Further details concerning the definition and sources for the variables used in the paper are given in Appendix A.

[7] More precisely, from the internal absorption group of variables, we include the fiscal deficit; from the income group, we select the external position from the rest of the world; from the competitiveness group of drivers, we choose the real exchange rate; from the financial factors, we select the real interest rate and the NFA position; from the demographic factors, we select the old dependency; and to measure oil dependency, we select the real oil price.

account for possible discontinuities in the series and the cointegrating relationships. Only the variable group related to institutional and regulatory quality is omitted due to data availability constraints.

Finally and as mentioned above, one distinctive feature of our study is that we include the increasingly important effects of the rest of the world on the adjustment of the current account. For this purpose, we compute a specific variable for each country in the sample: a weighted average of the current account position of the remaining $(N-1)$ countries, since the current account balance of one country is not only affected by domestic determinants but also by developments in the rest of the world.[8] Camarero et al. (2021) used a similar approach in the context of a GVAR model. The relevance of this variable and the critical degree of heterogeneity found recommended an individual analysis of the countries, using a country-specific weighted measure of the CA for the rest of the countries (CA^*) to capture the effect of foreign imbalances on the domestic economy.[9]

Based on the review of the theoretical and empirical contributions in the literature, data availability for the sample of countries studied and computation constraints, the variables that we have finally included in our analysis are the CA-to-GDP ratio ($cay_t = CA_t/GDP_t$), the NFA-to-GDP ratio ($nfay_t = NFA_t/GDP_t$), the fiscal deficit-to-GDP ratio ($fdy_t = FD_t/GDP_t$), the real interest rate (rir_t), a measure of population dependence that is defined as the ratio of people above 65 years old to the total population ($olddep_t$) and the CA of the rest of the countries-to-GDP ratio ($cay_t^* = CA_t^*/GDP_t$). These variables define the so-called country-specific variables, which are collected in the vector of stochastic regressors $x_t = (nfay_t, fdy_t, rir_t, olddep_t, cay_t^*)'$. Figure 1 depicts some time series by country. In addition, the analysis also includes the dollar/euro real exchange rate (rer_t) and the real oil price ($rpoil_t$) as global variables—which defines $\omega_t = (rer_t, rpoil_t)'$. The global variables are presented in Fig. 1 as well. Therefore, the analysis that is performed in the paper is based on the set of variables collected in the vector $Y_t = (cay_t, x_t', \omega_t')'$.

The decision to add the foreign influence of the current account ratio in the model specification is justified by the computation of the cross-sectional dependence that drives the international evolution of the current account ratio. This feature is expected if we consider that one important component of the current account is international trade, which involves commercial transactions among economies. The measure of the degree of cross-sectional dependence can be obtained through the computation of the average pairwise correlation coefficients that are associated with the current account ratio. Using this information, it is possible to test the null hypothesis of weak cross-sectional dependence against the alternative hypothesis of strong cross-sectional dependence using Pesaran (2015b) CSD statistic. The CSD statistic is reported in Panel A of Table 1 and is obtained (i) for every country (against the others) and (ii) for all countries together, using both a constant and a linear time trend as deterministic

8 We follow the methodology described in IMF (2012). Country weights are chosen according to country's trade partners. The selection of trade partners brings a trade perspective where relative movements to trade partners may influence current account developments. Due to a lack of statistical information, we exclude China and Russia when computing the weighted average of the current account balance of one country with respect to the others.

9 For China and Russia, CA^* is defined by the cross-sectional average of CA of the remaining countries.

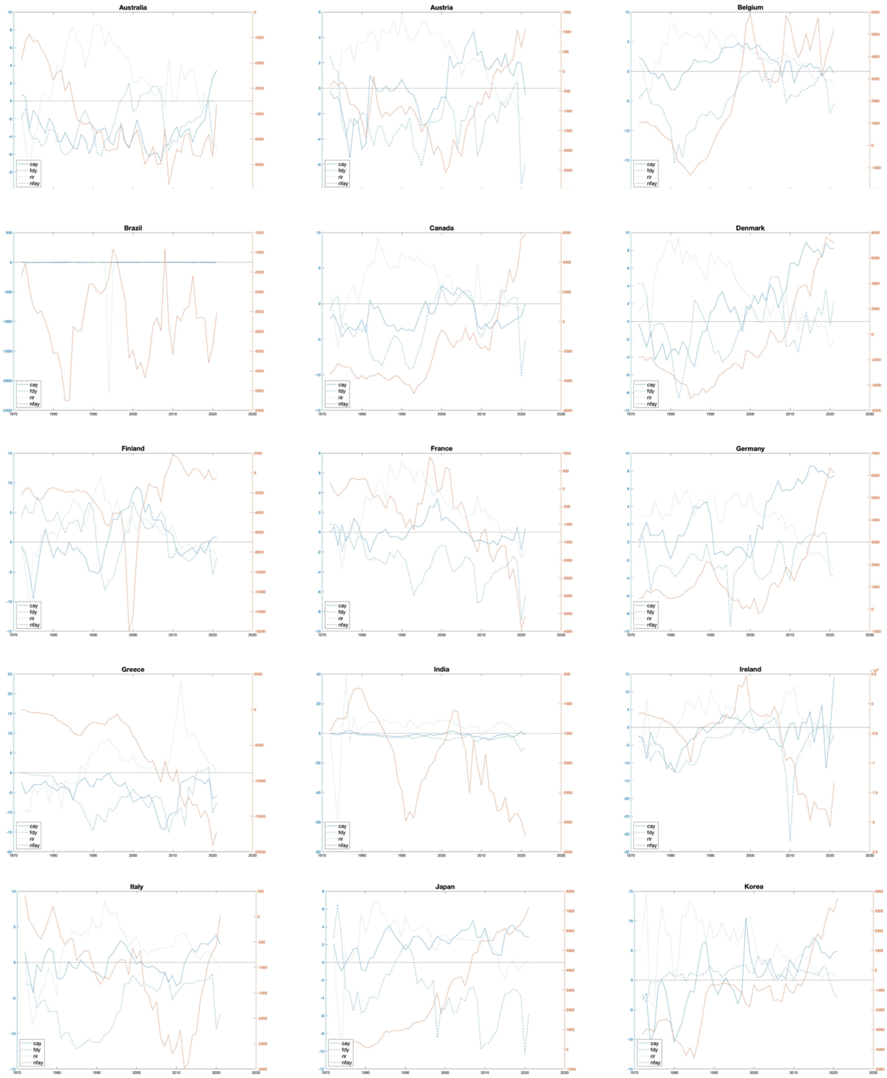


Fig. 1 Current account and some determinants. The solid blue line is the CA balance over GDP; the dashed blue line, the fiscal deficit over GDP; the dotted line is the real interest rate differential; and the solid orange line is the NFA position over GDP. (Color figure online)

components. The null hypothesis of weak cross-sectional dependence is rejected in most cases when the analysis focuses on individual countries and is rejected when all countries are considered together.

Fig. 1 continued

4 Empirical results

Concerning the empirical application, we apply well established unit root and cointegration tests, but more importantly, we follow a consistent step-by-step methodology for an accurate specification of the time series models that we are estimating:

- First, we test for the presence of structural breaks affecting the evolution of the variables using Perron and Yabu (2009) statistic that is robust to the order of integration of the variables.

