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Research Article

Do Agricultural Imports and Exports Cointegrate? Evidence from 13 OECD Countries

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Abstract. Previous studies have investigated the behaviour of trade flows at the aggregate level, thus they suffer from aggregation bias. In this paper, we use the sectoral data on agricultural exports and imports to examine whether they cointegrate. The likelihood-based panel cointegration technique is applied to investigate the long-run convergence between the variables for 13 industrialized countries. The results indicate that a long-run steady-state relationship exists between the variables for most countries in the sample. The policy implications of our findings are that agricultural trade does not lead to the violation of international budget constraints and, more importantly, there is no productivity gap in the agriculture sector between the domestic economy and the rest of the world, implying a lack of permanent technological shocks to the domestic economy. The results also provide support for intra-industry trade in the agriculture sector.

Keywords: agriculture, imports, exports, cointegration.
JEL codes: E60, F31, F14.

1. INTRODUCTION

A major indicator of a country's economic performance is the external account because significant external imbalances might predict future changes in a managed foreign exchange system. Empirical studies attempt to identify the sources of external imbalances by relating the external accounts to key macroeconomic variables such as government spending, private consumption, income, the net financial balance of the household sector, non-financial and financial corporations, etc. (Sachs, 1981; Ahmed, 1987; Razin, 1995; Elliott, Fatas, 1996; Chen *et al*., 2013; Allen, 2019).

Some authors argue that fiscal, monetary, and commercial policies (tariff, subsidy and exchange-rate policies) have aimed to reduce the size of external imbalances in several countries (e.g., Artis, Bayoumi, 1989; Ariza, Bahmani-Oskooee, 2018). In most cases, fiscal and monetary policies are used to alleviate domestic problems such as recession or inflation rather than external accounts problems. On the other hand, commercial policies such as currency devaluations or depreciations are used to deal with external problems such as reducing trade deficits.

It is not easy to separate the effectiveness of one policy in solving a problem over other policies. As far as the external accounts is concerned, one way to examine the effectiveness of all policies is to determine whether or not a country's exports and imports cointegrate in the long run (Husted, 1992). If they do, then we can believe that the combined effects of all macro policies are effective. Other studies state that there is evidence of external imbalance being the outcome of "bad policy" (Summers, 1988; Husted, 1992; Irandoust, Sjöö, 2000; Irandoust, Ericsson, 2004). They conclude that outflows and inflows in the current account cointegrate unless there are policy distortions or permanent productivity shocks to the domestic economy. Thus, in a well-functioning economy, external accounts deficits are temporary phenomena that will be balanced by future surpluses. In a country with distorted markets there is no tendency towards a balance of payments equilibrium and thus sustained external imbalances reflect "bad policy".

An external imbalance is regarded as sustainable when it does not violate the nation's solvency constraint; and a nation is said to be solvent if the present-value budget constraint, i.e., its intertemporal budget constraint holds. One way to analyze external imbalances applies the intertemporal approach to the current account (Sachs, 1981; Obstfeld, Rogoff, 1995; Razin, 1995; Irandoust, Sjöö, 2000; Raybaudi *et al.*, 2004; Chen, 2011, 2014; Afonso *et al.*, 2020). According to this approach, the current account equals the difference between savings and investment, and, because savings and investment decisions are based on intertemporal factors (such as life-cycle features, the expected returns of investment projects, and the like) the current account is necessarily an intertemporal phenomenon. Thus, a trade balance or current account balance would be sustainable if the series for exports and imports are found to be cointegrated (Trehan, Walsh, 1991; Hakkio, Rush, 1991; Wickens, Uctum, 1993; Wu *et al.*, 1996; Apergis *et al.*, 2000; Irandoust, Sjöö, 2000; Afonso *et al.*, 2020).

In this study, we focus on the agriculture sector and cointegration between agricultural imports and exports since anticipated growth in the demand for food and agricultural raw materials due to increasing world population and incomes will place significant demands upon the scarce natural resources, particularly land and water, used in the sector. Although agriculture is not expected to add significantly to job creation in the OECD countries due to the relatively small contribution that the sector makes to total employment, its use of purchased inputs and the supply of food and raw materials to other sectors are significant for employment and total economic activity. Thus, our research questions are: Are agricultural exports and agricultural imports cointegrated? What are the implications of agricultural imports and exports being cointegrated?

The cointegration between agricultural exports and imports also indicates intra-industry trade (IIT) within the agriculture sector. The creation and expansion of the European Union has contributed to an increase in IIT between European countries. Despite the importance of the topic, most literature examines IIT of industrial products and the agricultural sector is usually neglected in empirical works, possibly because agricultural markets are assumed to be competitive. However, recent studies support the view that agricultural markets can be characterized by imperfect competition and economies of scale (Sexton, 2013) and IIT plays an increasing role in agricultural trade (e.g., Leitao, 2011; Ferto, 2015a, b).

