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A cointegration and causality analysis of copper consumption and economic growth in rich countries

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ABSTRACT

The paper examines the copper consumption-economic growth nexus for 16 rich economies from the period 1966 to 2010. Various generations of panel unit root and cointegration tests are applied. Both series are found to be integrated of order one. Evidence of cointegration is found especially when controlling for breaks and long-run cross-sectional dependence. Causality is investigated using a vector error-correction mechanism (VECM) framework. At individual level, unidirectional causality running from economic growth to copper consumption is unraveled for Finland, France and UK in the long-run. Unidirectional causality is also found running from copper consumption to economic growth for Spain. Long-run bi-directionality between economic growth and copper consumption is found for Belgium, Greece, Italy, Japan and South Korea. The neutrality hypothesis holds for Australia, Austria, Canada, Netherlands, Portugal, Sweden and USA in the long-run. Taken as a whole, panel causality test reveals a long-run unidirectional causality running from economic growth to copper consumption.

Introduction

Copper is one of the oldest metals to have ever been exploited and is still a key input in many sectors such as in power, building, transport and many other sectors. For instance, given its malleability, electricity conductivity and corrosion-resistant capacity, copper is used in generators and cables to provide electricity. It is also utilized during construction of buildings for roofing, cladding, etc. and in the manufacturing of vehicles. These are essentially the products of rich countries and as such copper is closely connected to the economic development of a nation. Following Hricik (1988), while developed nations consume proportionally more copper than developing one, their demand has eventually leveled off. This is because there is a limit to the number of motor cars, washing machines, television sets and other equipment which a family can use.

According to the International Copper Study Group (2013), the copper market is expected to have a production surplus and this may of course affect the price of copper. But then again, this will affect copper consumption and potentially the economic growth of a country. It remains therefore necessary for policymakers to understand the causal link between copper consumption and economic growth in order manage their copper inventories and to meet their objectives of sustaining economic development. They can eventually use such information to make forecasts and enact mineral policies. So far, the literature has focused on time-series studies mainly and very few studies have been done for a group of countries. This paper presents a study of the link between copper consumption and economic growth using panel data from 16 rich countries[2] from the period of 1966–2010. Refined copper consumption data are obtained from the World Metal Statistics (various years) which are compiled from the World Bureau of Metal Statistics (2012) and those of real gross domestic product (GDP) are compiled from the World Development Indicators CD-ROM (2011).

The remainder of this paper is organized as follows: Section 2 reviews the existing literature. Section 3 discusses the testing framework. Section 4 presents the results. Section 5 concludes and provides some the policy implications.

Review of literature

The literature on the link between metal consumption and economic growth is somewhat scanty. Tilton (1989) investigates the consumption trend of six industrial metals in the OECD, USA,

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[2] The time frame and the selection of countries are purely dictated by the availability of data. The rich economies follow the classification of the World Bank (at http://go.worldbank.org/K2CKM78CC0).

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Japan and the EEC over two periods, 1960–1973 and 1973–1985. Such trend is found to be stimulated by increased economic activities. Roberts (1990) forecasts US steel consumption up to 2010 using data over the period 1963–1983 and pinpoints to the importance of the gross national product (GNP) in determining metal use. Labson and Crompton (1993) analyze the link between five industrial metals and income for the USA, UK, Japan and OECD for the period of 1960–1987 and discover little evidence of a long-run relationship between the two variables.

Ghosh (2006) studies the link between steel consumption and economic growth in India for the period beginning in 1951–1952 and ending in 2003–2004. He finds no evidence of cointegration but he finds existence of a unidirectional causal effect of economic growth on steel consumption. In similar fashion, Huh (2011) investigates the link between steel consumption and economic activity for Korea over the period of 1975–2008. He discovers evidence of a long-run relationship between total steel consumption and GDP, running from GDP to total steel consumption. Jaunky (2012) studies the link between aluminum consumption and economic growth for 20 rich countries. His study reveals a unidirectional causality running from aluminum consumption to economic growth in the short-run and vice-versa in the long-run.

The testing framework

Econometric tests such as unit root and cointegration tests are necessary before assessing any causal relationship between copper consumption and economic growth. Most of the panel unit root tests are based on an augmented Dickey-Fuller (ADF) unit root test type. Let $LCOP_{it}$ denotes a variable, thus:

$$\Delta LCOP_{it} = r_i + g_i t + \delta LCOP_{it-1} + Z_{im} \sum_{j=1}^{\rho_i} \Delta LCOP_{it-j} + e_{it} \tag{1}$$

where $\Delta LCOP_{it} = LCOP_{it} - LCOP_{it-1}$, t is the time trend, ρ is the lag length and e is the error term. If the null hypothesis (H_0) is accepted (i.e. H_0: $\widehat{\delta} = 0$), then the series is non-stationary. For example, $LCOP_{it}$ is be integrated order of d, i.e. $LCOP_{it} \sim I(d)$, if it were to be differenced by d times to come to be stationary.

As a preliminary step, some time series tests such as the ADF, Zivot and Andrews (1992) and Narayan and Popp (2010) will first be computed for individual countries. Various generations have of panel unit root tests are then to be considered. For instance, first generation tests include those of Levin et al. (2002, LLC), Im et al. (2003, IPS) and Im et al. (ILT, 2005). The ILT test can specifically accounts for endogenous breaks. But, these tests assume independence of individual cross-sections and this is unlikely to hold in practice. Pesaran (2007) further proposes a second generation test which allows for different forms of cross-sectional dependence. Finally as a third generation, Chang and Song (2009) propose a test which can tackle any forms of dependence whether short-run or long-run.

Unit root tests are generally computed via two different regressions. One regression includes a constant term only while the other one includes both a constant term and a time trend. Macroeconomic data usually tend to be non-stationary and exhibit a trend over time. It is therefore more fitting to consider a regression with a constant and a trend at level form. First-differencing tends to remove any deterministic trends in the series. The unit root regression should include a constant term only. For sake of comparison, both regressions are considered.

Next, assuming both series are non-stationary and integrated of the same order, panel cointegration tests can be performed. Pedroni (1999, 2004) was among the firsts to propose testing the panel cointegration. This first generation test assumes cross-sectional independence. With regard to second generation tests, Westerlund (2007) and Westerlund and Edgerton (2008) suggest some panel tests of the H_0 of no cointegration which allows for cross-sectional dependence. The latter allows for unknown structural breaks in both the intercept and slope of the cointegrating regression, which can be located at different periods for different countries. The Westerlund and Edgerton test is supplemented by a bootstrap panel cointegration test with breaks as proposed by Di Iorio and Fachin (2010). As a third generation test, Di Iorio and Fachin (2012) devise a bootstrap panel cointegration test which is robust to short-run and long-run dependence across countries. In the event the residual series is stationary, then the variables are cointegrated.

