An analysis of the trade balance for OECD countries using periodic integration and cointegration

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Abstract We analyze imbalances in external accounts that have historically affected most developed countries. The purpose of this study was to shed some light on the sustainability of the current account for a group of OECD countries by merging the popular Husted (Rev Econ Stat 74(1):159–166, 1992) testing procedure with recent econometric analysis dealing with seasonality. A necessary condition for current account sustainability is that exports and imports are cointegrated. Following previous empirical studies (Husted 1992; Arize in Int Rev Econ Financ 11:101–115, 2002; Hamori in Appl Econ Lett 16:1691–1694, 2009), we analyze the long-run relationship linking exports and imports, using quarterly data. In contrast to these studies, we explicitly deal with seasonal effects through the use of periodic integration and cointegration and find a long-run relationship for the majority of the countries.

Keywords Current account · Time series · Periodic integration · Periodic cointegration

JEL Classification F14 · F32 · C22
1 Motivation

Since the beginning of the 1970s, external imbalances have been widening considerably in the world economy. Economic globalization has meant an increase in international trade and capital mobility facilitating the financing of larger and more persistent disequilibria. Among the OECD countries, there is a clear trend toward larger external imbalances. The relevance of the disequilibrium in the external balances during the last decade has renewed the academic interest for this issue. Gourinchas and Rey (2007) have decomposed the external adjustment into a financial (valuation) channel and a trade (net export) channel and show that the deterioration in net exports or net foreign asset position of a country has to be matched either by future net export growth (trade adjustment channel) or by future increases in the returns of net foreign asset portfolio (financial adjustment channel). The valuation channel is important in the medium-term, whereas the net export channel matters in a long-time horizon. This paper analyzes imbalances in external accounts in the long-run from the traditional trade balance approach that postulates the trade channel as the main external adjustment mechanism. For this reason, the variables of interest are exports and imports of the countries analyzed.

Some previous empirical studies, such as Husted (1992), Arize (2002) and, more recently, Hamori (2009) have dealt with the long-run relationship between exports and imports using the cointegration methodology. Although the majority of the empirical evidence is based on annual or quarterly data and the latter can be affected by seasonal effects, to the best of our knowledge, the empirical literature has neglected the presence of seasonal nonstationary components. However, since the seminal work of Gupta (1965) the importance of seasonality in exports and imports time series is a well-documented fact.

The existence of seasonal products poses some significant challenges for the scholars. According to IMF (2004), seasonal commodities are products that are either not available in the marketplace during certain seasons of the year or available throughout the year but there are regular fluctuations in prices or quantities that are synchronized with the season of the time of the year. A commodity that satisfies the first condition is termed a “strongly seasonal” commodity, whereas a commodity that satisfies the second one will be called a “weakly seasonal.” These terms coincide with the distinction between “narrow” and “wide” seasonal products made earlier by Balk (1980) or by the taxonomic classification between “type 1” and “type 2” seasonality proposed in Diewert (1998).

According to Mitchell (1927), there are two main sources of seasonal fluctuations in prices and quantities that may cause fluctuations in the demand or supply for many products and consequently in trade flows: climate and custom. The importance of seasonality has been assessed by Alterman et al. (1999) quantifying that for a typical country, seasonal purchases will often amount to one-fifth to one-third of all consumer purchases.

Nevertheless, seasonality is a phenomenon that has not received sufficient attention in the economic literature in general. The standard treatment is either to assume that the seasonality that appears in the time series is deterministic or, alternatively, to use
a method to remove the seasonal component of the variables and estimate the models using seasonally adjusted variables.

Ghysels (1990), Ghysels and Perron (1993) and del Barrio Castro et al. (2002) show that the removal of seasonality with X-11 and SEATS standard procedures introduces excessive persistence in the series, which reduces the power of unit root tests. Maravall (1993) shows how seasonal adjustment procedures induce noninvertible moving average processes in the filtered series, invalidating the inference made in most of the unit root and cointegration tests.