- Second and depending on the previous results, we assess the order of integration (d) of the variables for all countries in our sample, using unit root tests either with or without structural breaks. This is of particular importance in the case of the dependent variable: in contrast to other studies, we do not estimate the model for some countries when the current account is found to be an integrated stochastic process of order zero ($I(0)$).
- Third, a similar approach is adopted to test for cointegration: starting by the simplest cointegration analysis without structural breaks—i.e., using the statistics in Engle and Granger (1987) and Shin (1994)—we also consider the case in which parameter instabilities might affect the deterministic component and/or the cointegrating vector—i.e., using the statistics in Gregory and Hansen (1996) and Carrion-i-Silvestre and Sansó (2006). If evidence of cointegration is found, the long-run relationship involving CA and its potential determinants is estimated. In some cases, parameter instabilities are considered, which implies dealing with two regimes—i.e., before and after the structural break.
- Finally, we calculate short-run multipliers by means of impulse response analysis, where the confidence bands are obtained considering the instabilities found. When the structural break affects the cointegrating vector, impulse responses are computed for each regime.[10]

4.1 Structural breaks and order of integration analysis

Visual inspection of the time series involved in the analysis reveals that, in most cases, the variables may have experienced the effect of structural breaks. It is well known that unaccounted-for structural breaks can bias the results of the order of integration analysis—see Perron (1989), among others. Perron and Yabu (2009) designed a methodology to test for the presence of a structural break on time series that is robust to the time series order of integration—i.e., robust conclusions about parameter stability can be obtained regardless of whether a generic time series y_t is integrated of order zero, $y_t \sim I(0)$, or of order one, $y_t \sim I(1)$. Panel B of Table 1 summarizes the results of the Perron and Yabu (2009) structural break test statistic considering the three different model specifications that can be assumed for the potential structural break effects—i.e., Model I establishes that the structural break only affects the level of the time trend, Model II considers that the structural break only affects the slope of the time trend, and Model III allows for the structural break to affect both the level and slope of the time trend. Note that the analysis is both country and variable-specific. The general qualitative conclusion that can be obtained is that the null hypothesis of no structural break is rejected in most cases, at least at the 10% level of significance and, at least, for one of the model specifications that are used. Detailed results of the computation of Perron and Yabu (2009) statistics can be found in Tables B.1 and B.2 in appendix.

Let us now focus on the results of the unit root statistics that have been computed. Generally, for the variables cay and fdy—defined in terms of the GDP—rir and rer,

[10] This has required the development of a GAUSS library that carries out the computations, which is available from the authors upon request.

the deterministic component assumed in the computation of the unit root tests is given by a (shifting) constant. The exception is China, where fdy shows a negative trend. Visual inspection reveals that a trending pattern is observed for $nfay$, $olddep$ and cay^*, so a (shifting) time trend is considered in these cases. It should be mentioned that, generally, the order of integration of the variables remains robust to the deterministic component that is specified once the potential presence of a structural break is accounted for.

The augmented Dickey–Fuller (ADF) version that is used might account for the presence of a structural break depending on the outcome of the Perron and Yabu (2009) procedure. For the no structural break case, the generalized least squares ADF (ADF-GLS) statistic is obtained as suggested in Ng and Perron (2001)—the order of autoregressive correction is selected by the modified Akaike's information criterion (AIC) with a maximum of 5 lags. The ADF statistic that considers one structural break is obtained as described in Perron and Vogelsang (1992)—one structural break affecting the level for non-trending variables (Model An)—and in Perron (1997)—one structural break affecting both the level and the slope of the time trend (Model C) for trending variables—where the order of the autoregressive correction is selected using the lags parameter individual significance (t-sig) strategy suggested by Ng and Perron (1995) with a maximum of 5 lags. Panel C of Table 1 reports qualitative conclusions about the order of integration analysis that has been performed using various unit root test statistics, depending on whether a structural break is considered or not.

The null hypothesis of a unit root cannot be rejected for most of time series, although there are some exceptions. For instance, the null hypothesis of unit root is rejected at the 5% significance level for the cay variable for France, Italy, New Zealand, Brazil, India, Korea and Russia. This result is important because it implies the exclusion of these countries from the cointegration analysis since cay is the dependent variable of the model. Similar results can be found for the other variables that are considered in our setup, which leads to excluding them as explanatory variables in the cointegration analysis that is conducted below. Consequently, the set of stochastic regressors that are defined for each country in the cointegration analysis might be different—for instance, for Austria we have $x_t = (nfay_t, fdy_t, rir_t, olddep_t, cay_t^*)'$, whereas for Germany $olddep_t$ and cay_t^* are excluded. Finally, for the real exchange rate, we reject the null hypothesis of unit root, which is not rejected for the real oil prices.

4.2 Cointegration analysis

The initial cointegration analysis is performed using the ADF test statistic of Engle and Granger (1987), which does not include the effect of structural breaks in the model—a constant gives the deterministic component. Except for Belgium, Switzerland and Mexico, Table 2 shows no evidence of cointegration at least at the 10% significance level. Similar qualitative conclusions are obtained with the computation of the statistic proposed by Shin (1994) that allows testing the null hypothesis of cointegration. Table 3 indicates that cointegration is rejected, at least at 10%, for all cases except for Ireland, the Netherlands, Portugal, the UK, Canada and Japan. Therefore, the overall evidence of a long-run equilibrium relationship is scarce.

Table 1 Summary of results

	Panel A		Panel B						Panel C					
	Pesaran CD test for *cay*		Evidence of structural break						Order of integration analysis (d)					
	Constant	Trend	cay	$nfay$	fdy	rir	$olddep$	cay^*	cay	$nfay$	fdy	rir	$olddep$	cay^*
AUT	−0.97	−1.78[c]	B	B		B	B	B	1	1	1	1	1	1
BEL	9.728[a]	10.19[a]	B		B	B	B		1	1	1	1	1	1
FIN	6.11[a]	5.05[a]		B	B	B	B	B	1	1	1	1	0	1
FRA	4.30[a]	4.40[a]		B		B	B	B	0	1	1	1	1	1
GER	2.06[a]	0.52	B	B	B	B	B		1	1	1	1	0	0
GRE	4.06[a]	4.61[a]	B	B	B	B	B	B	1	1	1	1	1	1
IRE	2.16[a]	2.13[a]	B		B		B		1	1	0	0	1	0
ITA	7.93[a]	7.13[a]		B		B	B		0	1	1	1	0	0
NET	1.80[c]	1.86[c]	B	B	B	B	B		1	1	1	1	1	1
POR	3.25[a]	3.77[a]	B	B		B	B	B	1	1	1	0	1	1
SPA	6.95[a]	6.50[a]	B	B	B	B		B	1	1	1	1	1	0
DNK	−0.11	0.39		B	B	B	B	B	1	1	1	1	1	1
SWE	4.47[a]	2.68[a]		B		B		B	1	1	0	1	1	0
UK	−2.63[a]	−1.89[c]		B		B	B		1	1	0	1	0	1
AUS	−0.70	−1.64[c]	B	B	B	B	B	B	1	1	1	1	0	0
CAN	−3.59[a]	−4.32[a]	B	B		B	B	B	1	1	0	1	1	1

Table 1 continued

	Panel A		Panel B						Panel C					
	Pesaran CD test for cay		Evidence of structural break						Order of integration analysis (d)					
	Constant	Trend	cay	$nfay$	fdy	rir	$olddep$	cay^*	cay	$nfay$	fdy	rir	$olddep$	cay^*
JAP	4.04[a]	3.43[a]		B		B	B	B	1	1	1	1	1	1
NZE	0.75	−0.40	B	B		B	B	B	0	1	0	1	1	0
SWI	3.40[a]	2.82[a]	B			B	B		1	1	1	1	1	0
USA	−2.26[a]	−1.13	B	B		B	B	B	1	1	0	1	0	1
BRA	3.88[a]	4.43[a]		B	B		B	B	0	1	1	0	1	1
IND	4.11[a]	3.68[a]		B	B	B	B	B	0	1	1	0	1	1
KOR	4.84[a]	3.76[a]	B	B	B	B		B	0	1	1	0	1	0
MEX	4.25[a]	3.26[a]	B	B	B	B	B	B	1	1	1	0	1	1
CHN			B	B			B		1	1	0	0	1	1
RUS			B	B	B	B			0	1	1	0	1	1
	All countries		Global variables						Global variables					
	6.92[a]	6.07[a]			rer						rer	0		
					$rpoil$	B					$rpoil$	1		

Superscripts a, b and c denote rejection of the null hypothesis at the 1, 5 and 10% significance levels, respectively

The cointegration analysis can be carried out by allowing for parameter instabilities in the model specification. First and following Gregory and Hansen (1996), the regression model in which the ADF cointegration statistic is based has been generalized with the inclusion of a structural break that affects either the level of the model (Model C) or both the level and the cointegrating vector (Model C/S). For instance, for Model C/S the specification is given by:

$$cay_t = \mu_0 + x_t'\alpha_0 + \omega_t'\beta_0 + \left(\mu_1 + x_t'\alpha_1 + \omega_t'\beta_1\right) DU_t + u_t,$$

with $DU_t = 1(t > T_b)$, where $1(\cdot)$ is the indicator function and T_b denotes the country-specific break date. The evidence against the null hypothesis of a spurious relationship is scarce—in Table 2 the null hypothesis is rejected at the 5% significance level for Belgium and Ireland (for both Model C and Model C/S specifications) and at the 10% significance level for Switzerland for Model C/S.