Examples of studies that have found evidence of cointegration between aggregate exports and aggregate imports include Bahmani-Oskooee (1994), who tested the hypothesis for Australia, Bahmani-Oskooee and Rhee (1997), for Korea, Arize and Bahmani Oskooee (2018), for 100 countries that supported nonlinear cointegration in most cases of bilateral trade, Irandoust and Ericsson (2004), for industrial countries. Previous studies suffer from aggregation bias since they use aggregate exports and imports to investigate trade flows and external accounts position. Thus, the purpose of this paper is to examine the behaviour of the agricultural trade flows in 13 OECD countries (France, Germany, Italy, Sweden, Switzerland, Austria, Denmark, Norway, Spain, Portugal, Finland, the Netherlands and the UK). We focus on these countries because of the fact that they are major exporters and importers of agricultural products in Western Europe. The total value of imports and exports of agricultural products between the EU and the rest of the world was EUR 275 billion in 2017 (Eurostat, 2018).

The departures from earlier studies are in disaggregate agricultural trade flows and the asymptotic theory of likelihood-based panel cointegration allowing for multiple cointegrating vectors. The main contribution of this study stems from the methodology used which is a likelihood-based panel cointegration under assumptions of cross-sectional dependence and slope homogeneity restrictions. This is an extension of the Johansen (1995) multivariate maximum likelihood developed by Larsson and Lyhagen (1999) and Larsson *et al.* (2001). They developed a likelihood-based panel test of the cointegrating rank and a general likelihood-based framework for inference in panel-VAR models with cointegration restriction, allowing for multiple cointegrating vectors. By using this method, the assumption of a unique cointegrating vector and the problem of normalization is relaxed. This is not the case with the usual residual-based tests of cointegration (e.g., Kao, 1999; Pedroni, 1999a, b). However, to the best of the author's knowledge, this study is the first attempt to test the cointegration between agricultural exports and imports using panel cointegration techniques based on likelihood inference of cointegrating vectors.

Our results indicate that agricultural trade flows are cointegrated for all countries in the sample except for Italy, Sweden and the Netherlands. The cointegration between agricultural exports and imports reveals that these countries are not in violation of their international budget constraints. Furthermore, macroeconomic policies have been effective in bringing agricultural imports and exports into equilibrium in the long run. More importantly, there is no productivity gap between the domestic economy and the rest of the world, implying a lack of permanent technological shocks to the domestic economy. The findings also provide support for ITT in most of the countries under review.

The paper is organized as follows. Section 2 outlines a simple model and we discuss the data and methodology used. In section 3, we present and interpret the results from the cointegration tests. In section 4 we discuss some policy implications. Conclusions are given in section 5.

2. MODEL, DATA AND METHODOLOGY

The international budget constraint for analyzing the dynamics of the exports and imports follows Husted (1992), Irandoust and Sjöö (2000) and Irandoust and Ericsson (2004). These studies show that the international budget constraint for a given country can be written as

$$EXP_t = a+bIMP_t + e_t \tag{1}$$

where EXP_t and IMP_t denote agricultural exports and imports respectively. The null hypothesis states that the economy satisfies its international budget constraint. Thus, it is expected that b = 1, and e_t is a stationary process that includes all short-term dynamics. In other words, if EXP_t and IMP_t are nonstationary and trending, then under the null hypothesis they are co-trending (cointegrating) with cointegrating vector b = (1, -1).

An important question here pertains to the policy implications of cointegration or lack of cointegration and convergence between agricultural imports and exports. The theory suggests that cointegration is to be expected under the maintained hypothesis that the economy is working properly and that breaking international budget constraints leads to a lack of cointegration.

An important reason why the time series paths of agricultural imports and exports might diverge, and not cointegrate, is technological shocks or the productivity gap hypothesis. Thus, finding cointegration for the variables rejects the assumption of a permanent technological or productivity gap between the economy and the rest of the world (Irandoust, Sjöö, 2000; Irandoust, Erisson, 2004). In other words, if agricultural trade flows are not cointegrating, this could be regarded as the outcome of permanent technological shocks to the domestic economy.

The data used in this study are agricultural (raw materials) exports and imports as a percentage of merchandise imports and exports, respectively. The sample consists of 13 European industrialized countries (France, Germany, Italy, Sweden, Switzerland, Austria, Denmark, Norway, Spain, Portugal, Finland, the Netherlands and the UK) and covers the period 1963-2020. The choice of the time period and sample countries are dictated by data availability. The variables are extracted from the World Bank database. Figures 1-13, Appendix A, illustrate the variables. Descriptive statistics for the variables under analysis is also reported in Table A.1., Appendix A.