Researchers confronted with nonstationary seasonal time series have two alternatives methods to deal with nonstationary seasonality: seasonal integration (SI) (Hylleberg et al. 1990; Hylleberg 1995; Rodrigues and Taylor 2007; Kunst 2009)\(^1\) and/or periodic integration (PI). As argued by Osborn (1988) and Hansen and Sargent (1993),\(^2\) periodic integration is more attractive than SI because PI can arise naturally from the application of economic theory when the underlying economic driving forces, such as preferences or technologies, vary seasonally. Secondly, according to Osborn (1991) and Franses (1994), from an econometric perspective, PI is attractive because it implies that the seasons of the year are cointegrated with each other and hence ensures that the patterns associated with the various seasons are linked in the long-run.

Based on the previous argument, in this paper we intend to use a seasonal treatment that includes periodic autoregressive models as well as PI tests (Boswijk and Franses 1996; del Barrio Castro and Osborn 2011) to determine the type of seasonality present in the nonstationary series analyzed. As shown in Ghysels and Osborn (2001), this point is crucial as it determines the type of cointegration between the set of variables analyzed. Specifically, if the series are seasonally integrated, long-term relationships can occur at each frequency, that is, “seasonal cointegration”, (Lee 1992; Johansen and Schaumburg 1998) or between the seasons of the time series, namely “periodic cointegration” (Boswijk and Franses 1995). However, if the series are periodically integrated, they can only be periodically cointegrated (del Barrio Castro and Osborn 2008). Moreover, if one does not take into account all the above-mentioned possibilities and ignores the univariate properties of the series analyzed, it may originate problems of spurious correlations and unstable parametrization.

Therefore, in this paper, the econometric analysis consists of first determining the order of integration of the trade flows and then, if nonstationary, to test and estimate the existence of a long-run relationship between a country’s exports and imports. The countries in our sample are Australia, Canada, Denmark, Sweden, the UK, Norway, Switzerland and Japan. Those are the OECD countries with nonseasonally adjusted quarterly data that were available. We have excluded France, Italy, the Netherlands, Finland and Spain, the five EMU members in the group, due to the presence of a structural change around 1999.\(^3\)

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1 See also Kunst (1997) and Osborn et al. (1988) for alternative methods of testing for seasonal unit roots.

2 In Gersovitz and McKinnon (1978) it is also possible to find arguments in favor of the use Periodic Autoregressive models.

3 The presence of a level shift recommends the use of unit root tests that explicitly allow for it in the alternative hypothesis. This will be the object of future research.
The rest of the paper is organized as follows. Section 2 briefly presents the theoretical background, while Sect. 3 reports the econometric tests and the empirical results. Section 4 contains the conclusions.

2 Theoretical model

In this paper, we follow Husted (1992) who presents a simple theoretical model of a small open economy with no government where there is a representative consumer. This economy produces and exports a composite good. The consumer can borrow and lend in the international markets using one-period instruments. His resources are output and profits from firms that are used for consumption and savings. The consumer’s budget constraint in the current period is:

\[ C_0 = Y_0 + B_0 - I_0 - (1 + r_0)B_{-1} \]  

(1)

where \( C_0 \) is current consumption, \( Y_0 \) is output, \( I_0 \) is investment, \( r_0 \) is the one-period world interest rate, \( B_0 \) is international borrowing that can be positive or negative, whereas \( (1 + r_0)B_{-1} \) is the stock of debt by the agent (or the country’s external debt). The budget constraint must hold for every period, and the usual transversality condition should be fulfilled. As expression (1) should hold for every period, this constraint can be combined to formulate an intertemporal budget constraint, such as:

\[ B_0 = \sum_{t=1}^{\infty} \mu_t T_A_t + \lim_{n \to \infty} \mu_t B_n \]  

(2)

where \( T_A_t = X_t - M_t (= Y_t - C_t - I_t) \) is the trade balance in period \( t \), that is, income minus absorption; \( X_t \) are the exports and \( M_t \) are imports, whereas \( \mu_t \) is the discount factor.

Husted (1992) arrives to a testable equation that relates exports with imports inclusive of interest payments on net debt:

\[ X_t = a + b^* M_Mt + e_t \]  

(3)

where \( M_Mt = M_t + r_t B_{t-1} \). Under the null hypothesis that the economy satisfies its intertemporal budget constraint, we expect \( b = 1 \) and \( e_t \) is stationary. Thus, if both variables are I(1), under the null, they are cointegrated, with a cointegrating vector \((1, -1)\).