The picture changes when the cointegration statistic in Carrion-i-Silvestre and Sansó (2006) is used, which extends Shin (1994) cointegration analysis, allowing for one structural break. Now the null hypothesis of cointegration with one structural break affecting the level (Model An) or both the level and the cointegrating vector (Model D) cannot be rejected in most cases. For the latter, the model specification is given by:

Table 2 Engle–Granger and Gregory–Hansen cointegration test statistics

	Engle and Granger		Gregory and Hansen					
			Model C			Model C/S		
	ADF	p	ADF	p	$\hat{T}_b$	ADF	p	$\hat{T}_b$
AUT	−3.66	4	−5.32	0	1989	−7.19	1	1999
BEL	−6.31[b]	0	−8.39[a]	0	2014	−8.55[b]	0	2011
FIN	−3.22	0	−4.87	1	1994	−6.66	0	2000
FRA								
GER	−3.29	2	−4.19	2	1980	−4.75	0	1989
GRE	−3.42	0	−5.18	1	2014	−7.67	2	2001
IRE	−2.34	2	−6.30[b]	1	1985	−6.70[b]	1	1989
ITA								
NET	−5.05	2	−5.52	2	2014	−6.38	0	1995
POR	−4.02	1	−5.25	1	1995	−6.48	1	1993
SPA	−2.88	0	−5.44	2	2011	−5.70	4	1994
DNK	−2.93	0	−5.25	0	1988	−6.39	0	1999
SWE	−1.73	0	−4.82	0	1995	−5.65	0	1994
UK	−3.50	0	−4.41	0	1986	−5.28	1	1985
AUS	−3.31	0	−4.63	4	2013	−6.51	1	2009
CAN	−3.91	0	−5.08	0	1984	−5.85	0	1993
JAP	−5.13	3	−5.96	3	1989	−6.37	0	2003
NZE								
SWI	−5.31[c]	0	−6.00	0	1997	−7.36[c]	0	2000
USA	−3.81	0	−5.91	0	1984	−5.94	0	1987
BRA								
IND								
KOR								
MEX	−5.59[b]	1	−6.11	2	1988	−6.54	0	1990
CHN	−2.76	0	−4.10	3	1989	−4.96	1	2003
RUS								

Superscripts a, b and c denote rejection of the null hypothesis at the 1, 5 and 10% significance levels, respectively

$$cay_t = \mu_0 + x_t'\alpha_0 + \omega_t'\beta_0 + \left(\mu_1 + x_t'\alpha_1 + \omega_t'\beta_1\right) DU_t + \sum_{j=-k_2}^{k_1} \Delta z_{t-j}'\delta_j + u_t, \tag{1}$$

with Δz_t the vector with the first difference of the stochastic regressors, k_1 and k_2 the number of lags and leads that are included, selected using the Bayesian information criterion (BIC) over all possible combinations of $\{k_1, k_2\} \in \{0, 1\}$. Table 3 shows that the null hypothesis of cointegration is only rejected at the 5% of significance for

Table 3 Shin and Carrion-i-Silvestre and Sansó cointegration test statistics

	Shin test statistic				Carrion-i-Silvestre and Sansó test statistic									
	No structural breaks				Model An					Model D				
	SC	lags	leads	BIC	SC	$\hat{T}_b$	lags	leads	BIC	SC	$\hat{T}_b$	lags	leads	BIC
AUT	0.14[a]	0	0	1.41	0.04	1981	0	0	0.74	0.03	1999	0	1	1.16
BEL	0.12[b]	0	0	0.95	0.07[b]	1992	0	0	0.64	0.07[b]	2011	1	0	0.85
FIN	0.08[c]	0	0	2.67	0.06	1982	1	0	2.69	0.04[c]	1994	0	0	2.17
FRA														
GER	0.19[b]	0	0	2.04	0.10[b]	2001	0	0	1.48	0.03	2005	1	1	1.91
GRE	0.12[b]	0	0	2.75	0.05[c]	1994	0	0	2.24	0.03	2005	0	1	2.36
IRE	0.09	0	1	3.02	0.10[c]	2009	0	1	3.02	0.13[b]	2012	1	1	3.20
ITA														
NET	0.04	0	0	1.84	0.03	1995	1	0	1.91	0.02	1990	1	0	2.11
POR	0.06	0	0	3.00	0.03	2010	0	1	2.74	0.03	2010	0	1	3.22
SPA	0.14[b]	1	0	2.39	0.03	2012	1	0	1.63	0.03	2010	1	0	1.52
DNK	0.23[a]	0	0	2.28	0.04	1989	0	0	1.04	0.05[b]	1994	1	1	1.38
SWE	0.20[b]	1	1	1.98	0.08[c]	2006	1	1	1.40	0.04	2006	1	1	1.58
UK	0.09	0	0	1.20	0.06	1986	0	0	1.03	0.07[b]	1987	0	1	1.11
AUS	0.17[b]	0	0	1.66	0.05	2012	0	0	1.45	0.03	1990	0	1	1.04
CAN	0.03	0	0	0.97	0.04	1984	0	0	0.91	0.02	2005	0	0	1.31
JAP	0.04	0	0	0.90	0.04	2003	0	0	0.83	0.02	2000	0	1	0.63
NZE														
SWI	0.11[b]	0	0	2.80	0.07[c]	1996	0	0	2.66	0.02	2006	1	0	2.79
USA	0.17[b]	1	0	0.01	0.07	2012	1	0	−0.50	0.03	1998	1	0	−0.38
BRA														
IND														
KOR														
MEX	0.08[c]	0	0	1.68	0.05	1985	0	0	1.31	0.03	1985	0	0	1.59
CHN	0.12[b]	0	0	2.04	0.04	1989	0	0	1.76	0.04	1995	0	0	1.94
RUS														

Superscripts a, b and c denote rejection of the null hypothesis at the 1, 5 and 10% significance levels, respectively

Belgium, regardless of the model specification that is used. For the other countries, we find cointegration at least for one of the model specifications—evidence of long-run relationships is reduced in some cases when working at the 10% significance level. In some cases, cointegration is found regardless of the model specification, in which case the BIC is used to select one of the specifications. Overall, there is evidence of cointegration for 17 out of 19 countries. The overwhelming difference between Gregory and Hansen (1996) and Carrion-i-Silvestre and Sansó (2006) test statistics results might be due to the asymmetric treatment of the structural break done in the former proposal. Thus, Gregory and Hansen (1996) only allows for one structural break under the alternative hypothesis of cointegration—i.e., the null hypothesis is a joint hypothesis of no cointegration and no structural breaks—whereas Carrion-i-Silvestre and Sansó (2006) considers a structural break under both the null and alternative hypotheses.

Additional evidence of structural breaks in the long-run relationship is obtained from the computation of the Wald statistic proposed by Kejriwal and Perron (2010). Table 4 presents the sup F statistic to test the null hypothesis of no structural break against the alternative hypothesis of one structural break. The model specification has been selected by the BIC statistic. The table also reports the estimated break date that maximizes the sequence of Wald statistics ($\hat{T}_b^F$), the one that minimizes the sum of the squared residuals (SSR) of (1)—i.e., $\hat{T}_b^{SSR}$, which coincides with the ones provided in Table 3—and the BIC statistic of the model with and without one structural break—i.e., BIC_1 and BIC_0, respectively. The null hypothesis of no structural break is rejected at the 5% significance level in most cases, except for Ireland, the UK and Canada. Note that for the UK and Canada, the BIC statistic selects the model that includes the structural break as the preferred one. This leads us to consider both possibilities in the subsequent analysis to obtain a better global picture of the convenience of allowing for one structural break in the estimated model for Ireland, the UK and Canada.