The process is estimated by implementing a likelihood-based panel framework developed by Larsson and Lyhagen (1999) and Larsson *et al.* (2001). By using this method, the assumption of a unique cointegrating vector and the problem of normalization is relaxed which is not the case with the usual residual-based tests of the cointegration approach. Let LR denote the cross-section-specific likelihood-ratio (trace) statistic of the hypothesis that there are at most r cointegrating vectors in the system. The standardized LR-bar statistic is given by:

$$Y_{\bar{LR}} = \frac{\sqrt{N\left(\bar{LR} - \mu\right)}}{\sqrt{V}} \tag{2}$$

where $\bar{LR}$ is the average of the N cross-section LR statistics, μ is the mean and v is the variance of the asymptotic trace statistic. Asymptotic values of μ and v (with and without constant and trend) can be obtained from stochastic simulations as described in Johansen (1995).[1]

Two steps should be followed before using any cointegration tests: testing the panel for cross-sectional dependence and testing for cross-country heterogeneity. The first issue means the transmission of shocks from

[1] This methodology is also used in Irandoust and Ericsson (2005).

one variable to another. In other words, all countries in the sample are affected by globalization and have common economic characteristics. The second issue shows that a significant economic connection in one country is not necessarily replicated by the others. A set of three tests is constructed to check the cross-sectional dependence assumption: the Breusch and Pagan (1980) cross-sectional dependence (CD_{BP}) test, the Pesaran (2004) cross-sectional dependence (CD_P) test, and the Pesaran *et al.* (2008) bias-adjusted LM test (LM_{adj}). Regarding the country-specific heterogeneity assumption, the slope homogeneity tests ($\tilde{\Delta}$ and $\tilde{\Delta}_{adj}$) of Pesaran and Yamagata (2008) are used (Appendix B provides more information about these tests).

The traditional panel unit root tests do not consider cross-sectional dependence of the contemporaneous error terms. Failing to take into account cross-sectional dependence may lead to misleading results. Thus, to eliminate this problem, we use the cross-sectionally augmented panel unit root test (CIPS) that allows for parameter heterogeneity and serial correlation between the cross-sections (Pesaran, 2007).[2] Finally, we check diagnostic tests, i.e., if the residuals are normally distributed and there is no autocorrelation. The normality test stems from a multivariate extension of the Bowman-Shenton test developed by Doornik and Hansen (1994) and the test for autocorrelation is the Ljung-Box test statistics.

3. ESTIMATION RESULTS

As a pre-test for the cointegration analysis, we first examine cross-sectional dependence and slope homogeneity assumptions. Table 1 indicates the results of cross-sectional dependence tests (CD_{BP}, CD_p, and LM_{adj}) and slope homogeneity tests ($\tilde{\Delta}$ and $\tilde{\Delta}_{adj}$). The first set of tests, for cross-sectional dependence, clearly shows that the null hypothesis of no cross-sectional dependence is rejected for all significance levels. This implies that there is a cross-sectional dependence in the case of our sample countries. Any shock in one country is transmitted to others. The second part of the table shows that the null hypothesis of slope homogeneity is rejected for both tests and all significance levels. This means that the economic relationship in one country is not replicated by the others. As there are both cross-sectional dependence and slope heterogeneity, the cointegration tests can be used.

Tab. 1. Cross-sectional dependence and slope homogeneity tests.

Method	Test statistic
Cross-sectional dependence test	
CD_{BP}	377.126*** (0.000)
CD_P	54.392*** (0.000)
LMadj	65.287*** (0.000)
Slope homogeneity test	
$\tilde{\Delta}$ test	19.205*** (0.000)
$\tilde{\Delta}_{adj}$ test	16.211*** (0.000)

*** indicate significance for 0.01 levels. The numbers within parentheses show *p*-values.
CD_{BP} test, CD_P test and LM_{adj} test show the cross-sectional dependence tests of Breusch and Pagan (1980), Pesaran (2004), and Pesaran *et al.* (2008), respectively.
The slope homogeneity tests are proposed by Pesaran and Yamagata (2008).

Tab. 2. Panel unit root test.

Variable	CIPS statistic
EXP	-1.926
IMP	-1.823

Critical values for the CIPS test are -2.15 (1%), -2.07 (5%), and -2.02 (10%), Pesaran (2007).

We test for panel non-stationarity among the variables before applying the cointegration test. The results of the cross-sectionally augmented IPS test are reported in Table 2. After inspection of the data, we only include a constant term (mainly due to measurement errors). When applying the Schwartz criterion to decide the optimal lag length, the common lag length was set to four. The table shows that all variables support the null hypothesis of panel non-stationarity. Furthermore, note that our approach does not exclude the possibility of including stationary variables.[3]

The likelihood ratio tests are reported in Table 3. The Bartlett corrected critical values are obtained by using the estimated model as data generating process when calculating the sample mean. Using the Bartlett corrected critical values, the test rejects the null of 0 cointegrating ranks but accepts the null of 1 cointegrating vector. Since the panel cointegration tests show that

[2] The CIPS panel unit root test is based on the Im, Pesaran and Shin (2001) test (IPS), which controls for cross-sectional heterogeneity in the estimated coefficients. The CIPS is the average of the individual country's cross-sectionally augmented ADF (CADF) statistics.

[3] The effect of one stationary variable in the system is that the rank order increases with one.

Tab. 3. Test for the cointegrating rank.

H_o	ACV^{a}	BCV^{b}	$-2logQ_T$
$R = 0$	536.11	647.22	605.49
$R \leq 1$	270.19	471.20	393.52
$R \leq 2$	103.35	275.18	167.38

Notes:
a. The asymptotic critical values at 5% significance level.
b. Bartlett corrected critical values at 5% significance level.