We assume that the world interest rate is stationary and that the transversality condition holds in Eq. (2). Therefore, the accumulated discounted debt is stationary, and then, the term \( r_t B_{t-1} \) would also be stationary. In practice, we can test for cointegration between exports and imports when we believe that the adjustment works essentially through the trade channel.
3 Econometric techniques

In order to explicitly acknowledge the role of seasonality, it is often convenient to represent a univariate time series as $y_{s\tau}$, where the first subscript refers to the season ($s$) and the second subscript to the year ($\tau$), as we have quarterly data $s = 1, 2, 3, 4$. For simplicity of exposition, we assume that data are available for precisely $N$ years, so that the total sample size is $T = 4N$. Note that, throughout the paper, it is understood that $y_{s-k,\tau} = y_{4-s+k,\tau-1}$ for $s-k \leq 0$.

Applications of periodic processes within economics have focused on the autoregressive case, with the $p$th order periodic autoregressive, or PAR($p$) process, defined by

$$y_{s\tau} = \alpha_s + \phi_{1s}y_{s-1,\tau} + \phi_{2s}y_{s-2,\tau} + \cdots + \phi_{ps}y_{s-p,\tau} + e_{s\tau}, \quad s = 1, 2, 3, 4$$

where $e_{s\tau}$ is white noise. In (4), we only consider seasonal intercepts $\alpha_s$ due to the nature of the analyzed data that are ratios. Note that all the coefficients in this process may vary over seasons $s = 1, \ldots, 4$. The conventional (nonperiodic) AR($p$) process is a special case with $\phi_{is} = \phi_i$ ($s = 1, 2, 3, 4$) for all $i = 1, 2, \ldots, p$. However, in the presence of seasonality, it is important to consider the possibility that the process may be periodic, with at least some AR coefficients in (4) varying over the year.

Under the assumption that $y_{s\tau}$ is integrated of order 1, and using a similar notation to Boswijk and Franses (1996), (4) can also be written as

$$\left(y_{s\tau} - \varphi_0 y_{s-1,\tau}\right) = \alpha_s^* + \sum_{j=1}^{p-1} \psi_{js} \left(y_{s-j,\tau} - \varphi_{s-j} y_{s-j-1,\tau}\right) + e_{s\tau}$$

where $\prod_{s=1}^4 \varphi_s = 1$ with the quasi-difference $y_{s\tau} - \varphi_0 y_{s-1,\tau}$ being stationary. Note also that it is understood that $\varphi_{s-j} = \varphi_{4+s-j}$ for $s-j \leq 0$ and also for $\alpha_s^*$ in (13). Boswijk and Franses (1996) analyze the distribution of the likelihood ratio test statistic for the null of PI $\prod_{s=1}^4 \varphi_s = 1$ in (5), with this statistic defined by

$$LR_{PI} = T \ln \left( \frac{RSS_0}{RSS_1} \right)$$

where RSS$_0$ and RSS$_1$ denote the residual sum of squares under the null hypothesis and from the unrestricted form (4), respectively. Under the null hypothesis of a PI(1) process, they show that this statistic has the same asymptotic distribution as the squared Dickey–Fuller $t$ statistic for a conventional (nonperiodic) I(1) process.

To implement the previous test (6), we need to determine the order $p$ for the unrestricted and restricted models (4) and (5). To do that, we follow Franses and Paap (1994), Franses and Paap (2004) and use the Schwarz criterion in conjunction with diagnostic tests for neglected periodic serial correlation to determine $p$ with a maximum value of 5. Boswijk and Franses (1996) also proposed a F-type statistic $F_{per}$

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4 See Ghysels and Osborn (2001, pp. 153–155) for details about the models nested in (4)/(5)
to test the null of nonperiodic variation in the coefficients of (4) \( H_0 : \phi_{js} = \phi_j \) for \( j = 1, \ldots, p \). The implementation of the LR\(_{PI}\) Boswijk and Franses test by practitioners has two problems: in first place, the models (4) and (5) tend to have a large number of parameters, and in second place, to fit model (5) we will need nonlinear methods of estimation. Recently, del Barrio Castro and Osborn (2011) have proposed two nonparametric tests (based on the Breitung 2002 and Stock 1999 unit root tests) that allow us to circumvent the limitations of the Boswijk and Franses (1996) test. They propose to compute a variance ratio statistic for a given season \( s \) as