4.3 Estimation of the current account long-run relationship

Tables 5, 6 summarize the dynamic ordinary least squares (DOLS) estimates of the model specification for which the null hypothesis of cointegration is not rejected at the 5% significance level—in those cases for which cointegration is found for both Models An and D specifications, the estimates reported are the ones selected by the BIC statistic that appears in Table 3.

Let us first focus on the estimation results without structural breaks. Table 5 shows that the Irish CA is driven by $olddep$ and $rpoil$, with a negative contribution to CA. The CA is driven by cay^* for the UK, whereas Canadian CA is determined by $nfay$, rir, cay^* and, at the 10% significance level, $rpoil$. Note that the constant term is not statistically significant at the 10% level for Canada—which would indicate a non-systematic deviation from the external deficit sustainability condition—whereas it is positive (negative) and statistically significant for Ireland (the UK).

The estimation results with one structural break can be found in Table 6. First, it is worth noting the heterogeneity that characterizes the estimated break dates obtained. Thus, even for the European (and eurozone) countries, implementing the estimation procedure does not point to a common structural break. This might reflect quite idiosyncratic behavior regarding the evolution of the CA. Second, for 13 out of 19 countries, the preferred model specification only includes a level shift (Model An), which indicates that for these countries, the cointegrating vector is stable throughout the time period. This is not the case for Germany, Spain, Sweden, Australia, Japan and Switzerland, for which the effect of the CA determinants might have changed during the analyzed period.

The effect of $nfay$ is positive for Greece, Portugal, Australia, Japan, the USA and China. In contrast, it is negative for Finland, the Netherlands, Sweden (after 2006),

Table 4 Kejriwal and Perron $supF$ statistic

	Model	sup F	$\hat{T}_b^F$	$\hat{T}_b^{SSR}$	BIC_0	BIC_1
AUT	An	16.72[a]	1981	1981	1.41	0.74
BEL	An	15.07[a]	1991	1992	0.95	0.64
FIN	An	10.84[c]	1982	1982	2.67	2.69
FRA						
GER	D	24.64[a]	2005	2005	2.04	1.91
GRE	An	10.45[c]	1994	1994	2.75	2.24
IRE	An	4.98	2000	2009	3.02	3.20
ITA						
NET	An	16.65[a]	1995	1995	1.84	1.91
POR	An	30.17[a]	2010	2010	3.00	2.74
SPA	D	34.25[a]	2004	2010	2.39	1.52
DNK	An	34.46[a]	1988	1989	2.28	1.04
SWE	D	30.86[a]	2006	2006	1.98	1.58
UK	An	4.12	1986	1986	1.20	1.03
AUS	D	40.46[a]	1996	1990	1.66	1.04
CAN	An	7.45	1984	1984	0.97	0.91
JAP	D	31.64[a]	1999	2000	0.90	0.63
NZE						
SWI	D	30.62[a]	2006	2006	2.80	2.79
USA	An	10.42[c]	1998	2012	0.01	−0.50
BRA						
IND						
KOR						
MEX	An	12.90[b]	1985	1985	1.68	1.31
CHN	An	12.64[b]	2004	1989	2.04	1.76
RUS						

Superscripts a, b and c denote rejection of the null hypothesis at the 1, 5 and 10% significance levels, respectively

Table 5 DOLS estimation of the cointegrating relationship without structural breaks

		Parameter associated to regressor					
	μ	$nfay$	fdy	rir	$olddep$	cay^*	$rpoil$
IRE	33.21	−0.03			−1.65		−0.16
	(4.72)	(−1.63)			(−3.78)		(−6.55)
UK	−1.50	0.01		0.04		−0.97	0.00
	(−1.79)	(0.27)		(0.36)		(−2.82)	(0.23)
CAN	−2.66	0.05		0.17	0.11	−0.66	0.01
	(−1.24)	(3.03)		(2.29)	(0.93)	(−3.90)	(1.75)

Table 6 DOLS estimation of the cointegrating relationship with one structural break. Models An or D, selected by the BIC

	$\hat{T}_b$	Parameters for the whole period							Parameters associated to the structural change						
		μ_1	$nfay$	fdy	rir	$olddep$	cay^*	$rpoil$	μ_2	$nfay$	fdy	rir	$olddep$	cay^*	$rpoil$
AUT	1981	9.82	−0.03	0.41	−0.57	−0.56	0.25	0.04	3.99						
		(3.04)	(−1.10)	(4.17)	(−3.31)	(−4.34)	(1.57)	(5.78)	(7.12)						
BEL	1992	12.53	0.01	0.16	0.20	−0.43	0.04	−0.05	2.92						
		(4.28)	(0.39)	(2.43)	(2.99)	(−3.18)	(0.20)	(−6.38)	(4.84)						
FIN	1982	−5.12	−0.14	0.15	0.54		2.26	0.03	−4.85						
		(−3.24)	(−9.00)	(1.21)	(2.69)		(4.56)	(1.66)	(−2.44)						
GER	2005	−0.60	0.17	−0.48	−0.20			−0.01	5.57	−0.22	−0.10	−0.05			0.01
		(−0.37)	(5.39)	(−2.45)	(−0.69)			(−1.00)	(2.04)	(−4.68)	(−0.47)	(−0.13)			(0.46)
GRE	1994	−7.75	0.07	0.31	0.40	0.49	1.57	−0.03	−6.79						
		(−1.18)	(2.16)	(3.43)	(5.49)	(1.51)	(2.90)	(−2.15)	(−6.35)						
IRE	2009	33.19	−0.01			−1.68		−0.14	5.93						
		(5.18)	(−0.57)			(−4.24)		(−6.09)	(2.49)						
NET	1995	−19.15	−0.08	−1.13	−0.84	1.29	0.96	0.02	−3.41						
		(−5.44)	(−3.77)	(−5.47)	(−4.78)	(5.87)	(2.00)	(1.67)	(−3.16)						
POR	2010	0.55	0.08	−0.40		−0.11	−0.34	−0.02	13.52						
		(0.16)	(3.79)	(−2.04)		(−0.57)	(−0.66)	(−0.66)	(5.68)						

Table 6 continued

	$\hat{T}_b$	Parameters for the whole period							Parameters associated to the structural change						
		μ_1	$nfay$	fdy	rir	$olddep$	cay^*	$rpoil$	μ_2	$nfay$	fdy	rir	$olddep$	cay^*	$rpoil$
SPA	2010	1.18	0.04	−1.32	−0.44	−0.15		−0.01	−167.42	−0.12	1.10	3.05	5.68		−0.01
		(0.44)	(2.28)	(−10.17)	(−6.98)	(−1.09)		(−1.14)	(−4.99)	(−2.01)	(7.97)	(4.55)	(5.15)		(−0.22)
DNK	1989	−15.68	0.00	0.01	−0.18	0.63	−0.66	−0.01	4.68						
		(−4.20)	(0.16)	(0.12)	(−1.27)	(3.98)	(−2.25)	(−1.00)	(11.21)						
SWE	2006	11.68	0.02		−0.27	−0.31		0.01	−18.00	−0.16		0.46	0.69		0.04
		(1.57)	(0.56)		(−1.50)	(−1.10)		(1.29)	(−1.42)	(−2.79)		(0.84)	(1.48)		(1.92)
UK	1986	−0.24	−0.04		0.21		−0.81	−0.00	−2.12						
		(−0.31)	(−1.06)		(1.93)		(−2.94)	(−0.27)	(−2.83)						
AUS	1990	−0.08	0.09	−0.00	−0.08			−0.01	19.83	0.29	−0.75	−0.70			−0.03
		(−0.04)	(1.70)	(−0.00)	(−0.53)			(−1.29)	(6.22)	(4.14)	(−3.25)	(−4.66)			(−2.69)
CAN	1984	−7.88	0.01		0.19	0.46	−1.17	−0.00	−2.97						
		(−3.05)	(0.38)		(2.89)	(2.89)	(−5.11)	(−0.50)	(−2.93)						
JAP	2000	9.37	0.31	−0.05	0.42	−0.66	−1.94	−0.07	−17.66	−0.16	0.35	0.67	0.79	0.51	0.04
		(7.94)	(5.21)	(−1.20)	(10.15)	(−5.55)	(−4.88)	(−9.28)	(−3.19)	(−2.75)	(3.57)	(3.56)	(3.74)	(0.92)	(2.40)
SWI	2006	−71.93	−0.00	0.09	−0.48	3.62		0.01	111.56	−0.15	−0.73	−0.71	−4.48		0.09
		(−7.26)	(−0.10)	(0.43)	(−1.62)	(6.57)		(0.70)	(4.83)	(−4.27)	(−1.09)	(−0.90)	(−4.86)		(3.66)
USA	2012	−2.54	0.03		0.13		−1.61	−0.01	2.18						
		(−7.70)	(2.45)		(2.68)		(−13.07)	(−2.67)	(5.48)						
MEX	1985	−10.79	−0.14	−0.17		0.70	−1.26	−0.03	−3.81						
		(−5.17)	(−4.23)	(−2.15)		(2.37)	(−6.12)	(−3.00)	(−5.11)						
CHN	1989	−0.49	0.16			−0.02	−0.87	−0.01	3.01						
		(−0.19)	(3.24)			(−0.06)	(−1.81)	(−0.60)	(3.97)						