Tab. 5. Diagnostic tests[a].

Normality[b]	Autocorrelation[c]
0.038	0.507

Notes:
a. The table reports the p-values.
b. The test is a multivariate extension of the Bowman-Shenton test developed by Doornik and Hansen (1994).
c. This is the Ljung-Box test statistics for autocorrelation.

the common cointegrating rank is one, it is thus interesting to estimate the cointegrated vectors. The estimated cointegrating vectors, normalized for *IMP*, are presented in Table 4.

According to Table 4, we can assert that *EXP* is positively associated with *IMP* for almost all countries in the sample. Exceptions are Sweden, Italy and the Netherlands. In these countries, the coefficients have a very low value and are not significant. This implies that there is no long-run relationship between agricultural imports and exports in these countries. The lack of cointegration is probably a result of policy distortions or technological shocks. However, the magnitude of parameters varies from country to country.

The results from the diagnostic tests are given in Table 5. It seems that there is no problem with autocorrelation since the p-value is very high but the null hypothesis of normality is rejected and this problem could not be solved by using more lags.

4. POLICY IMPLICATIONS AND DISCUSSIONS

All theories of the trade balance assert that sustained deficits or surpluses might signal underlying policy problems. The elasticity approach suggests the real exchange rate and its effect on the demand and supply of traded goods as the key factor, while the absorption approach proposes that total expenditure is the most critical factor for understanding and correcting external account imbalances. The dynamics of the external accounts are explained by agents' responses to transitory and permanent shocks, in particular shocks in productivity. In the case of favourable productivity or technological shocks, investment booms tend to boost output growth but worsen the external accounts (Glick, Rogoff, 1995).

What does cointegration or lack of cointegration between agricultural imports and exports in the trade balance tell us about the state of the economy? The theory states that cointegration is to be expected under the maintained hypothesis that the economy is working properly and that breaking international budget constraints causes a lack of cointegration. (e.g., Trehan, Walsh, 1991; Hakkio, Rush, 1991; Husted, 1992; Bahmani-Oskooee, 1997; Irandoust, Sjöö, 2000; Herzer, Nowak-Lehman, 2006; Ariza, Bahmani-Oskooee, 2018; Afonso *et al*., 2020). This means that sustained external imbalances are the outcome of distorted markets or "bad policy". For understanding the cointegration results based on the international budget constraints, the conclusion is that lack of cointegration reveals fundamental policy problems unless there are permanent productivity shocks that lead to a non-stationary agricultural import-export relationship. In a well-functioning economy without permanent one-sided productivity shocks, cointegration is to be expected.

What are the policy implications of our findings? First, our findings of cointegration indicate that ten OECD countries, out of 13 under review, are not in violation of their international budget constraint as far as

Tab. 4. Cointegrating vectors normalized on *IMP*.

	Finland	Spain	Portugal	Italy	Austria	Denmark	Norway	Sweden
IMP	-1.000	-1.000	-1.000	-1.000	-1.000	-1.000	-1.000	-1.000
EXP	0.727	0.833	0.945	0.066	0.857	0.565	0.632	0.078
	Germany	France	UK	Switzerland	Netherlands			
IMP	-1.000	-1.000	-1.000	-1.000	-1.000			
EXP	0.510	1.052	1.116	0.912	0.059			

agricultural trade is concerned. Second, macroeconomic policies (such as fiscal and monetary policies) have been effective in bringing agricultural imports and exports to converge towards equilibrium in the long run. In the case of Sweden, Italy and the Netherlands, the lack of cointegration is a sign of bad policy or the existence of permanent technological shocks to the domestic economy. In other words, fundamental policy problems and the permanent productivity gap hypothesis lead to long-run agricultural trade imbalances. Third, the cointegration between agricultural imports and exports also provide support for intra-industry trade in the agriculture sector for almost all countries under review.

However, countries that suffer from longer-term "structural" external imbalances have to strongly concentrate their policy attention on a recovery of the tradable sector such as agriculture. This is not simply subject to real exchange rate adjustments or fiscal and monetary policies, as the expansion of export capacities requires strong investment in the tradable sector. This can be achieved by foreign direct investment (FDI), but since FDI flows have become smaller in the post-crisis period (Hunya, 2015), other domestic and policy instruments have to be applied. The focus here is on other industrial policy instruments that have to be adjusted to the specific requirements of OECD's peripheral economies (e.g., Landesmann, 2015). Combined with the use of innovative industrial policy instruments, there has to be an emphasis on institutional upgrading so that industrial policy intervention might show positive rather than negative results (Stöllinger, Holzner, 2013).