\[
VRT_s = N^{-2} \sum_{\tau=1}^{N} \frac{\hat{U}_{s\tau}^2}{\sum_{\tau=1}^{N} \hat{u}_{s\tau}^2} \quad s = 1, \ldots, 4
\]

where \( \hat{U}_{s\tau} \) is the season-specific partial sum \( \hat{u}_{s1} + \hat{u}_{s2} + \cdots + \hat{u}_{s\tau} \), with \( \hat{u}_{s\tau} \) obtained as the OLS residuals \( \hat{u}_{s\tau} = y_{s\tau} - \hat{\beta}'_s z_{\tau} \) from a regression of observations for season \( s \), \( y_{s\tau} (\tau = 1, \ldots, N) \), on \( z_{\tau} \) that collects the deterministic part, in our case \( z_{\tau} = 1 \). In order to test the PI(1)/I(1) null hypothesis, they use the average variance ratio statistic

\[
VRT_{PI} = 4^{-1} \sum_{s=1}^{4} VRT_s
\]

where each \( VRT_s \) is defined in (7).

Additionally, based on Perron and Ng (1996) and Stock (1999), del Barrio Castro and Osborn (2011) propose to apply for a single season \( s \), the corresponding season-specific modified Sargan and Bhargava (1983) test statistic:

\[
MSB_s = \left( N^{-2} \sum_{\tau=1}^{N} \frac{\hat{u}_{s,\tau-1}^2}{\hat{\gamma}_{sl}} \right)^{\frac{1}{2}} \quad s = 1, \ldots, 4
\]

which requires an appropriate long-run variance estimator \( \hat{\gamma}_{sl} \) for the annual difference \( \Delta u_{s\tau} = u_{s\tau} - u_{s,\tau-1} \) relating to season \( s \). \( \hat{\gamma}_{sl} \) is obtained based on sample autocovariances using the Bartlett and quadratic spectral kernels, following (Newey and West (1994), equations (3.8) to (3.15) and Table 1) data-dependent bandwidth procedure.

As in the previous case, they propose the use of the average MSB\(_{PI}\) statistic

\[
MSB_{PI} = 4^{-1} \sum_{s=1}^{4} MSB_s.
\]

del Barrio Castro and Osborn (2011) show that the VRT\(_{PI}\) (8) and MSB\(_{PI}\) (10) tests under the null of PI \( \prod_{s=1}^{4} \phi_s = 1 \) (8) have the same distribution of the variance ratio test proposed by Breitung (2002) in the case of VRT\(_{PI}\) (8) and of the modified Sargan–Bhargava test proposed by Stock (1999) in the case MSB\(_{PI}\) (10).

As mentioned in Sect. 1 above, Osborn (1991) and Franses (1994) show that the main characteristic of a periodically integrated process is that the nonstationary behavior is caused by a common stochastic trend shared by the quarters of the time series. Based on that Franses (1994) and del Barrio Castro and Osborn (2012) proposed the
Table 1  Test for the null of not periodicity in (4) and of PI