Switzerland (after 2006) and Mexico—the significance level is set at 10%. The sign of the effect has switched from positive to negative for Germany and Spain—i.e., an initial positive effect on CA turned out to become negative after the structural break. This heterogeneity is related to the different position (debtor or creditor) of the countries in the sample and their role in the international financial markets. In addition, during the sample period, some countries worsened, and others improved their relative position. Most significant changes have implied a worsening after the breaks and, therefore, a negative sign.

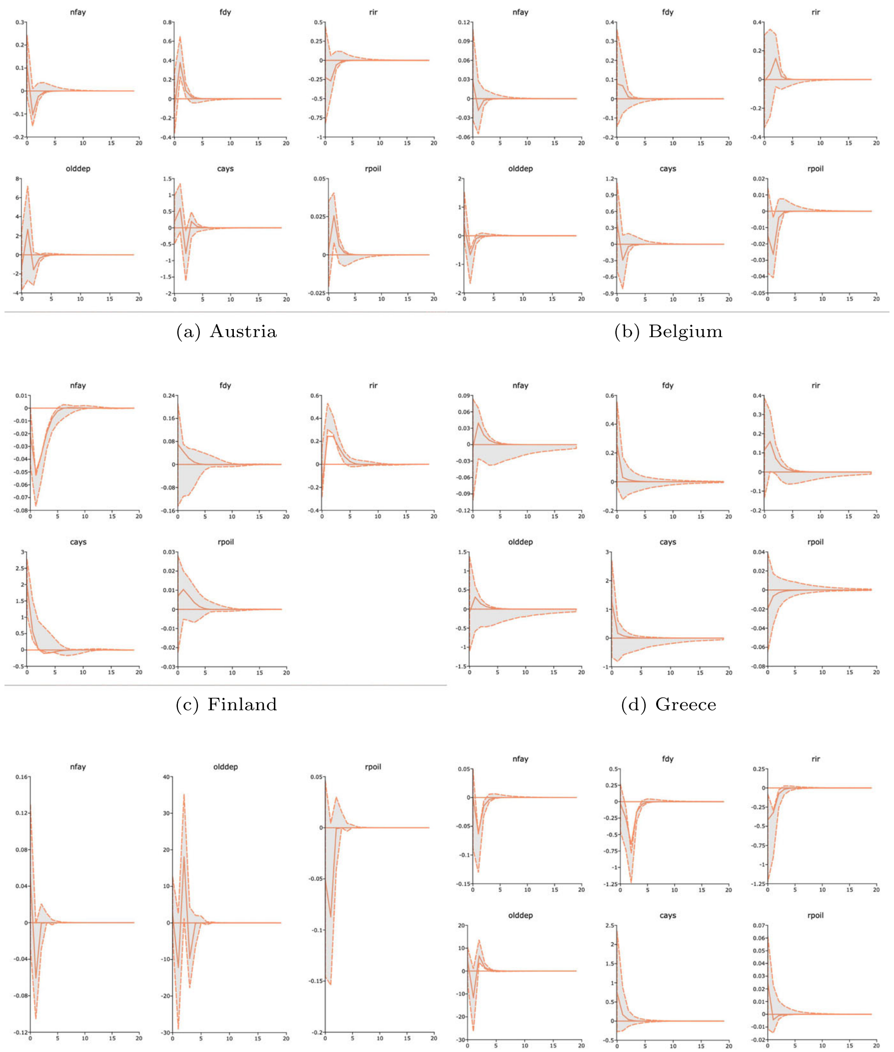
If we focus on the effect of the fiscal deficit, fdy, the empirical evidence supports the twin deficit hypothesis (a positive relationship between the current account and the fiscal balance) for Austria, Belgium, Greece and Japan (after 2000). For Germany, the Netherlands, Portugal, Spain, Australia (after 1990) and Mexico, the effect is negative.

According to the theory, the sign linking the current account and the real interest rate differential (rir) is uncertain. Higher interest rates appreciate the exchange rate and have a negative effect on the current account; however, this will reduce domestic demand, and this contractionary effect may improve the current account balance. Therefore, the final effect will depend on which one prevails. Our results concerning this variable are mixed. The real interest rate produces a significant negative effect on the CA for Austria and the Netherlands, whereas the opposite is found for Belgium, Finland, Greece, the UK, Canada, Japan and the USA. For Spain, the initial negative effect became positive after 2010.

The parameter of the population dependence ($olddep$) is statistically significant in few cases. According to the theory, an increase in this variable deteriorates the current account as the aged population decreases savings. On the other hand, the effects of demography in the current account are much more complex. Life expectancy has also increased in the countries included in the analysis, mostly OECD members. The latter has changed the saving behavior of middle-aged, as Dao and Jones (2018) points out. Therefore, even if a more significant proportion of the population reaches old age, their savings may have increased during their lifespan. Indeed, we have found that the sign and magnitude of the parameters are quite heterogeneous. The estimated effect is negative for Austria, Belgium and Ireland, whereas for the Netherlands, Denmark, Canada and Mexico, it is positive. For Japan, the initial negative effect changes to a positive determinant (after 2000), and the opposite is found for Switzerland (after 2006).

Concerning the real oil price ($rpoil$), its effect is positive for Austria, Finland, the Netherlands, Sweden (after 2006) and Switzerland (after 2006), and negative for Belgium, Greece, Ireland, Australia (after 1990), Japan, the USA and Mexico.

Finally, the parameter of cay^* is positive and statistically significant, at least at the 10% significance level, for Finland, Greece and the Netherlands, whereas it represents a negative and statistically significant effect for Denmark, the UK, Canada, Japan, the USA, Mexico and China.



(a) Austria (b) Belgium

(c) Finland (d) Greece

(e) Ireland (f) Netherlands

Fig. 2 Multiplier analysis. Model An, BIC-based order selection model

4.4 Multiplier analysis

This section describes the estimation of the error correction mechanism (ECM) equation *à la* Engle and Granger (1987), but imposing the DOLS estimated cointegration vector that has been obtained in the previous section in the definition of the error correction term. This avoids assuming that stochastic regressors are weakly exogenous when conducting the estimation of the single-equation ECM. Considering the most