The relationship between refined copper consumption and economic growth can subsequently be investigated. In the incidence of lack of theoretical support to empirically test this relationship, Granger causality results can provide useful information which can assist the development of new theories or in the refinement of existing ones (Farr et al., 1998). Since, the Granger causality test requires the variables to be stationary, a panel vector error-correction mechanism (VECM) causality test is employed.

The ρth order of a panel VECM structure can be algebraically represented as follows:

$$\begin{bmatrix} \Delta LCOP_{it} \\ \Delta LGDP_{it} \end{bmatrix} = \begin{bmatrix} \alpha_1 \\ \alpha_2 \end{bmatrix} + \sum_{k=1}^{\rho} \begin{bmatrix} \phi_{11k} & \phi_{12k} \\ \phi_{21k} & \phi_{22k} \end{bmatrix} \begin{bmatrix} \Delta LCOP_{it-k} \\ \Delta LGDP_{it-k} \end{bmatrix} + \begin{bmatrix} \omega_1 \\ \omega_2 \end{bmatrix} [ECM_{it-1}] + \begin{bmatrix} \kappa_1 \\ \kappa_2 \end{bmatrix} [D_{itb}] + \begin{bmatrix} f_{1i} \\ f_{2i} \end{bmatrix} + \begin{bmatrix} \varepsilon_1 \\ \varepsilon_2 \end{bmatrix} \tag{2}$$

where $LCOP_{it}$ and $LGDP_{it}$ denote the natural logarithm of refined copper consumption (in metric tonnes) and GDP (at constant 2000) for country i over year t respectively. The later is usually used to capture economic growth. $i = 1, 2, \ldots, N$, $t = \rho + 1, \rho + 2, \ldots, T$, Δ denotes first differences while α's, ϕ_k's, ω's and κ's are parameters to be estimated. The descriptive statistics of both variables are reported in Table 1. ECM_{it-1} represents the one period lagged error-term derived from the cointegrating vector and the coefficients on the ECMs represent how fast deviations from the long-run equilibrium are eliminated. f_{1i} and f_{2i} are the fixed effects components and the error terms ε_1 and ε_2 are serially independent with mean zero and finite covariance matrix. D_{itb} is the dummy capturing economic shocks which an economy i has been subjected to at a given point in time t. b denotes the point at which the break occurs. The break dates can be obtained from the time-series unit root tests. These dates are utilized as a proxy for economic shocks.

A Wald test for joint significance can be applied to determine the direction of any causal relationship. The results from this test should be interpreted as indicating whether previous changes in one variable contribute significantly to the prediction of the future value of the other variable. Economic growth does not Granger-cause copper consumption if and only if all of the coefficients $\phi_{12k}; \forall = 1, 2, \ldots, \rho$ are not significantly different from zero in Equation (3). The dependent variable reacts only to short-term shocks. These can be referred to as the "*short-run causality*" tests. Another channel of causality can be investigated by testing the significance of the ECMs. This test can be denoted as the "*long-run causality*" tests. If no causality is found, then the "*neutrality hypothesis*" holds. Granger causality does not imply true causality (Granger, 1969). For example, if $LCOP_{it}$ Granger-causes $LGDP_{it}$, then $LCOP_{it}$ is useful in forecasting $LGDP_{it}$. According to this framework, "useful" implies the ability of $LCOP_{it}$ to increase the accuracy of the prediction of $LGDP_{it}$ with respect to a forecast, considering only past values of $LGDP_{it}$ (Foresti, 2006).

Dynamic panels tend to suffer from endogeneity problem. The lagged dependent variable appears as an explanatory variable in these models and the former contains the fixed effects such as f_{1i} and f_{2i}. Traditional panel data techniques such as pooled ordinary least square (OLS), fixed-effects or random-effects models cannot effectively deal with this problem as they tend to produced biased estimates. To remedy the correlation and endogeneity problems, Arellano and Bond (1991) propose a two-step difference generalized method of moments (GMM) technique where the lags of explanatory variables in levels are utilized like instruments. But this GMM estimator suffers from a lack of power of the internal instruments.[3] Blundell and Bond (1998) recommended a two-step system GMM estimator which has superior finite-sample properties.

The causality results can be applied to investigate whether there is any potential predictability power of one variable for the other. Moreover, the direction of has significant policy implications especially for mineral conservation policies. Mineral conservation is not only concerned with efficient extraction and use but also preservation and saving of minerals (Myers, 1946). If there is no causality, then adopting mineral conservation policies to limit the copper consumption can be implemented as these will not impacting negatively on economic growth. A fall in copper consumption can lead copper industries to a reduction in production. This will eventually cause a reduction in the exploitation of natural resources and environmental degradation.[4] If causality runs from economic growth to copper consumption, mineral conservation policies can be implemented. Mineral conservation policies[5] could be enacted to control the activities of the copper industries. These policies will have no impact on economic growth. If a unidirectional causality running from copper consumption to economic growth exists, then mineral conservation policies will have an adverse impact on economic growth.

Results

The ADF unit root statistics for individual countries are reported in Table 2(a)(i) and (a)(ii). Following the above-discussed economic perception, both the $LCOP_t$ and $LGDP_t$ series are found to be I(1) for Australia, Canada, Italy, Netherlands, Portugal, Sweden, UK and USA. The ADF test ignores the presence of structural breaks in the series and this can lead to the test to be biased towards the non-rejection of the null hypothesis (Perron, 1989). These breaks can be attributed to international recession, oil price hikes, technological advances, etc. Furthermore, the Zivot-Andrews test can account for one endogenous structural break in the series and its statistics are in Table 2(b)(i) and (b)(ii), With the exception of France, Japan, Portugal, Spain and UK, copper and income series for the remaining countries are I(1).

In addition, the Narayan-Popp test accounts for the presence of two endogenous structural breaks and adds more power to the testing framework. As per Table 2(c)(i) and (c)(ii), the $M1_{B,L}$ test reflects the test equation for two breaks in the level of a trending series while the $M2_{B,L}$ test captures the test equation for two breaks in the level and slope of a trending series. The $M1_{B,L}$ test reveals an I(1) process for both $LCOP_t$ and $LGDP_t$ series, excluding Austria, Belgium, France and Spain. Same process applies to Canada, Japan, South Africa and USA when computing the $M2_{B,L}$ test. For many countries the first break tends to occur around the mid 1970s and early 1980s. This coincides with the supply shocks caused by the Intergovernmental Council of Copper Exporting Countries (CIPEC) to limit production in 1975 (Stürmer, 2013). The break points can as well be explained by the demand shocks caused by the excessive hike in oil price following the 1979 Iranian Revolution and the 1980–1981 Iran–Iraq War. The second break tends to occur around early 1990s and 2000s. These periods coincide with another spike in oil price mainly due to the 1990 Gulf War and to the 1997 Kyoto Protocol.