Although concern about real exchange rate developments is still valid, this has to be directed towards a joint sustained move towards supply-side improvements (i.e., targeting structural change and productivity improvements) as well as a consideration of balanced wage-productivity and human capital developments (Hanzl-Weiss, Landesmann, 2016). Growth and incomes policies combined with education, training and labour market policies should be included in a targeted policy that aims at competitive real exchange rate developments and not simply wage setting. Finally, capital markets policies or policies oriented towards attracting FDI should allocate capital towards the tradable sector rather than non-tradable activities (Hanzl-Weiss, Landesmann, 2016).

Generally speaking, if deficit countries are looking to improve their external balances permanently, they should assure that the capital flows stemming from abroad are allocated to tradable industries with high added value, avoiding the concentration of resources in non-tradable sectors in which the potential for increasing productivity is restricted. In other words, such countries should develop non-price competitive industries (Carrasco, Hernandez-del-Valle, 2017). This implies that a European industrial policy would create benefits by targeting resources towards the development of these industries. On the other hand, surplus countries should implement an expansive economic policy so as to boost domestic demand. An increase of domestic demand and a deterioration of the external balance in surplus countries could relieve the burden of deficit countries when trying to address external imbalances. Thus, addressing the persistent external European imbalances requires asymmetric responses from deficit and surplus countries, and the collaboration and coordination of economic policy between both groups of countries (Carrasco, Hernandez-del-Valle, 2017).

5. CONCLUSIONS

The purpose of this paper was to examine the long-run convergence of agricultural exports and imports in 13 industrialized OECD countries (France, Germany, Italy, Sweden, the UK, Switzerland, Denmark, Spain, Portugal, Finland, Austria, Norway and the Netherlands) over the period 1963-2020. Economic theory suggests that non-stationary agricultural trade flows in the trade balance will cointegrate in the long run. This is not the case if policy distortions exist or permanent technological shocks to the domestic economy. Thus, a natural tendency towards cointegration and convergence between agricultural exports and imports are expected in a well-functioning economy where there are neither permanent productivity shocks nor policy distortions.

The departures from earlier studies are in disaggregate agricultural trade flows and the asymptotic theory of likelihood-based panel cointegration allowing for multiple cointegrating vectors. The main contribution of this study stems from its methodology, which is a likelihood-based panel cointegration under assumptions of cross-sectional dependence and slope homogeneity restrictions. By using this method, the assumption of a unique cointegrating vector and the problem of normalization is relaxed, which is not the case with the usual residual-based tests of the cointegration approach. To the best of our knowledge, this is the first attempt to study cointegration between agricultural imports and exports.

Based on the likelihood-based panel cointegration technique, we found cointegration and convergence between agricultural exports and imports for almost all countries in the sample, but it was rejected for Sweden, Italy and the Netherlands. Our findings support the view that there is a stable underlying trend towards convergence between agricultural exports and imports in 10 OECD countries out of 13. The results also provide support for intra-industry trade in the agriculture sector for most of the countries under review.

It is worth to mentioning that other studies focus on the cointegration between the aggregate variables. Examples are Wickens and Uctum (1993), Bahmani-Oskooee (1994), Wu *et al*., (1996), Bahmani-Oskooee and Rhee (1997), Apergis *et al*., (2000); Irandoust and Sjöö, (2000); Holmes (2006), Konya and Singh, 2008; Chen (2011, 2014), Camarero *et al*. (2013), and Afonso *et al*., (2020).

This study has a few limitations. These stem from the fact that we used a linear, bivariate model without considering structural breaks since the likelihood panel cointegration model does not allow for structural shifts. Future studies should consider nonlinear and multivariate estimation methodology to account for structural breaks and regime shifts.

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APPENDIX A

Tab. A.1. Descriptive statistics of the variables, 1963-2020, n = 58 for each individual country.

Country	Mean	S.D.	Skewness	Kurtosis
Austria				
EXP	4.7765	3.2231	0.7910	2.4961
IMP	3.4139	1.3596	1.0229	3.2689
Switzerland				
EXP	0.7742	0.4479	0.1956	1.8689
IMP	2.2761	1.4511	0.8457	3.0635
Germany				
EXP	1.1335	0.3171	0.6510	2.3388
IMP	3.5216	2.3830	1.2061	3.5550
Denmark				
EXP	3.9372	1.2582	0.2954	1.4938
IMP	3.5420	1.1707	0.8616	2.8535
Spain				
EXP	1.7775	0.8603	2.2008	8.6913
IMP	3.9147	2.7023	0.6241	1.9390
Finland				
EXP	13.1630	8.9774	1.3772	4.0694
IMP	3.2135	1.0239	1.2382	3.9876
France				
EXP	1.9881	1.0933	0.7657	2.6097
IMP	3.5428	2.8339	1.5753	4.8509
Italy				
EXP	1.0619	0.5482	1.4370	4.4943
IMP	5.9746	3.6801	0.9097	3.0388
Netherlands				
EXP	3.7425	0.9809	1.0680	3.5263
IMP	2.7472	1.4851	1.5524	4.5914
Norway				
EXP	2.8134	2.8470	1.3975	3.8755
IMP	2.3061	0.9567	1.1063	3.3454
Portugal				
EXP	5.8596	3.1595	0.2035	1.4046
IMP	5.1661	3.7925	1.0891	3.4066
Sweden				
EXP	8.2365	4.8663	0.8971	2.7031
IMP	2.0890	0.7636	1.3368	4.9992
UK				
EXP	1.1938	0.7218	0.7107	2.2111
IMP	4.0373	3.6073	1.4568	4.2595

Fig. 1.-13. Agricultural exports and imports in the sample countries (1963-2020).