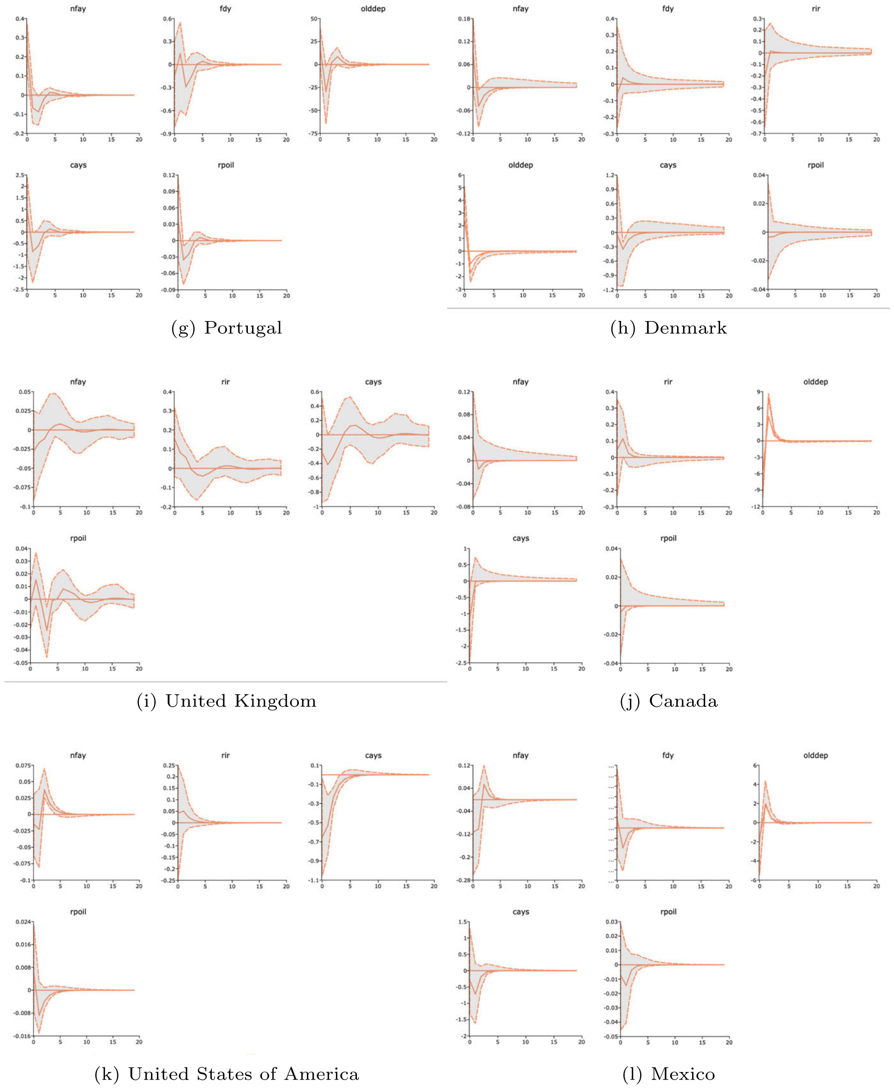
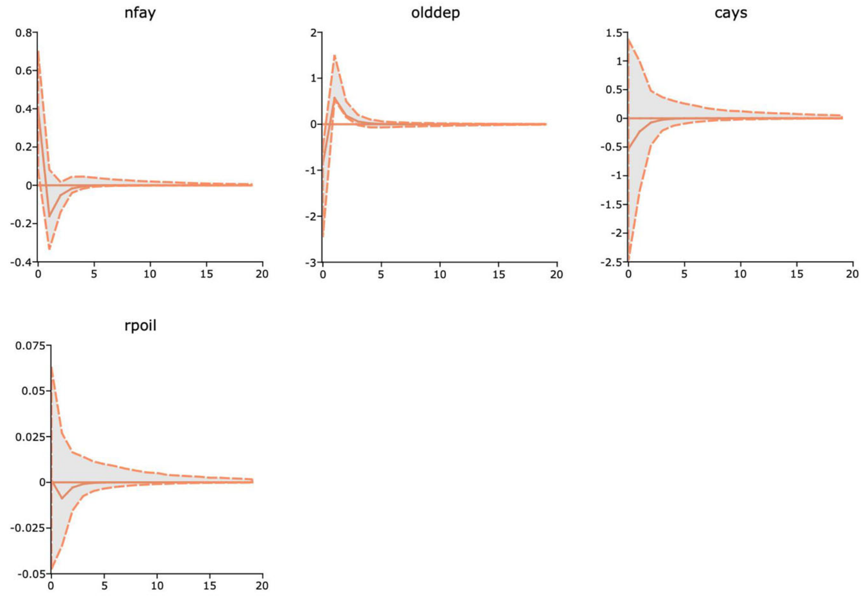


Fig. 2 continued

general specification of Model D, the error correction term is given by:

$$\hat{e}_t = cay_t - (\hat{\mu}_0 + x_t'\hat{\alpha}_0 + \omega_t'\hat{\beta}_0) - (\hat{\mu}_1 + x_t'\hat{\alpha}_1 + \omega_t'\hat{\beta}_1)DU_t, \tag{2}$$

where the estimated parameters are the ones from the DOLS estimation of (1)—for Model An, $\hat{\alpha}_1 = \hat{\beta}_1 = 0$ in (2). The single-equation ECM specification for generic



(m) China

Fig. 2 continued

orders p_x, p_ω and p_c is given by:

$$\Delta cay_t = \sum_{j=0}^{p_x} \Delta x'_{t-j} A_j + \sum_{j=0}^{p_\omega} \Delta \omega'_{t-j} B_j + \sum_{j=1}^{p_c} \Delta cay_{t-j} H_j + \delta \hat{e}_{t-1} + v_t, \qquad (3)$$

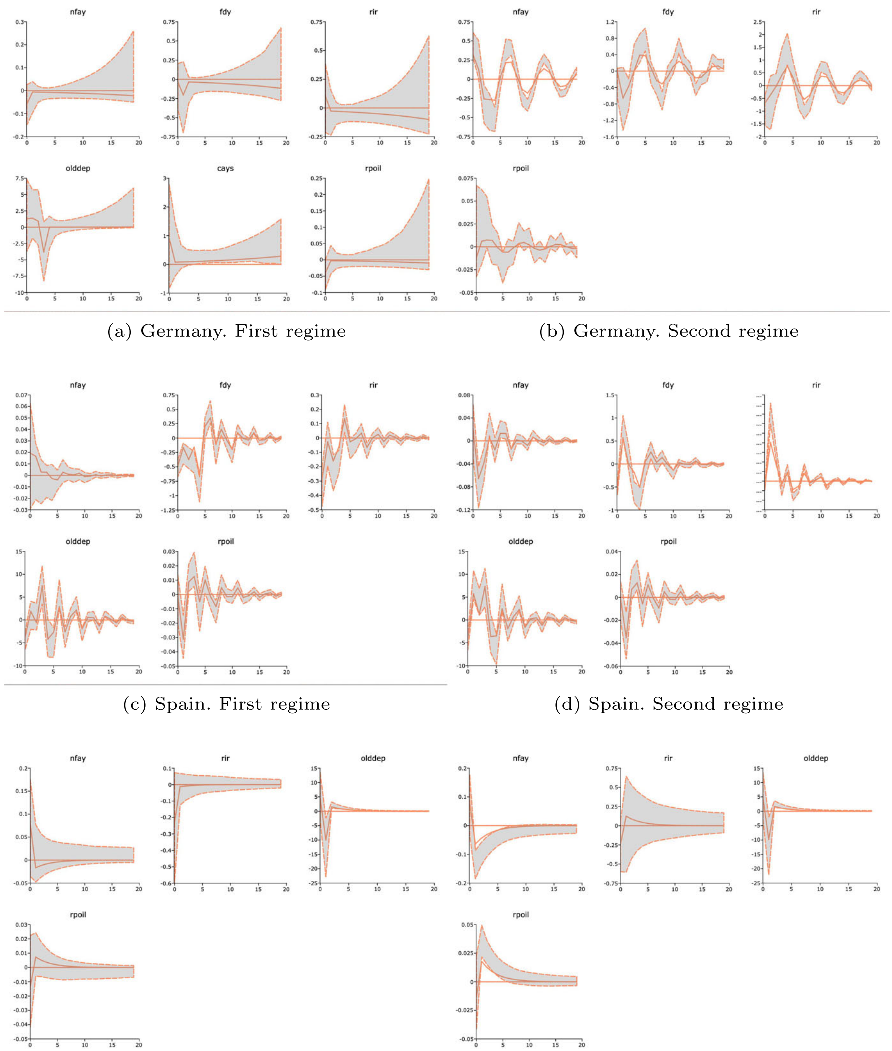
a model specification that is estimated by OLS with the (variable-specific) number of lags that are selected by the BIC, allowing for a maximum of $p_{\max} = 4$ lags—note that all potential combinations of lag orders have been checked for each regressor. Table 7 collects the selected orders, the R^2 and $\bar{R}^2$ adjustment coefficients, the Breusch–Godfrey Lagrange multiplier (LM) autocorrelation statistics (up to order 2) and the Jarque–Bera (JB) normality statistic of the estimated model given in (3). Except for Spain and Japan, the estimated order of the dynamic component does not reach the maximum of lags for any regressor. The degree of fit varies among models with values of $R^2 \in [0.34, 0.90]$ and $\bar{R}^2 \in [0.24, 0.82]$. In general, there is no evidence of autocorrelation in the disturbance term[11] and the Jarque–Bera statistic does not reject the null hypothesis that the disturbance term follows a normal distribution in most cases—the exceptions are found for Germany, Ireland and Portugal.