Yet, as argued by Toda (1995), even 100 observations are not enough to ensure good performance of the time-series tests. Consequently, the latter can be subject to the criticism of low power and spurious results. This raises need to apply panel data techniques which allow for a sizeable increase in testing power especially in context of testing for unit roots and cointegration amongst the data. Table 4(a) reports the LLC test statistics for both series are presented, where $LCOP_{it} \sim I(1)$ and $LGDP_{it} \sim I(0)$. Nonetheless, the assumptions of the LLC test are rather restrictive. The LLC test is based on the assumption of homogeneity in the autoregressive of order one (AR(1)) coefficients of the ADF specifications. It ignores structural breaks in addition to cross-sectional dependence issues.

Corresponding to Banerjee et al. (2004), cross-sectional dependence biases the panel data unit root tests towards the alternative hypothesis. The degree of cross-sectional dependence can be assessed by inspecting the pair-wise correlations between changes in the variables (Koedijk et al., 2004). As presented in Table 3, the pair-wise correlations of the first-differences in two series are generally positive and in some cases rather large. For instance, the pair-wise correlation of $\Delta LCOP_t$ between Australia and Canada is 0.46 while for $\Delta LGDP_t$, it is equal to 0.78 between France and Netherlands. The pair-wise correlation coefficients of $\Delta LCOP_t$ and $\Delta LGDP_t$ range from -0.30 to 0.66 and from 0.05 to 0.88 respectively. This provides strong evidence of cross-sectional dependence.

The IPS test allows for heterogeneity between groups and control for cross-sectional dependence using demeaned data. As shown in Table 4(b), the IPS test provides evidence of an I(1) process for both variables, especially when using the demeaned data. This result is contrasted when controlling for structural breaks. The ILT test indicates a stationary process for both $\Delta LCOP_{it}$ and $\Delta LGDP_{it}$. All the three first-generation tests suffer from size distortions and have low power in the presence of cross-sectional dependence (Herwartz and Siedenburg, 2008). The IPS test exploits demeaned data to tackle this issue but this method can be problematic. Assuming the existence of one common factor with the same effect on all the individuals and as such is rather unrealistic and the demeaning may not get rid of the size problem (Strauss and Yigit, 2003). Even the ILT test fails to efficiently control for cross-sectional dependence.

As a second generation test, Pesaran (2007) proposes a test which allows for the presence of more general cross-sectional dependence patterns. For instance, the standard ADF regression models are augmented with the cross-section averages of lagged levels and first-differences of the individual series. This test is based on the averages of the individual cross-sectionally augmented ADF (CADF) statistics and is found to have good size and power properties even if N and T are relatively small. As reported in Table 4(d), both $LCOP_{it}$ and $LGDP_{it}$ series are found to be I(1).

A third generation panel unit test controls for cross-sectional cointegration which arises when there are short-run and long-run

[3] For these instruments to be valid there should not be serial correlation in ε_1 and ε_2. The optimal lag length, ρ, is chosen when no serial correlation is obtained in the residuals. The autocorrelation test has a null hypothesis of no autocorrelation and is applied to the differenced residuals. The tests of AR(1) and AR(2) process in the first-differences are computed. Tthe test for AR(2) in first differences is considered to be more important since it will detect autocorrelation in levels (Mileva, 2007). To test for the joint validity of the instruments the Hansen (1982) J and the Sargan (1958) tests are employed. Too many weak instruments can overload the endogenous variables and reduce the accuracy of the tests (Roodman, 2009). A rule of thumb for is to keep the number of instruments less than or equal to the number of groups (Mileva, 2007).

[4] Copper production entails greenhouse gases which tend to be principally sulfur dioxide (SO_2) emissions (Ayres et al., 2002)

[5] Myers (1946) discusses the various implications of adopting mineral conservation policies in the USA.

Table 1
Descriptive Statistics.
Source: Computed.

Countries	$LCOP_{it}$				$LGDP_{it}$			
	Mean	Std. Dev.	Minimum	Maximum	Mean	Std. Dev.	Minimum	Maximum
Australia	11.781	0.185	11.393	12.147	26.342	0.433	25.541	27.057
Austria	10.260	0.283	9.480	10.756	25.648	0.341	24.951	26.151
Belgium	12.470	0.363	11.503	12.918	25.870	0.309	25.221	26.317
Canada	12.302	0.179	11.864	12.613	26.926	0.389	26.190	27.494
Finland	11.049	0.445	10.278	11.660	25.177	0.370	24.461	25.754
France	12.923	0.249	12.173	13.260	27.607	0.321	26.911	28.040
Greece	10.588	0.860	8.612	11.810	25.291	0.337	24.509	25.828
Italy	12.960	0.369	12.181	13.593	27.431	0.314	26.717	27.817
Japan	13.948	0.252	13.087	14.294	28.826	0.399	27.846	29.280
Netherlands	10.397	0.404	9.642	10.882	26.305	0.348	25.606	26.832
Portugal	8.393	1.863	3.828	10.240	25.039	0.425	24.115	25.567
South Korea	11.961	1.749	8.039	13.753	26.147	0.899	24.497	27.408
Spain	12.002	0.456	11.091	12.716	26.722	0.392	25.933	27.331
Sweden	11.725	0.277	11.303	12.223	25.966	0.283	25.439	26.441
UK	12.619	0.734	10.035	13.292	27.690	0.327	27.165	28.223
USA	14.570	0.163	14.149	14.922	29.510	0.395	28.845	30.088
Panel	11.872	1.642	3.829	14.923	26.656	1.311	24.115	30.088

co-movements among the variable. Chang and Song (2009) propose a test which employs of a set of orthogonal functions as instrument generating function (IGF) for controlling any long-run dependence. As exposed in Table 4(e), two different types of panel unit root test statistics are calculated. The average tests relate to the testing of the H_0 of non-stationarity for all individual countries while the minimum tests evaluate the H_0 of non-stationarity of some individual countries within the panel. With the exception of ta_a, all test statistics confirm an I(1) process for $LCOP_{it}$. Similar result is obtained for $LGDP_{it}$ when considering the ta_c and tm_c panel statistics.