| Country   | $F_{\text{per}}$ | $\text{LR}_{\text{PI}}$ | PI order | $\text{MSB}^{b}_{\text{PI}}$ | $\text{MSB}^{c}_{\text{PI}}$ | VRT$_{\text{PI}}$ |
|-----------|------------------|--------------------------|----------|-------------------------------|-------------------------------|------------------|
| Australia | 6.9374**         | 12.6438**                | 2        | 0.2798                        | 0.2825                        | 0.0600           |
|           | 11.2127**        | 7.1483                   | 1        | 0.3329                        | 0.3462                        | 0.0618           |
|           | 1.8039           | 18.7643**                | 3        | 0.1817**                      | 0.1944*                       | 0.0098**         |
| Canada    | 7.0393**         | 3.6664                   | 2        | 0.4465                        | 0.4474                         | 0.0797           |
|           | 7.3068**         | 3.3667                   | 2        | 0.4291                        | 0.4283                         | 0.0814           |
|           | 2.6053**         | 13.0961**                | 2        | 0.1826**                      | 0.1792**                       | 0.0095**         |
| Denmark   | 3.4351**         | 6.9169                   | 2        | 0.3046                        | 0.3077                         | 0.0624           |
|           | 3.4995**         | 13.5095**                | 1        | 0.1981                        | 0.2115                         | 0.0149           |
|           | 3.1794**         | 8.1569                   | 2        | 0.4362                        | 0.4385                         | 0.0637           |
| Sweden    | 3.9594**         | 3.5997                   | 2        | 0.2732                        | 0.2675                         | 0.0348           |
|           | 3.9580**         | 6.2928                   | 1        | 0.2523                        | 0.2501                         | 0.0331           |
|           | 5.2382**         | 6.7620                   | 3        | 0.3021                        | 0.3313                         | 0.0217           |
| United Kingdom | 0.3915   | 5.8637                   | 2        | 0.2752                        | 0.2635                         | 0.0274           |
|           | 2.7003**         | 8.6335*                  | 1        | 0.1964*                       | 0.2038*                        | 0.0351           |
|           | 3.7082**         | 15.3947**                | 1        | 0.1555*                       | 0.1591**                       | 0.0115*          |
| Norway    | 1.5214           | 6.4674                   | 2        | 0.1928*                       | 0.1944*                        | 0.0382           |
|           | 3.0904**         | 4.7313                   | 2        | 0.2746                        | 0.2721                         | 0.0655           |
|           | 2.4200**         | 3.3746                   | 3        | 0.2762                        | 0.2785                         | 0.0670           |
| Switzerland | 4.6170**        | 1.1163                   | 2        | 0.2329                        | 0.2342                         | 0.0111*          |
|           | 0.7261           | 4.0753                   | 2        | 0.2037*                       | 0.2031*                        | 0.0085**         |
|           | 6.2268**         | 1.9552                   | 2        | 0.2381                        | 0.2330                         | 0.0124*          |
| Japan     | 2.8044**         | 2.9987                   | 1        | 0.3039                        | 0.3022                         | 0.0248           |
|           | 0.2432           | 3.2508                   | 1        | 0.3208                        | 0.3159                         | 0.0255           |
|           | 1.0003           | 8.1958*                  | 1        | 0.2420                        | 0.2418                         | 0.0173           |

***, * Statistically significant at 5 and 10 %, respectively

use of the Johansen (1988) and the Breitung (2002) procedures, respectively, to determine the number of cointegration relationship (common trends) between the quarters (seasons) of the time series. The results of a small Monte Carlo experiment (avail-

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5 We have not applied these procedures to our data because these methods perform well in this context for sample sizes of around 75 years in the case of the Johansen tests and 100 years in the case of the Breitung tests.
able upon request) demonstrate that the Johansen procedure when applied to quarterly PI processes with sample between 30 and 50 years (as in our case) is able to determine the presence of cointegration relationship between the quarters but is unable to determine whether there are 3 cointegration relationship between the quarters. This same situation was reported by del Barrio Castro and Osborn (2012) for the Breitung cointegration tests.

As shown by Ghysels and Osborn (2001, pp. 168–171) and del Barrio Castro and Osborn (2008), when the series follow PI processes, the only cointegration possibilities are periodic cointegration or nonperiodic cointegration, with cointegration for any one season implying cointegration for all seasons, that is, full cointegration. They also show that in order to have full nonperiodic cointegration (equivalent here to conventional cointegration), the involved processes must share the same $\phi_s$ coefficients in (5). Hence, if two variables are cointegrated with a $(1, -1)$ vector, both processes must share the same $\varphi_s$ coefficients in (5). Thus, the LR$_{PI}$, MSB$_{PI}$ and VRT$_{PI}$ tests can be applied to the difference between two variables.

Finally, del Barrio Castro and Osborn (2008) propose a residual-based likelihood ratio test (LRCR) for the null of not full periodic cointegration between periodically integrated processes and obtain their asymptotic distribution; in particular, they show that the LRCR statistics follow the squared distribution reported by Phillips and Ouliaris (1988) for the residual-based ADF cointegration test. This test is based on testing the null of PI in the residuals of the following model (in the case of two variables $y_{s\tau}$ and $x_{s\tau}$):

$$
y_{s\tau} = \alpha_s + \beta_s x_{s\tau} + u_{s\tau} \quad s = 1, 2, 3, 4.
$$

(11)