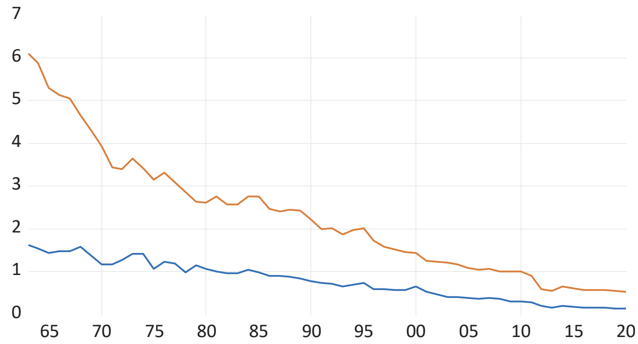

Fig. 1: Agricultural EXP and IMP in Switzerland (1963-2020).

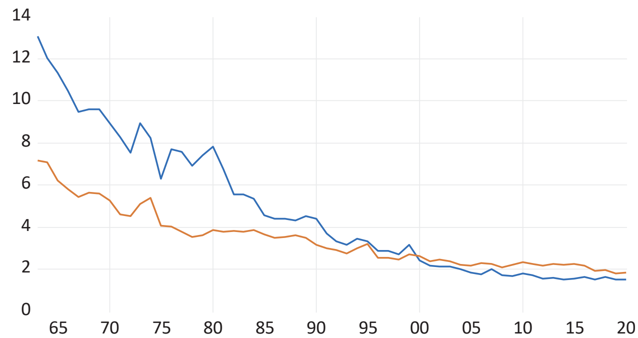

Fig. 2: Agricultural EXP and IMP in Austria (1963-2020).

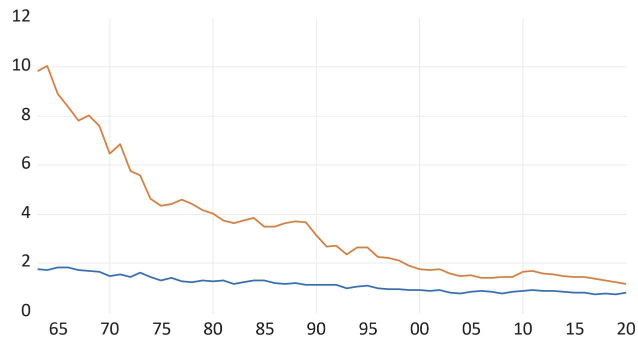

Fig. 3: Agricultural EXP and IMP in Germany (1963-2020).

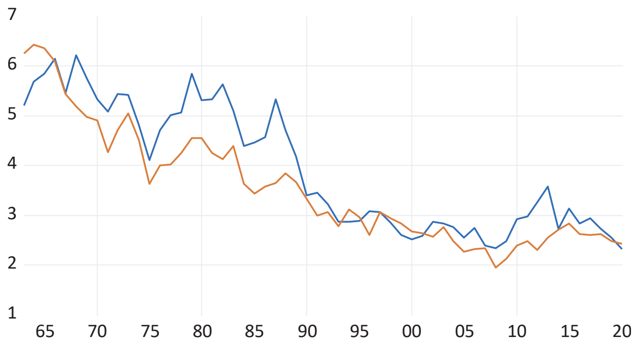

Fig. 4: Agricultural EXP and IMP in Denmark (1963-2020).

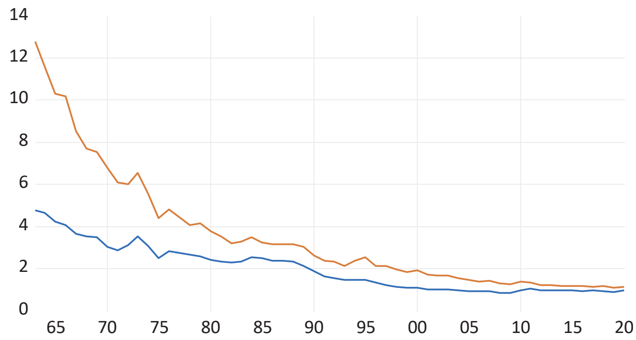

Fig. 7: Agricultural EXP and IMP in France (1963-2020).

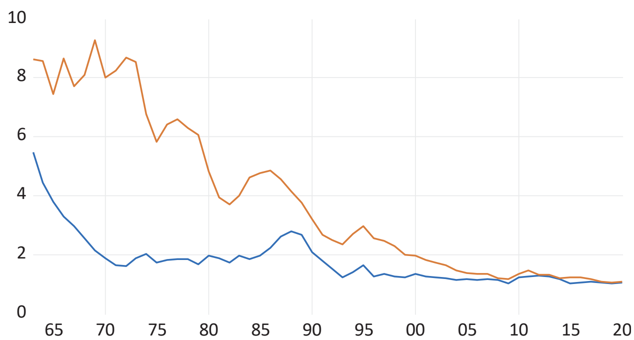

Fig. 5: Agricultural EXP and IMP in Spain (1963-2020).