For consistent inferences, panel unit root tests require a non-stationary process for individual series. According to Karlsson and Lothgren (2000), rejection of the panel unit root null may be driven by a few stationary series and the whole panel can be incorrectly modeled as stationary. For instance, the Chang-Song test are recomputed $LCOP_{it}$ by excluding the French and British series. These are I(0) as per the Narayan-Popp $M2_{B,L}$ test. No major difference to the results has been identified. The panel unit root tests allow for a substantial increase in testing power and the results corroborate with the *a-priori* expectation of an I(1) process.[6] Panel cointegration tests are next computed.

As shown in Table 5(a), seven Pedroni test statistics with the H_0 of no cointegration. Four of these statistics are called panel cointegration statistics. These are panel-ν, panel-ρ, and panel-pp which denote the non-parametric variance ratio, Phillips-Perron ρ, and student's t-statistics respectively while panel-adf is a parametric statistic based on the ADF statistic. The extra three statistics are called group mean panel cointegration statistics. These are the group-ρ, group-pp and group-adf which correspond to Phillips-Perron ρ-statistic, Phillips-Perron t-statistic and the ADF-statistic respectively. The three statistics allow the modeling of potential heterogeneity across the panel. Only the panel-ν statistic without trend rejects the null. In general, no evidence of a cointegrating relationship is found.

As reported in Table 5(b), four Westerlund panel test statistics of the H_0 of no cointegration are computed. The panel tests denoted by Gt and Ga are performed under the alternative hypothesis of panel cointegration, while Pt and Pa are performed under the alternative hypothesis that at least one element of the panel is cointegrated. The H_0 of no cointegration which assumes that the error-correction term in a conditional error-correction model is equal to zero is tested. The H_0 cannot be rejected especially after controlling for cross-sectional dependence. Similar result of no cointegration is obtained when computing the Westerlund and Edgerton test. As presented in Table 5(c), the $Z_\tau(N)$ and $Z_\phi(N)$ tests control for heteroskedasticity, serial correlation, cross-sectional dependence. These tests allow for unknown endogenous breaks in both the intercept and slope of the cointegrating regression, which may be located at different dates for different countries. All the test statistics fail to of reject the H_0 of no cointegration.

By and large, to obtain good performances of the cointegration tests, large sample sizes are often needed. According to Di Iorio and Fachin (2010), the power of the Westerlund and Edgerton test is found to be unsatisfactory even when $T=100$. Through a bootstrap algorithm, the former also implements a cointegration test which can control for both endogenously structural breaks and cross-sectional dependence. Their test is found to have relatively good size and power properties. When referring to the mean, median and maximum ADF cointegration test statistics in Table 5(d), the H_0 of no cointegration is rejected. Evidence of a long-run relationship between copper consumption and economic growth is supported for the 16 high-income countries.

Existence of cross-sectional cointegration may have biased the results. The Di Iorio and Fachin (2012) test for long-run cross-sectional dependence can be referred to as a residual-based Stationary Bootstrap (RSB) test and is based on the unit root test of Parker et al. (2006). The RSB test does not require any assumptions on the form of time dependence of the residuals and has the advantage of using stationary resampled pseudo-series. As shown in Table 5(e), all the mean, median and maximum ADF cointegration test statistics strongly reject the H_0 of no cointegration. This study therefore highlights the importance for policymakers to employ state-of-the-art panel data techniques in analyzing the relationship copper consumption and economic growth. These recently devised tests have provided evidence of an I(1) process for both $LCOP_{it}$ and $LGDP_{it}$ series while supporting the prevalence of a cointegrating relationship.

Results of the causality test are illustrated in Table 6. The economic shocks dummy is constructed by using the breaks dates of the $M2_{B,L}$ Narayan-Popp model for both $LCOP_t$ and $LGDP_t$ at level form. These shocks have a significant impact for Greece, South

[6] Jaunky (2013) uncovers a non-stationary process for copper consumption for a group of 37 countries.

Table 2
(a)(i): ADF statistics for individual countries at level form. (a)(ii): ADF statistics for individual countries at first-difference. (b)(i): Zivot-Andrews time series unit root tests at level form. (b)(ii): Zivot-Andrews time series unit root tests at first-difference. (c)(i): Narayan-Popp Time series unit root tests at level form. (c)(ii): Narayan-Popp time series unit root tests at first difference.
Source: Computed.

(a)(i)

Countries	$LCOP_t$				$LGDP_t$			
	With constant and without trend		With constant and with trend		With constant and without trend		With constant and with trend	
	ADF	ρ	ADF	ρ	ADF	ρ	ADF	ρ
Australia	−2.26	0	−2.685	0	0.276	0	−2.031	0
Austria	−1.925	1	−2.056	1	−2.351	0	−3.335‡	0
Belgium	−2.638‡	0	−1.686	0	−2.168	0	−3.283‡	0
Canada	−2.221	0	−2.152	0	−1.704	1	−3.215	1
Finland	−1.669	0	−0.666	0	−1.212	2	−3.221‡	3
France	0−0.660	0	1.986	3	−2.481	1	−2.577	1
Greece	−1.988	0	1.470	0	−1.565	1	−2.909	3
Italy	−1.286	0	−2.850	1	−3.671*	0	−0.480	0
Japan	−2.661‡	2	−1.502	2	−3.630+	2	−1.001	2
Netherlands	−1.294	0	−1.637	0	−1.030	1	−2.321	1
Portugal	−1.337	4	−1.720	4	−2.528	4	−1.099	4
South Korea	−4.661*	1	−1.462	1	−2.588	0	−0.113	0
Spain	−0.122	3	−2.367	0	−1.293	1	−4.347*	4
Sweden	−1.337	0	−1.819	4	0.343	2	−2.528	1
UK	1.110	3	−2.457	2	−0.105	2	−2.627	1
USA	−2.158	0	−1.948	0	−1.362	2	−3.026	1

(a)(ii)

Countries	$LCOP_t$				$LGDP_t$			
	With constant and without trend		With constant and with trend		With constant and without trend		With constant and with trend	
	ADF	ρ	ADF	ρ	ADF	ρ	ADF	ρ
Australia	−7.236*	0	−7.221*	0	−5.724*	0	−5.691*	0
Austria	−8.647*	0	−8.760*	0	−5.494*	0	−5.796*	0
Belgium	−6.081*	0	−6.770*	0	−5.931*	0	−6.284*	0
Canada	−7.270*	0	−7.301*	0	−4.337*	0	−4.543*	0
Finland	−6.781*	0	−7.635*	0	−4.491*	1	−4.533*	1
France	−0.771	1	−1.492	1	−4.266*	0	−4.822*	0
Greece	−0.083	4	−5.509*	0	−4.423*	0	−4.453*	0
Italy	−3.553+	4	−3.851+	2	−4.809*	0	−5.574*	1
Japan	−6.468*	2	−7.305*	1	−2.373	2	−5.269*	1
Netherlands	−6.761*	0	−6.940*	0	−4.580*	0	−4.602*	0
Portugal	−6.833*	1	−6.751*	1	−3.289+	3	−4.398*	1
South Korea	−0.965	4	−6.682*	1	−5.033*	0	−5.712*	0
Spain	−2.756‡	2	−2.746	2	−2.767‡	0	−2.851	0
Sweden	−6.113*	0	−6.034*	0	−4.773*	1	−4.729*	1
UK	−3.550‡	2	−3.986+	2	−4.510*	1	−4.443*	1
USA	−5.595*	1	−5.746*	1	−4.823*	1	−4.945*	1