Under the null of non-full periodic cointegration, del Barrio Castro and Osborn (2008) show (see Lemma 4) that the residuals $\hat{u}_{s\tau}$ of (11) asymptotically retain the same nonstationary periodic coefficients of the univariate process for $y_{s\tau}$. Hence, the strategy is to test for PI in the residuals $\hat{u}_{s\tau}$ using the unrestricted model:

$$
\hat{u}_{s\tau} = \sum_{j=1}^{p^*} \phi^*_{js}\hat{u}_{s-j,\tau} + \epsilon_{s\tau} \quad s = 1, 2, 3, 4 
$$

(12)

with residual sum of squares $\text{RSS}_1$. And the restricted model:

$$
\hat{u}_{s\tau} = \varphi^*_{s}\hat{u}_{s-1,\tau} + \sum_{j=1}^{p^*-1} \psi^*_{js} \left(\hat{u}_{s-j,\tau} - \varphi^*_{s-j}\hat{u}_{s-j-1,\tau}\right) + \epsilon_{s\tau}
$$

(13)

subject to $\varphi^*_{1}\varphi^*_{2}\varphi^*_{3}\varphi^*_{4} = 1$, with residual sum of squares $\text{RSS}_0$. Finally, the LRCR to the test the null of non-full periodic cointegration has the expression:

$$
\text{LRCR} = T \ln \left(\frac{\text{RSS}_0}{\text{RSS}_1}\right).
$$

(14)
4 Empirical results

As in Arize (2002), we analyze the natural logarithms of the nominal ratio exports to GDP (ln(exp/gdp) hereafter) and imports to GDP (ln(imp/gdp)). We have collected quarterly data (not seasonally adjusted) for the following non-Eurozone countries: Australia, Canada, Denmark, Sweden, United Kingdom, Norway, Switzerland and Japan. The evolution of the ratios is depicted in Figs. 1 and 2. The sample ends in 2009Q1 for all the countries considered, but it has different starting dates: 1960Q1 for Australia, 1961Q1 for the UK, 1977Q1 for Canada, 1978Q1 for Denmark and finally 1980Q1 for the remaining countries.

From the graphs, we can observe that the ratios ln(exp/gdp) and ln(imp/gdp) show clear seasonal variation but not very large seasonal oscillations. Note also that from the evolution of the time series, we do not observe a trending behavior in our data (with the exception of some weak evidence in the cases of Canada and Sweden). Hence, we consider only seasonal dummies in the deterministic part.

Taking into account the previous arguments and the evolution of the ratios ln(exp/gdp) and ln(imp/gdp) for each country (Figs. 1, 2), we focus on PI as the potential source of nonstationarity in our data.

The results obtained for the unit root tests described in the previous section are reported in Table 1. The first column corresponds to the $F_{per}$ test, the second to Boswijk and Franses (1996) LR$_{PI}$ or likelihood ratio test, followed by the order of the fitted
PAR ($p$) for each time series. MSB$^p_{\text{PI}}$ and MSB$^q_{\text{PI}}$ denote the statistic MSB$_{\text{PI}}$ with the Bartlett and quadratic spectral kernels, respectively. Finally, we present the results of VRT$_{\text{PI}}$, that is, del Barrio Castro and Osborn (2011) variance ratio test. For each country, we also report the results for the difference between ln(exp/gdp) and ln(imp/gdp).

From the results of the $F_{\text{per}}$ test, we find clear evidence of periodicity in both ln(exp/gdp) and ln(imp/gdp) for the majority of the countries. Exceptions are the cases of Norway for ln(exp/gdp) and Switzerland and Japan for ln(imp/gdp). The tests LR$_{\text{PI}}$, MSB$_{\text{PI}}$ and VRT$_{\text{PI}}$ do not reject the null of PI for Canada, Sweden and Japan. In the case of Australia, the tests MSB$_{\text{PI}}$ and VRT$_{\text{PI}}$ do not reject the existence of PI. Concerning the LR$_{\text{PI}}$ test, it does not reject the null of PI for ln(imp/gdp) but it rejects the null for ln(exp/gdp). For Denmark, only the MSB$_{\text{PI}}$ tests reject the null of PI for ln(imp/gdp). In the UK, the null is rejected for ln(imp/gdp) at the 10% with the LR$_{\text{PI}}$ and the MSB$_{\text{PI}}$ tests. For Norwegian variables, only the MSB$_{\text{PI}}$ tests reject the null at 10% for ln(exp/gdp). Finally, in the case of Switzerland, the VRT$_{\text{PI}}$ test for ln(exp/gdp) rejects the null at 10% level of significance and for ln(exp/gdp) at a 5% level with the VRT$_{\text{PI}}$ test and at a 10% level for the MSB$_{\text{PI}}$ tests. Overall, we can conclude that we have found reasonable empirical evidence in favor of the hypothesis that the two ratios follow periodically integrated processes for all the countries. In order to take into account the weak trending behavior found in the cases of Canada and Sweden for ln(exp/gdp) and ln(imp/gdp), we have computed the PI tests including