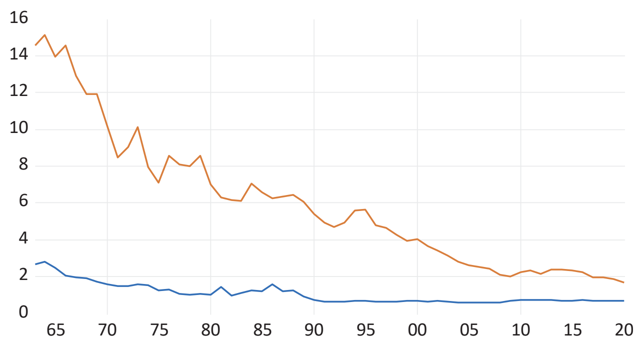

Fig. 8: Agricultural EXP and IMP in Italy (1963-2020).

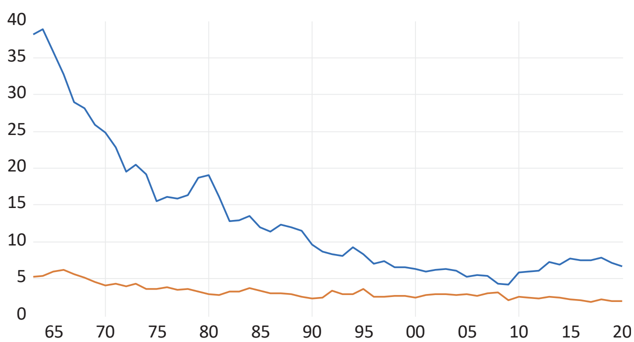

Fig. 6: Agricultural EXP and IMP in Finland (1963-2020).

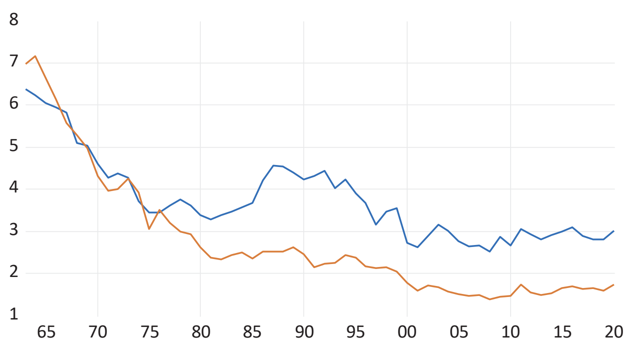

Fig. 9: Agricultural EXP and IMP in the Netherlands (1963-2020).

Fig. 10: Agricultural EXP and IMP in Portugal (1963-2020).

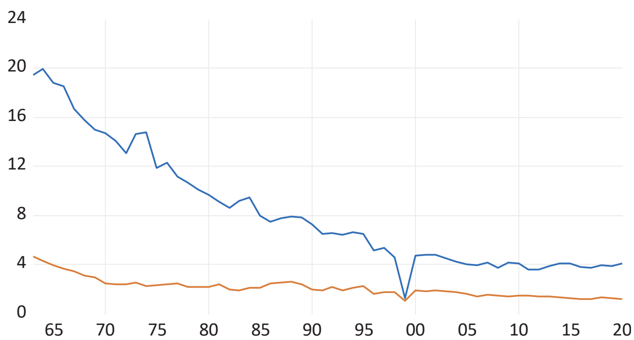

Fig. 11: Agricultural EXP and IMP in Sweden (1963-2020).

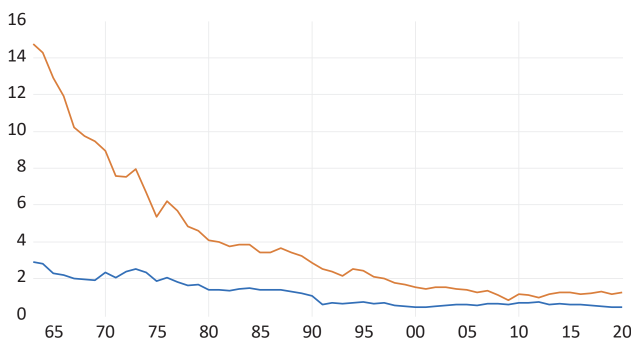

Fig. 12: Agricultural EXP and IMP in the UK (1963-2020).