Table 2 (*continued*)

(b)(i)

Countries	$LCOP_t$						$LGDP_t$					
	Without trend			With trend			Without trend			With trend		
	Min. *t*-stat.	Break date	ρ	Min. *t*-stat.	Break date	ρ	Min. *t*-stat.	Break date	ρ	Min. *t*-stat.	Break date	ρ
Australia	−3.550	1993	0	−4.405	2002	0	−4.250	1998	1	−4.085	1998	1
Austria	−3.978	1975	1	−4.071	1987	1	−2.644	2001	0	−3.062	1973	0
Belgium	−3.260	1976	0	−3.875	1976	0	−2.951	2001	0	−3.516	1973	0
Canada	−3.067	1996	0	−3.477	1998	0	−3.306	1990	1	−3.530	1990	1
Finland	−2.437	2001	0	−3.985	1997	0	−5.331+	1991	1	−5.343	1991	1
France	1.146	2003	1	−2.453	2003	1	−3.314	1976	1	−3.337	2001	1
Greece	−0.483	2003	0	−3.581	2003	0	−3.467	1981	1	−3.334	2003	1
Italy	−2.842	1981	1	−3.691	2000	1	−2.441	1976	0	−2.947	2002	0
Japan	−3.254	1978	2	−4.276	1990	2	−3.099	1985	2	−5.874*	1990	2
Netherlands	−4.388	1995	0	−4.511	1995	0	−3.685	1981	1	−3.970	1997	1
Portugal	−6.058*	1975	0	−5.931*	1995	0	−3.028	2003	1	−3.545	2000	1
South Korea	−1.153	1976	2	−2.268	1977	2	−1.887	1998	0	−3.463	1994	0
Spain	−3.350	1993	0	−3.193	1998	0	−3.945	1981	1	−3.879	2003	1
Sweden	−3.653	1992	0	−3.705	1997	0	−3.972	2002	1	−4.708	1991	1
UK	−3.081	2003	2	−5.541+	2003	2	−4.157	1997	1	−4.048	1980	1
USA	−3.640	2002	0	−5.052	1997	0	−3.398	1987	1	−4.768	1999	1

(b)(ii)

Countries	$LCOP_t$						$LGDP_t$					
	Without trend			With trend			Without trend			With trend		
	Min. *t*-stat.	Break date	ρ	Min. *t*-stat.	Break date	ρ	Min. *t*-stat.	Break date	ρ	Min. *t*-stat.	Break date	ρ
Australia	−6.920*	1992	1	−7.189*	1992	1	−5.772*	1993	0	−6.337*	1984	0
Austria	−9.646*	1984	0	−9.921*	1984	0	−6.678*	1975	0	−6.630*	1985	0
Belgium	−7.893*	1979	0	−7.947*	1978	0	−7.926*	1975	0	−7.792*	1975	0
Canada	−8.809*	1993	0	−9.566*	1993	0	−5.599*	1994	0	−5.566*	1994	0
Finland	−8.445*	2001	0	−8.360*	1979	0	−6.151*	1996	1	−6.077*	1996	1
France	−2.674	2003	1	−4.214	2003	1	−5.796*	1975	0	−5.681*	1975	0
Greece	−7.086*	2003	0	−8.500*	1993	0	−5.565*	1974	0	−5.461+	1974	0
Italy	−5.459*	1986	2	−5.717*	1998	2	−5.878*	1975	1	−5.959*	1985	1
Japan	−7.276*	1980	1	−7.599*	1976	1	−5.346+	2003	1	−6.134*	1977	1
Netherlands	−7.795*	1994	0	−8.026*	1994	0	−4.983+	1988	0	−5.714*	1983	0
Portugal	−8.009*	2002	1	−9.383*	1995	1	−5.700*	1974	0	−5.772*	1985	0
South Korea	−9.118*	1982	0	−9.482*	1979	0	−6.936*	1983	0	−6.939*	1983	0
Spain	−3.469	1986	2	−3.698	1982	2	−3.999	1974	0	−4.001	1983	0
Sweden	−6.040*	2001	1	−6.189*	1978	1	−5.659*	1994	1	−5.594*	1984	1
UK	−4.941+	2002	2	−5.903*	1995	2	−5.013+	1985	1	−5.578*	1997	1
USA	−7.307*	2001	1	−7.151*	2001	1	−5.848*	1983	1	−5.954*	1996	1

(c)(i)

Countries	$LCOP_t$								$LGDP_t$							
	$M1_{B,L}$				$M2_{B,L}$				$M1_{B,L}$				$M2_{B,L}$			
	t-value	T_{B1}	T_{B2}	ρ	*t*-value	T_{B1}	T_{B2}	ρ	*t*-value	T_{B1}	T_{B2}	ρ	*t*-value	T_{B1}	T_{B2}	ρ
Australia	−3.745	1974	1992	0	−3.939	1974	1990	1	−3.552	1982	1991	4	−2.262	1982	1992	0
Austria	−0.214	1979	1986	4	−1.305	1979	1986	3	−1.942	1974	1977	0	−3.755	1974	1978	1

Table 2 (*continued*)

(c)(i)