Fig. 2 Evolution of ln(exp/gdp) and ln(imp/gdp) for United Kingdom, Norway, Switzerland and Japan
seasonal dummies and trends. The results are that the tests do not reject the null of PI in any of the cases.\(^6\)

As shown by Ghysels and Osborn (2001, pp. 168–171) and del Barrio Castro and Osborn (2008), when the series follow PI processes, the only cointegration possibilities are periodic cointegration or nonperiodic cointegration, with cointegration for any one season implying cointegration for all seasons, that is, full cointegration. They also show that in order to have full nonperiodic cointegration, the involved processes must share the same \(\phi_s\) coefficients in (5). Note that full nonperiodic cointegration is equivalent to conventional cointegration. Hence, if ln(exp/gdp) and ln(imp/gdp) are cointegrated with a \((1, -1)\) vector, both processes must share the same \(\phi_s\) coefficients in (5). We report these coefficients in Table 2, including all the time series for all the countries in our sample. We also report in Table 1 (see the last row for each country) the results

\(^6\) The role played by intercepts and trends in periodically integrated processes is more complicated than in the case of standard integrated processes. For an in-depth analysis of this, see Paap and Franses (1999)

\[\text{Table 2 Coefficients of (5) and results of the LR}_{\text{CR}} \text{ tests}\]

|        | \(\hat{\phi}_1\) | \(\hat{\phi}_2\) | \(\hat{\phi}_3\) | \(\hat{\phi}_4\) | LR\(_{\text{CR}}\) |
|--------|------------------|------------------|------------------|------------------|-----------------|
| **Australia** |                  |                  |                  |                  |                 |
| ln(exp/gdp)  | 0.713            | 1.006            | 1.251            | 1.115            | 33.1368**       |
| ln(imp/gdp)  | 0.758            | 1.062            | 1.098            | 1.132            |                 |
| **Canada**   |                  |                  |                  |                  |                 |
| ln(exp/gdp)  | 1.1376           | 0.8967           | 1.0797           | 0.908            | 29.5145**       |
| ln(imp/gdp)  | 1.0770           | 0.8723           | 1.2030           | 0.8848           |                 |
| **Denmark**  |                  |                  |                  |                  |                 |
| ln(exp/gdp)  | 0.9613           | 0.9996           | 1.3155           | 0.7911           | 7.8410          |
| ln(imp/gdp)  | 1.3120           | 0.8743           | 1.0011           | 0.8708           |                 |
| **Sweden**   |                  |                  |                  |                  |                 |
| ln(exp/gdp)  | 1.0823           | 0.9574           | 1.2000           | 0.8042           | 6.8975          |
| ln(imp/gdp)  | 1.1039           | 0.9999           | 1.1213           | 0.8036           |                 |
| **UK**       |                  |                  |                  |                  |                 |
| ln(exp/gdp)  | 0.9578           | 1.0768           | 1.0082           | 0.9617           | 15.4674**       |
| ln(imp/gdp)  | 0.9077           | 1.1893           | 0.9651           | 0.9598           |                 |
| **Norway**   |                  |                  |                  |                  |                 |
| ln(exp/gdp)  | 1.0272           | 0.8724           | 1.3097           | 0.8520           | 9.5170*         |
| ln(imp/gdp)  | 0.9821           | 0.9804           | 0.7535           | 1.3784           |                 |
| **Switzerland** |                |                  |                  |                  |                 |
| ln(exp/gdp)  | 1.2062           | 1.0615           | 1.1387           | 0.8190           | 1.7272          |
| ln(imp/gdp)  | 1.1041           | 1.0119           | 1.0928           | 0.6859           |                 |
| **Japan**    |                  |                  |                  |                  |                 |
| ln(exp/gdp)  | 0.9054           | 1.1573           | 0.9961           | 0.9582           | 12.1204**       |
| ln(imp/gdp)  | 0.9786           | 1.0267           | 1.0125           | 0.9831           |                 |

\* Statistically significant at 5 and 10 \%, respectively
obtained for the LR_{PI}, MSB_{PI} and VRT_{PI} when applied to the difference between ln(exp/gdp) and ln(imp/gdp), denoted as diff\_ln. The results about the residual-based likelihood ratio test proposed by del Barrio Castro and Osborn (2008) are also reported in Table 2.