Equations (2) and (3) can be rewritten in terms of an ARDL model specification—see, for instance, Lütkepohl (2005), pp. 249, and Pesaran (2015a), pp. 124—from

[11] The null hypothesis of no autocorrelation is not rejected against the alternative hypothesis of autocorrelation of order 1 (see the p-values for the LM(1) statistic in Table 7) in any case, although it is rejected against the alternative of autocorrelation of order 2 (LM(2)) for the Netherlands and Japan.

Table 7 BIC-based selected orders and specification statistics for the estimated ECM

	Model	cay	$nfay$	fdy	rir	$olddep$	cay^*	$rpoil$	R^2	$\bar{R}^2$	$LM(1)$	p-val	$LM(2)$	p-val	JB	p-val
AUT	An	0	0	0	0	1	2	0	0.72	0.65	0.64	0.42	1.46	0.48	3.12	0.21
BEL	An	0	0	0	1	0	0	0	0.58	0.51	0.03	0.85	0.23	0.89	1.84	0.40
FIN	An	1	0	0	0		0	0	0.62	0.56	0.17	0.68	0.57	0.75	0.00	1.00
GER	D	2	3	0	3			0	0.62	0.48	1.18	0.28	1.18	0.55	9.13	0.01
GRE	An	0	0	0	0	0	0	0	0.52	0.45	2.40	0.12	4.06	0.13	0.00	1.00
IRE	An	0	0			2		0	0.60	0.55	0.41	0.52	1.15	0.56	19.72	0.00
NET	An	0	0	1	0	1	0	0	0.65	0.58	1.59	0.21	7.51	0.02	0.09	0.96
POR	An	1	0	1		1	0	0	0.53	0.44	1.04	0.31	1.69	0.43	6.38	0.04
SPA	D	4	0	4	2	4		1	0.90	0.82	1.49	0.22	2.99	0.22	0.40	0.82
DNK	An	0	0	0	0	0	0	0	0.53	0.47	0.24	0.63	0.24	0.89	4.97	0.08
SWE	D	0	0		0	1		0	0.41	0.34	3.11	0.08	3.99	0.14	0.00	1.00
UK	An	2	0		0		0	3	0.60	0.50	1.85	0.17	2.87	0.24	1.78	0.41
AUS	D	2	0	0	0			0	0.34	0.24	0.04	0.84	0.40	0.82	0.00	1.00
CAN	An	0	0		0	0	0	0	0.51	0.45	0.09	0.76	0.71	0.70	0.90	0.64
JAP	D	4	0	0	0	2	0	0	0.53	0.36	3.68	0.06	15.21	0.00	0.46	0.80
SWI	D	0	0	1	1	0		0	0.56	0.49	0.76	0.38	3.19	0.20	3.72	0.16
USA	An	0	1		0		0	0	0.64	0.60	0.27	0.61	0.67	0.72	0.03	0.99
MEX	An	0	1	0		0	0	0	0.63	0.57	3.84	0.05	3.92	0.14	5.64	0.06
CHN	An	0	0			0	0	0	0.50	0.44	0.21	0.64	0.63	0.73	0.00	1.00



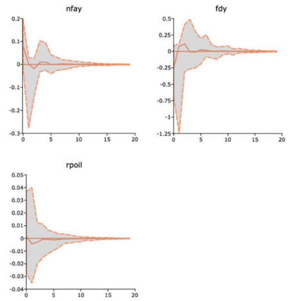
(a) Germany. First regime (b) Germany. Second regime

(c) Spain. First regime (d) Spain. Second regime

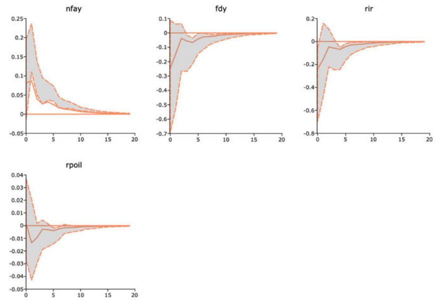
(e) Sweden. First regime (f) Sweden. Second regime

Fig. 3 Multiplier analysis. Model D, BIC-based order selection model

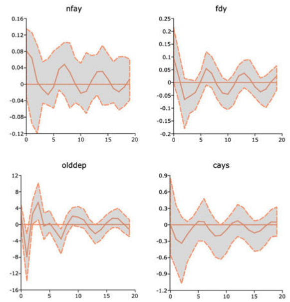
which the short-run multipliers of the different explanatory variables can be obtained—the long-run multipliers are given by the cointegrating vector. It is worth noting that the short-run dynamics (A_j, B_j and H_j) and the speed of adjustment coefficient (δ) are assumed to be constant through time—i.e., the structural break affects the elements inside the error correction term. This is not restrictive for Model An since this specification assumes a change in the level of the long-run relationship. In this case, the multiplier analysis is constant throughout the time period. To be specific, the



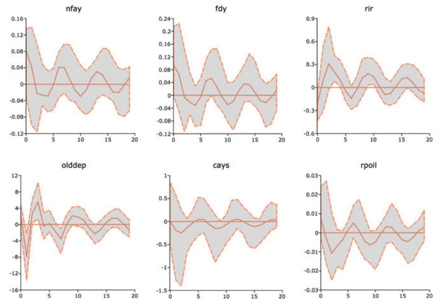
(g) Australia. First regime



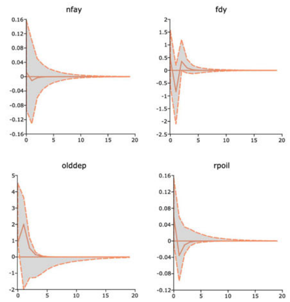
(h) Australia. Second regime



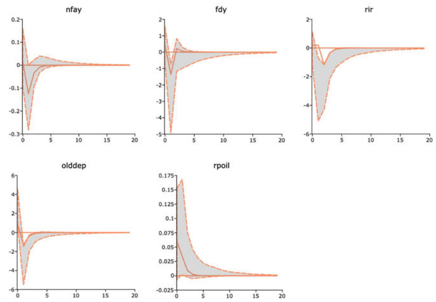
(i) Japan. First regime



(j) Japan. Second regime



(k) Switzerland. First regime



(l) Switzerland. Second regime

Fig. 3 continued

model in levels that derives from Model An is given by:

$$cay_t = \sum_{j=0}^{p_x+1} x'_{t-j}\hat{a}_j + \sum_{j=0}^{p_\omega+1} \omega'_{t-j}\hat{b}_j + \sum_{j=1}^{p_c+1} cay_{t-j}\hat{h}_j + \hat{w}_t, \tag{4}$$

with $\hat{a}_0 = \hat{A}_0$, $\hat{a}_{p_x+1} = -\hat{A}_{p_x}$, $\hat{a}_j = \hat{A}_j - \hat{A}_{j-1}$ for $j = 2, \ldots, p_x - 1$ and $\hat{a}_1 = -\hat{\delta}\hat{\alpha}_0 - \sum_{j=2}^{p_x+1} \hat{a}_j$. The same transformations are applied to obtain $\hat{b}_j$ and $\hat{h}_j$

coefficients, with the exception that $\hat{h}_1 = \hat{\delta} + 1$ when $p_c = 0$, and $\hat{h}_1 = \hat{\delta} + 1 + \hat{h}_2$ when $p_c > 0$.