Countries	*LCOP$_t$*								*LGDP$_t$*							
	M1$_{B,L}$				M2$_{B,L}$				M1$_{B,L}$				M2$_{B,L}$			
	t-value	T$_{B1}$	T$_{B2}$	ρ	*t*-value	T$_{B1}$	T$_{B2}$	ρ	*t*-value	T$_{B1}$	T$_{B2}$	ρ	*t*-value	T$_{B1}$	T$_{B2}$	ρ
Belgium	−4.294‡	1975	1988	0	−4.754	1980	1992	0	−1.395	1974	1976	0	−3.184	1974	1987	0
Canada	−1.41	1974	1981	0	−0.169	1974	1981	0	−2.176	1981	1990	1	−4.314	1981	1998	1
Finland	−0.074	1975	1978	3	−3.312	1974	2001	0	−3.841	1990	1992	3	−3.643	1990	1993	3
France	3.535	1974	1979	1	4.794‡	1978	1983	1	−4.167‡	1975	1987	1	−3.36	1974	1987	1
Greece	2.321	1979	1992	4	2.039	1979	1997	4	−3.972	1974	1986	0	−2.403	1974	1978	3
Italy	−2.057	1975	1997	0	−3.334	1974	1997	0	−0.905	1974	1976	2	−0.635	1974	1997	1
Japan	−3.349	1977	1988	2	−3.982	1997	2000	0	−1.738	1987	1997	2	−4.03	1974	1997	1
Netherlands	−3.345	1979	1994	0	−4.538	1979	1994	0	−2.751	1975	1980	1	−2.399	1975	1980	1
Portugal	−3.865	1994	1997	0	−4.2	1994	1998	4	−1.763	1975	1990	4	0.163	1975	1979	4
South Korea	0.604	1976	1980	4	−0.878	1974	1980	1	0.708	1979	1997	0	−3.108	1979	1997	0
Spain	−1.092	1974	1980	0	−2.38	1974	1981	3	−3.283	1974	1980	1	−3.385	1975	1993	1
Sweden	−2.032	1974	1977	0	−2.322	1977	2001	0	−1.535	1976	1992	0	−3.074	1976	1992	1
UK	−2.729	1981	1992	2	−6.829*	1994	2001	2	−3.491	1979	1996	1	−3.159	1979	1996	1
USA	−2.438	1974	1981	0	−1.148	1975	1983	0	−3.213	1981	1998	1	−2.282	1975	1981	1

(c)(ii)

Countries	*LCOP$_t$*								*LGDP$_t$*							
	M1$_{B,L}$				M2$_{B,L}$				M1$_{B,L}$				M2$_{B,L}$			
	t-value	T$_{B1}$	T$_{B2}$	ρ	*t*-value	T$_{B1}$	T$_{B2}$	ρ	*t*-value	T$_{B1}$	T$_{B2}$	ρ	*t*-value	T$_{B1}$	T$_{B2}$	ρ
Australia	−6.834*	1977	1992	1	−6.750*	1977	1992	1	−8.015*	1982	1990	0	−1.708	1982	1992	3
Austria	−1.678	1979	1986	4	−2.8	1979	1986	2	−4.306‡	1977	1979	4	−4.917‡	1975	1978	0
Belgium	−7.837*	1976	1980	0	−7.217*	1977	1992	0	−6.378*	1980	1987	0	−4.687	1975	1980	0
Canada	−8.897*	1981	1992	0	−8.965*	1981	1992	0	−6.114*	1981	1990	0	−6.145*	1981	1993	1
Finland	−5.043*	1978	1997	4	−6.294*	1978	1997	4	−4.308‡	1990	1996	3	−4.636	1978	1990	0
France	−0.954	1978	1980	1	−1.04	1978	1982	1	−4.899+	1987	1998	3	−3.743	1975	1987	0
Greece	−6.417*	1979	1992	0	−7.291*	1979	1992	0	−4.477‡	1980	1987	0	−4.44	1975	1987	0
Italy	−4.431‡	1988	1997	4	−4.988‡	1975	1980	2	−6.203*	1975	1977	1	−4.65	1975	1979	1
Japan	−7.385*	1977	1979	1	−5.023‡	1975	1988	4	−5.276*	1987	1997	0	−5.734+	1975	1992	1
Netherlands	−8.353*	1982	1994	0	−9.059*	1977	1994	0	−4.917+	1980	2001	0	−3.703	1975	1980	1
Portugal	−7.486*	1994	1997	1	−4.641	1994	1998	4	−4.177‡	1975	1990	3	−5.842+	1975	1984	3
South Korea	−6.879*	1979	1981	1	−8.665*	1980	1998	0	−9.132*	1979	1997	3	−9.560*	1986	1993	0
Spain	−3.511	1980	1997	2	−4.36	1975	1981	0	−2.983	1986	1993	0	−3.415	1975	1992	0
Sweden	−6.435*	1977	2001	0	−7.664	1977	2001	0	−4.633+	1983	1993	0	−2.367	1991	1993	0
UK	−4.804+	1981	1991	2	−6.079*	1994	2001	2	−4.387‡	1979	1996	0	−4.36	1975	1981	0
USA	−5.160*	1975	1983	0	−6.501*	1975	1983	0	−6.997*	1977	1981	1	−6.056*	1975	1981	0

Note: To select the order of lag ρ, we start with a maximum lag length of 4 and pare it down as per the Akaike Information Criterion (AIC). There is no general rule on how to choose the maximum lag to start with. The bandwidth and maximum lag length are chosen according to the Bartlett kernel which is equal to $4(T/100)^{2/9} \approx 4$, where $T=45$ (Basher and Westerlund, 2008). The MacKinnon (1991) one-sided critical values for the ADF unit root tests with a constant and without a time are −3.64, −2.96 and −2.61 at 1%, 5% and 10% significance level respectively while those with a constant and a time trend are −4.24, −3.54 and −3.20 respectively. *,+ and ‡ denotes 1%, 5% and 10% significance level correspondingly.

Note: The MacKinnon (1991) critical values for the ADF unit root tests with a constant and without a time are −3.66, −2.96 and −2.61 at 1%, 5% and 10% significance level respectively while those with a constant and a time trend are −4.25, −3.54 and −3.21 respectively.

Note: The critical values for the ZA test, which allows for a break in the intercept only, are −5.43, −4.80 and −4.58 while those for a break in the intercept and trend are −5.57, −5.08 and −4.82 at 1%, 5% and 10% level of significance correspondingly.

Note: M1$_{B,L}$: Test equation for two breaks in the level of a trending series. M2$_{B,L}$: Test equation for two breaks in the level and slope of a trending series. T$_{B1}$ and T$_{B2}$ are the dates of the structural breaks. The one-sided critical values are −5.259, −4.514 and −4.143 respectively for model M1$_{B,L}$ and −5.949, −5.181 and −4.789 at 1%, 5% and 10% level of significance ($T=50$) for model M2$_{B,L}$.

Table 3
(a): Pair-wise correlation matrix of $\Delta LCOP_t$. (b): Pair-wise correlation matrix of $\Delta LGDP_t$.
Source: Computed.