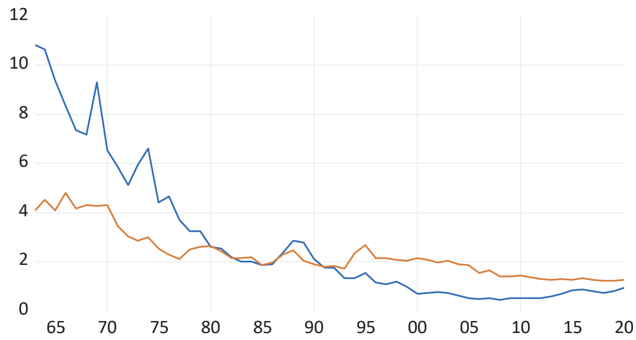

Fig. 13: Agricultural EXP and IMP in Norway (1963-2020).

APPENDIX B

Cross-sectional dependence tests

Breusch and Pagan's (1980) LM test has been used in many empirical studies to test cross-sectional dependency. LM statistics can be calculated using the following panel model:

$$y_{it}=\alpha_i+\beta^{\circ}_{it}\, x_{it}+\mu_{it}, \quad i=1,2,\ldots, N \quad t=1,2,\ldots, T, \qquad \text{1A}$$

where i is the cross-section dimension, t is the time dimension, x_{it} is $k \times 1$ vector of explanatory variables. while α_i and β_i are the individual intercepts and slope coefficients, respectively, that are allowed to differ across states. In the LM test, the null hypothesis of no cross-sectional dependence $H_0 : Cov(\mu_{it},\mu_{jt}) = 0$ for all t and $i \neq j$ is tested against the alternative hypothesis of cross-sectional dependence $H_1 : Cov(\mu_{it},\mu_{jt}) \neq 0$ for at least one pair of $i \neq j$. For testing the null hypothesis, Breusch and Pagan (1980) developed the following test:

$$CD_{BP}=T\sum_{i=1}^{N-1}\sum_{j=i+1}^{N}\hat{\rho}^{2}_{ij}, \qquad \text{2A}$$

where $\hat{\rho}^{2}_{ij}$ is the estimated correlation coefficient among the residuals obtained from individual OLS estimation of Eq. (1A). Under the null hypothesis, the LM statistic has an asymptotic chi-square distribution with $N(N-1)/2$ degrees of freedom. Pesaran (2004) proposes that the LM test is only valid when N is relatively small and T is sufficiently large. To overcome this problem, Pesaran (2004) introduces the following LM statistic for the cross-section dependency test:

$$CD_P = \sqrt{\frac{1}{N(N-1)}} \sum_{i=1}^{N-1} \sum_{j=i+1}^{N} \left(T \hat{\rho}_{ij}^2 - 1 \right) \quad 3A$$

However, Pesaran *et al.* (2008) state that while the population average pair-wise correlations are zero, the CD test will have less power. Therefore, they proposed a bias-adjusted test that is a modified version of the LM test by using the exact mean and variance of the LM statistic. The bias-adjusted LM statistic is calculated as follows:

$$LM_{adj} = \sqrt{\frac{2T}{N(N-1)}} \sum_{i=1}^{N-1} \sum_{j=i+1}^{N} \hat{\rho}_{ij}^2 \frac{(T-k)\hat{\rho}_{ij}^2 - u_{Tij}}{\sqrt{v_{Tij}^2}}, \quad 4A$$

where u_{Tij} and $v^2{}_{Tij}$ are the exact mean and variance of $(T\text{-}k)\hat{\rho}_{ij}^2$, which are provided in Pesaran *et al.* (2008). Under the null hypothesis of no cross-sectional dependence with $T \rightarrow \infty$ first followed by $N \rightarrow \infty$, the results of this test follow an asymptotic standard normal distribution.

Slope homogeneity tests

In order to relax the assumption of homoscedasticity in the F-test, Swamy (1970) developed the slope homogeneity test that examines the dispersion of individual slope estimates from a suitable pooled estimator. Pesaran and Yamagata (2008) state that both the F-test and Swamy's test require panel data models where N is relatively small compared to *T*. To overcome this problem, they proposed a standardized version of Swamy's test (the so-called $\tilde{\Delta}$ test) for testing slope homogeneity in large panels. The $\tilde{\Delta}$ test is valid when $(N, T) \rightarrow \infty$ without any restrictions on the relative expansion rates of *N* and *T* when the error terms are normally distributed. Pesaran and Yamagata (2008) then develop the following standardized dispersion statistic:

$$\bar{\Delta} = \sqrt{N}\left(\frac{N^{-1}\tilde{S} - k}{\sqrt{2k}} \right), \quad 5A$$

where $\tilde{S}$ is Swamy's statistic. Under the null hypothesis with the condition of $(N, T) \rightarrow \infty$ and when the error terms are normally distributed, the $\tilde{\Delta}$ test has an asymptotic standard normal distribution. The small sample properties of the $\tilde{\Delta}$ test can be improved when there are normally distributed errors by using the following mean and variance bias adjusted version:

$$\bar{\Delta}_{adj} = \sqrt{N}\left(\frac{N^{-1}\tilde{S} - E(\tilde{z}_{it})}{\sqrt{\operatorname{var}(\tilde{z}_{it})}} \right) \quad 6A$$

where $E(\tilde{z}_{it}) = k, var(\tilde{z}_{it}) = 2k(T-k-1))/(T+1)$.