In the case of Australia, Canada and the UK, we find clear evidence of cointegration with a (1, −1) vector. Note that the coefficients for ln(exp/gdp) and ln(imp/gdp) in Table 2 are quite similar. Moreover, applying the LR_{CR} test, we also find evidence of full periodic cointegration as expected. In the case of Norway, there is no (1, −1) cointegration but we detect full periodic cointegration at a 10% level. Also note that in this case, the \( \varphi_s \) coefficients are quite different. For Japan, we find weak evidence of (1, −1) cointegration (at 10% with the LR_{PI} test), but strong full periodic cointegration with the LR_{CR} test. Finally for Denmark, Sweden and Switzerland, we do not find nonperiodic cointegration with vector (1, −1) nor full periodic cointegration.7

5 Concluding remarks

The issue of external imbalances has regained interest in the last years due to the impact of the international financial crisis. In this paper, we assess whether the trade channel, as an adjustment mechanism in the long-run, has been working as postulated by the theory. The contribution of the paper to previous literature is twofold. First, from an econometric point of view, we refine previous analysis considering the seasonal components of the variables involved. This is not trivial because seasonality is important in exports and imports and should be addressed to guarantee a robust empirical analysis. Second, we obtain more evidence in favor of cointegration than in previous studies, which has important consequences from an economic policy view. In a well-functioning economy, deficits are temporary phenomena that will be balanced by future surpluses but in a country with distorted markets, there is no tendency toward balance of payments equilibrium, and thus, sustained external imbalances may reflect the implementation of a bad economic policy or just “bubble-financing”.

Therefore, we analyze the long-run relationship linking exports and imports for a group of developed countries between 1960/1970 and 2009 using quarterly data non-adjusted for seasonality. As many official statistical offices only provide information of seasonally adjusted data, we are restricted to twelve countries and different data spans: Australia, Canada, Denmark, France, Finland, Italy, Japan, the Netherlands, Norway, Spain, Switzerland, Sweden and UK. We have finally excluded the five EMU members due to the presence of a clear level shift in 1999.

In the empirical literature, the issue of seasonal nonstationary components has been frequently neglected. Thus, the aim of the paper was to apply a flexible approach to cointegration, where we allow for the presence of these type of nonstationary components (instead of deterministic seasonality).

7 Following the suggestion of one referee, we have conducted the cointegration analysis for the annual data using the Johansen procedure and Engle and Granger two steps approach. The results are available upon request from the authors. As expected, it is possible to say the aggregation process do not alter the long-run relationships between the analyzed variables.
The general conclusion of the univariate analysis is that both export and import ratios over GDP are periodically integrated processes for all the countries. Consequently, the only cointegration possibilities are periodic cointegration or nonperiodic cointegration, with cointegration for any one season implying cointegration for all seasons (i.e., full cointegration). Using a residual-based \( \text{LR}_{CR} \) test proposed by del Barrio Castro and Osborn (2008) for the null hypothesis of non-full periodic cointegration, we conclude that we find either full nonperiodic cointegration, such as in the cases of Canada, the UK and Australia, or full periodic cointegration, as in Norway and Japan. However, neither nonperiodic nor full periodic cointegration is found for Denmark, Sweden and Switzerland signaling the existence of persistent disequilibria and the need of prospective adjustments. These three countries are European, two of them EU members. This last result may be the consequence of the close commercial link between them and the Eurozone which has been affected by a progressive increase of financial integration among EU members (singularly important for euroarea members). The consequence has been an improvement in the financial conditions of EU countries and a widening of external disequilibria among them (i.e., persistent surpluses in the case of Denmark and Sweden) that has only been corrected from 2007. Moreover, the case of Switzerland presents peculiarities related to the importance of the financial sector in the economy and its capacity for permanent financing of its external imbalances.

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