	AUS	AUT	BEL	CAN	FIN	FRA	GRC	ITA	JPN	NLD	PRT	KRO	ESA	SWE	GRB	USA
(a)																
AUS	1.000															
AUT	−0.172	1.000														
BEL	0.077	0.117	1.000													
CAN	0.430	−0.001	0.121	1.000												
FIN	0.118	−0.035	0.112	0.250	1.000											
FRA	0.309	0.223	0.342	0.373	0.119	1.000										
GRC	0.129	0.006	0.038	0.083	0.075	0.391	1.000									
ITA	0.103	0.344	0.239	0.365	0.298	0.429	0.389	1.000								
JPN	0.211	−0.091	0.328	0.247	0.080	0.360	0.232	0.396	1.000							
NLD	0.095	−0.019	−0.050	0.157	−0.110	−0.039	−0.087	0.030	0.128	1.000						
PRT	−0.099	0.059	0.040	−0.027	0.061	−0.121	−0.065	−0.054	0.125	−0.296	1.000					
KOR	0.273	−0.249	0.089	0.225	0.092	0.011	0.304	0.023	0.475	0.195	0.128	1.000				
SPA	0.136	0.147	0.334	0.167	0.257	0.274	0.015	0.361	0.281	0.141	−0.004	−0.022	1.000			
ESP	0.367	−0.178	0.032	0.360	0.209	0.398	0.317	0.371	0.386	−0.080	0.057	0.129	0.207	1.000		
GRB	0.165	−0.018	0.182	0.480	0.088	0.480	0.032	0.391	0.316	0.060	−0.099	0.000	0.261	0.352	1.000	
USA	0.464	0.060	0.414	0.663	0.250	0.319	0.015	0.318	0.482	0.199	0.060	0.258	0.344	0.440	0.231	1.000
(b)																
AUS	1.000															
AUT	0.398	1.000														
BEL	0.476	0.751	1.000													
CAN	0.397	0.422	0.599	1.000												
FIN	0.486	0.556	0.622	0.658	1.000											
FRA	0.474	0.810	0.879	0.591	0.623	1.000										
GRC	0.291	0.480	0.502	0.423	0.406	0.546	1.000									
ITA	0.419	0.706	0.840	0.624	0.564	0.858	0.417	1.000								
JPN	0.357	0.562	0.595	0.442	0.467	0.668	0.578	0.727	1.000							
NLD	0.471	0.746	0.843	0.596	0.545	0.781	0.482	0.784	0.575	1.000						
PRT	0.390	0.715	0.721	0.381	0.394	0.771	0.455	0.735	0.574	0.652	1.000					
KOR	0.051	0.331	0.337	0.305	0.121	0.418	0.183	0.463	0.550	0.342	0.277	1.000				
SPA	0.423	0.716	0.793	0.574	0.620	0.808	0.554	0.675	0.586	0.758	0.621	0.331	1.000			
ESP	0.448	0.493	0.603	0.599	0.805	0.549	0.337	0.564	0.393	0.592	0.335	0.185	0.565	1.000		
GRB	0.264	0.400	0.429	0.634	0.623	0.488	0.433	0.452	0.374	0.468	0.440	0.344	0.535	0.582	1.000	
USA	0.143	0.337	0.413	0.805	0.442	0.437	0.415	0.467	0.417	0.487	0.312	0.331	0.412	0.391	0.727	1.000

Note: AUS, AUT, BEL, CAN, FIN, FRA, GRC, ITA, JPN, NLD, PRT, KOR, SPA, ESP and, GRB denote Australia, Austria, Belgium, Canada, Finland, France, Greece, Italy, Japan, Netherlands, Portugal, South Korea, Spain, Sweden and UK respectively.

Table 4
(a): LLC panel unit root test statistics. (b): IPS panel unit root test statistics. (c): ILT panel LM unit root test statistics. (d): Pesaran CADF panel unit root test statistics. (e): Chang-Song panel unit root test statistics.
Source: Computed.

(a)

Variables	Deterministics	Level Form		First-Difference	
		t-value	t^*	*t*-value	t^*
$LCOP_{it}$	Constant	−6.241	−2.818 [0.002]*	−25.771	−20.889 [0.000]*
	Constant+Trend	−7.463	−0.560 [0.288]	−28.514	−22.886 [0.000]*
$LGDP_{it}$	Constant	−3.622	−0.812 [0.208]	−21.071	−16.969 [0.000]*
	Constant+Trend	−7.085	−1.773 [0.038]$^+$	−22.355	−16.536 [0.000]*

(b)

Variables	Data	Deterministics	Level Form		First-Difference	
			t-bar	Ψ_t	*t*-bar	Ψ_t
$LCOP_{it}$	Raw	Constant	−1.783	−1.215 [0.112]	−5.152	−16.445 [0.000]*
		Constant+Trend	−1.459	3.466 [1.000]	−6.278	−20.463 [0.000]*
	Demeaned	Constant	−1.756	−1.094 [0.137]	−5.879	−19.721 [0.000]*
		Constant+Trend	−2.004	0.764 [0.778]	−6.709	−22.613 [0.000]*
$LGDP_{it}$	Raw	Constant	−2.372	−3.889 [0.000]*	−4.411	−13.161 [0.000]*
		Constant+Trend	−2.595	−2.077 [0.019]$^+$	−4.952	−13.846 [0.000]*
	Demeaned	Constant	−1.351	0.721 [0.765]	−5.177	−16.650 [0.000]*
		Constant+Trend	−1.917	1.290 [0.901]	−5.415	−16.157 [0.000]*

(c)

Variables	With One Break	With Two Breaks
$LCOP_{it}$	−3.803*	−6.082*
$LGDP_{it}$	−3.030*	−4.800*

(d)

Variables	Deterministics	Level Form		First-Difference	
		t-bar	*Z*	*t*-bar	*Z*
$LCOP_{it}$	Constant	−1.927	−0.668 [0.252]	−4.683	−12.395 [0.000]*
	Constant+Trend	−2.217	0.559 [0.712]	−5.416	−13.982 [0.000]*
$LGDP_{it}$	Constant	−1.550	0.937 [0.826]	−5.086	−14.112 [0.000]*
	Constant+Trend	−2.285	0.248 [0.598]	−5.205	−13.022 [0.000]*

(e)

Statistics	$LCOP_{it}$		$LGDP_{it}$	
	Level Form	First-Difference	Level Form	First-Difference
ta_c	1.104	−10.798*	−1.993	−3.837*
ta_h	−0.125	−2.532*	0.006	−0.549
ta_a	0.957	−1.045	0.443	0.099
tm_c	−1.523	−4.686*	−1.409	−2.787$^+$
tm_h	−1.129	−3.255*	−0.755	−1.287
tm_a	−2.280	−2.671‡	−2.081	−1.599