The situation is slightly different for Model D since the change in the cointegrating vector induces a change in the coefficient of the first lag of the regressors of the ARDL model representation—i.e., $\hat{a}_1 = -\hat{\delta}\hat{\alpha}_0 - \sum_{j=2}^{p_x+1} \hat{a}_j$ for the first regime ($t \leq T_b$) and $\hat{a}_1 = -\hat{\delta}(\hat{\alpha}_0 + \hat{\alpha}_1) - \sum_{j=2}^{p_x+1} \hat{a}_j$ for the second regime ($t > T_b$), with $\hat{a}_j$ for $j = 0, \ldots, p_x + 1$ but $j \neq 1$, as defined above. The same applies to $\hat{b}_j$ coefficients. Therefore, the changing cointegrating vector leads to a change in the dynamics of the corresponding ARDL model that involves the variables in levels.

Once the ARDL model specification is obtained, the multiplier analysis can be easily carried out to measure the effect of a unit change of a given regressor over the dependent variable. The computation of the short-run multipliers is accompanied by 95% bootstrap confidence intervals. In this regard, we have computed Hall (1992) studentized percentile interval using 1000 bootstrap replications of the resampled estimated residuals—parametric Gaussian iid bootstrap and wild Gaussian bootstrap produce similar results. It is worth noting that a bootstrap-after-bootstrap bias correction type as described in Kilian (1998) is implemented to account for the estimation bias that is expected to appear when dealing with dynamic models in finite samples—see Appendix C for further details.

Figure 2 provides the estimated multiplier analysis plots for the countries for which Model An has been selected. In the interest of brevity, we do not present the estimated coefficients of the short-run multipliers directly in the text. However, these coefficients can be obtained from the authors upon request. Below, we comment on the main findings derived from the multipliers' analysis for the various countries under consideration. To ensure conciseness, we initially focus on providing a more detailed account of the results for the first country in our sample that displays a level shift (Austria). Subsequently, we offer a more generalized overview of the common features observed among the remaining countries in this subgroup of 13 economies. For Austria, the contemporaneous effect of a change in $nfay$ over cay is positive, although it quickly represents negative effects in the subsequent periods. Notwithstanding, these effects are not statistically significant at the 5% significance level—we also find (see Table 6) that the parameter associated with $nfay$ in the cointegrating relationship is not statistically significant. In the short run, a change on fdy only has statistically significant positive effects on cay after one (0.376) and two (0.086) periods, which is in accordance with the estimated long-run effect (0.41). Changes of rir and $olddep$ have not statistical significant effects on cay, whereas a change in the external position of the other countries (cay^*) only affects cay after two periods. Finally, an increase in the real price of oil improves the current account position of Austria for the next period, with an amount of 0.025 units. In the case of Belgium, only the changes in $olddep$ and $rpoil$ are statistically significant after one period, presenting mixed results between internal and external drivers. This is also the case of the Netherlands, Portugal or Denmark, where the variable $olddep$ seems to play a key role shared with the variable $nfay$ or, in Mexico, where $olddep$ stands out for its explanatory power in the short run. Something similar happened in China, where the parameter is not significant in the long run but has a negative instantaneous effect that becomes positive and disappears after the second period. While this variable also shows a relevant role in the case of Canada or Sweden, other external drivers also play a vital influence in determining the cay: financial factors, represented by rir, and external current account (cay^*) in the first case, and $nfay$, cay^* and $rpoil$ for the case of Sweden. Finally, for the USA and the UK, only the external drivers affect the short-run dynamics of cay. While in the case of the USA, $nfay$ and cay^* are statistically significant and present some inertia, as they affect the cay after three or four periods, for the UK the only external explanatory variable is, not surprisingly, the real oil prices ($rpoil$). For China, cay^* was significant in the long run but not significant short-run effect is found in the impulse response analysis.

It is worth noting that in eight out of the thirteen countries considered, the most relevant internal driver is $olddep$. Only in one case (the Netherlands) does the second internal driver, namely fdy, show important and persistent effects over time. (The effects are significant after three periods and present the expected sign.) On the contrary, when we focus on the external drivers, both financial ($nfay$) and real ones (cay^*) are significant in six and seven cases, respectively. Moreover, in most cases, both drivers show a combined dynamic effect on the cay of the incumbent countries. A global external driver, $rpoil$, is significant in six countries, according to the relative dependence on oil of the countries analyzed.

As for the countries exhibiting a changing cointegrating vector, Fig. 3 shows the short-run multipliers for each regime.[12] The degree of heterogeneity between the two periods depends on the specific country analyzed. For example, in the case of Germany, the internal driver linked to the twin deficit hypothesis seems to hold in the short run. However, the parameter of the demographic variable $olddep$ is not significant either before or after the break. In the case of Spain, internal and external drivers are at play, and the most striking change is the increasing importance of the financial driver $nfay$ after the break date, coinciding with the European debt crisis around 2010. This is also the case for Sweden, Australia and Switzerland for different reasons. Two special cases are Japan and Greece. For Japan, the most important drivers are the demographic one ($olddep$), which shows a high persistence for up to seven periods, and real oil prices ($rpoil$), especially before the break in 1990, with an inertia that lasts up to four periods. Finally, in the case of Greece, the cay is driven by external factors, cay^* with high persistence and the real oil prices ($rpoil$), only with contemporaneous effects before the structural break. However, after the euro's inception, the real interest rate surged as a very influential driver with persistent effects for up to five periods.

5 Summary and conclusions

Current account imbalances often reflect cross-country differences in saving and investment, aligned with economic fundamentals and international flows of goods and

[12] In some cases, as in the response during the first regime of $nfay$ and rir in Germany, some lags are outside the significance bands. The reason for this finding is the choice of BIC to select the lags. This criterion provides a more parsimonious number of lags, but sometimes, they are not enough to recover the right dynamics. We have also obtained the lags using AIC, and the response stays within the bands for all the periods considered. These results are available from the authors upon request.

finance. However, they can also arise from economic and financial distortions, signaling vulnerabilities in the short run. The significance of these imbalances depends on the nature of the underlying shocks that drive them. In this paper, we study the long-run determinants of current account balances to identify the structural factors explaining the persistent imbalances we observed in the last few decades. In addition to examining traditional determinants of the current account, it advances the existing empirical literature in different respects. First, by using a comprehensive annual dataset spanning from 1972 to 2021 for a group of 26 developed and emerging economies—shorter for China and Russia. Second, we unveil the relative importance of internal and external drivers of the current account. In particular, as a distinctive feature, we introduce in the econometric model a variable measuring the external disequilibria in third countries to capture their effect on the magnitude and persistence in each country's current account. Third, from an econometric point of view, to uncover the dominant drivers of current account movements at different frequencies, our study incorporates cointegration analysis with consideration for one structural break, enabling the exploration of changes in the relevance of determinants over time. Notably, the research identifies parameter instabilities, demonstrating their pivotal role in modeling external imbalances while highlighting the explicit heterogeneity in countries' responses to estimated break dates. Fourth, a salient characteristic of our study is that specifying the estimated single-equation ECM model in terms of the equivalent ARDL model representation for each country has allowed us to obtain the dynamic multipliers of the explanatory variables. The paper distinguishes two subsamples when the cointegrating vector that defines the long-run relationship has experienced the effect of a structural break. Finally, the findings underscore the increasing significance of the external drivers, emphasizing its substantial impact on current account imbalances across the examined countries.

As for the long-run relationships, we have found evidence that supports a (changing) long-run relation for the nineteen countries that have been analyzed—the initial sample of twenty-six countries has been reduced since for seven countries, the dependent variable is found to be I(0). An interesting trait is the heterogeneity that characterizes the estimated break dates, which tend to be country-specific. As already mentioned, another distinctive contribution of the paper is that the model specification has accounted for the influence of the current account of trade-related foreign countries, which plays the role of a common factor that captures the effect of global foreign external balance on the domestic one. The other most relevant external driver is the real oil price. This variable is relevant for the cointegrating vector in twelve cases out of nineteen; these tend to be aligned with the openness of the economies. When we focus only on the fiscal and demographic internal drivers, there are only nine significant cases. Our results point to the existence of differences between the short-run current account determinants and those in the long run. In the short run, the external drivers of the current account, like the competitiveness effect and the oil price, are more significant as explanatory variables, while in the long run, the fiscal balance, demographics and the level of financial market development play a more relevant role in explaining the current account balance along with other determinants.

Our framework provides a comprehensive and nuanced understanding of the intricate interplay between long-run trends and short-run fluctuations, enriching our comprehension of the complex factors influencing external imbalances.

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