Note: The lag lengths for the panel test are based on those employed in the univariate ADF test. Assuming no cross-country correlation and *T* is the same for all countries, the normalized t^* test statistic is computed by using the *t*-value statistics. After transformation by factors provided by LLC, the t^* tests is distributed standard normal under the H_0 of non-stationarity. It is then compared to the 1%, 5% and 10% significance levels with the one-sided critical values of −2.326, −1.645 and −1.282 correspondingly. The p-values are in square brackets.
Note: The lag lengths for the panel test are based on those employed in the univariate ADF test. The IPS test statistics are computed as the average ADF statistics across the sample. These statistics are distributed as standard normal as both *N* and *T* grow large. *t*-bar is the panel test based on the ADF statistics. Critical values for the *t*-bar statistics without trend at 1%, 5% and 10% significance levels are −1.980, −1.850 and −1.780 while with inclusion of a time trend, the critical values are −2.590, −2.480 and −2.410 respectively. Assuming no cross-country correlation and *T* is the same for all countries; the normalized Ψ_t test statistic is computed by using the *t*-bar statistics. The Ψ_t tests for H_0 of joint non-stationarity and is compared to the 1%, 5% and 10% significance levels with critical values of −2.330, −1.645 and −1.282 correspondingly.
Note: The maximum lag length is based on the Bartlett kernel. Critical values for the LM panel unit root test (without or with breaks) are distributed asymptotic standard normal and are −2.326, −1.645, and −1.282 at the 1%, 5%, and 10% levels, respectively. The minimum LM unit root test which accounts for a break in the data is employed to test for the H_0 of non-stationarity. Time dummies are included when performing the panel unit root test in the presence of one structural break.
Note: The lag lengths for the panel test are based on those employed in the univariate ADF test. The Pesaran CADF test of the H_0 of non-stationarity is based on the mean of individual DF (or ADF) *t*-statistics of each unit in the panel. Critical values for the *t*-bar statistics without and with trend at 1%, 5% and 10% significance levels are −2.360, −2.200 and −2.110; and −2.850, −2.710 and −2.630 respectively. Assuming cross-section dependence and *T* is the same for all countries. The normalized *Z* test statistic is computed by using the *t*-bar statistics. The *Z* test statistic is compared to the 1%, 5% and 10% significance levels with the one-sided critical values of −2.326, −1.645 and −1.282 correspondingly.
Note: The maximum lag length is based on the Bartlett kernel. The nonlinear IV average and minimum tests are denoted by the *ta* and *tm* while the subscripts *c*, *h* and *a* refer to those tests with single IGF and no covariate, with single IGF and covariate and orthogonal IGF with no covariate respectively. As per Chang and Song (2009), the tests include a constant term only. The H_0 of non-stationarity is tested. Each test statistic is compared to the 1%, 5% and 10% significance levels with the one-sided critical values of −2.326, −1.645 and −1.282 for the average test while these are −3.243, −2.746 and −2.502 for minimum test (N=17) respectively.

Table 5
(a): Pedroni panel cointegration test statistics. (b): Westerlund cointegration test statistics. (c): Westerlund and Edgerton panel cointegration test statistics. (d): Di Iorio and Fachin Panel cointegration test with breaks. (e): Di Iorio and Fachin Panel cointegration test with long-run cross-section dependence.
Source: Computed.

(a)

Statistics	Without trend	With trend
Panel ν-statistic	2.600*	−0.357
Panel ρ-statistic	−1.048	0.969
Panel pp-statistic	−0.556	1.022
Panel adf-statistic	0.038	0.941
Group ρ-statistic	0.285	1.260
Group pp-statistic	0.563	1.181
Group adf-statistic	1.087	1.245

(b)

Statistics	Without trend				With trend			
	Value	*Z*	*P*-value	Robust *P*-value	Value	*Z*	*P*-value	Robust *P*-value
Gt	−1.213	−0.910	0.181	0.143	−1.339	5.060	1.000	0.999
Ga	−2.214	1.397	0.919	0.736	−4.120	4.642	1.000	1.000
Pt	−5.755	−3.181	0.001*	0.143	−6.761	1.930	0.973	0.779
Pa	−3.613	−3.573	0.000*	0.201	−6.469	1.607	0.946	0.804

(c)

Statistics	No break	Level break	Regime shift
Z_{τ}(N)	2.393 [0.992]	3.014 [0.999]	1.900 [0.971]
Z_{ϕ}(N)	2.086 [0.981]	2.665 [0.996]	1.619 [0.947]

(d)

Statistics	Without trend	With trend
Median ADF	−5.981 [0.000]*	−6.778 [0.000]*
Mean ADF	−5.771 [0.000]*	−6.810 [0.000]*
Maximum ADF	−3.539 [0.000]*	−4.519 [0.000]*

(e)

Statistics	Without trend	With trend
Median ADF	−6.691 [0.000]*	−6.068 [0.000]*
Mean ADF	−6.137 [0.000]*	−6.123 [0.000]*
Maximum ADF	−2.260 [0.000]*	−3.730 [0.000]*

Note: The panel statistics are the within-dimension statistics while group statistics are between-dimension ones. These are one-sided standard normal test with critical values of 1%, 5% and 10% given by −2.326, −1.645 and −1.282. A special case is the panel ν-statistic which diverges to positive infinity under the alternative hypothesis. Rejection of the H_0 of no cointegration requires values larger than 2.326, 1.645 and 1.282 at 1%, 5% and 10% significance level. The critical values for the mean and variance of each statistic are obtained from Pedroni (1999). H_0 corresponds to no cointegration.
Note: All of these statistics are distributed standard normal. Critical values of one-sided tests for 1%, 5% and 10% significance levels are −2.326, −1.645 and −1.282, respectively. The lag and lead lengths are set to one. Choosing too many lags and leads can result in a deterioration of the small-sample properties of the test. To control for cross-sectional dependence, robust critical values are obtained through 5000 bootstrap replications.
The maximum lag length is based on the Bartlett kernel. The H_0 of no cointegration is tested. The statistics test is distributed as a one sided standard normal with critical values of one-sided tests for 1%, 5% and 10% significance levels are −2.326, −1.645 and −1.282 respectively.
Note: The maximum lag length is based on the Bartlett kernel. The panel statistics are compared to one-sided standard normal test with critical values of 1%, 5% and 10% given by −2.326, −1.645 and −1.282. The *p*-values are obtained through 5000 bootstrap replications. The breakpoint is estimated on the basis of the residual sum of squares.

Korea, Spain, Sweden, UK and for the overall panel. At individual level, unidirectional causality running from economic growth to copper consumption is detected for South Korea, Sweden, UK and USA in the short-run and for Finland, France and UK in the long-run. Unidirectional causality is also found running from copper consumption to economic growth for Australia, Finland and Italy in the short-run and for Spain in the long-run. Bi-directionality between economic growth and copper consumption is found for Belgium, Greece, Italy, Japan and South Korea in the long-run. The neutrality hypothesis holds for Australia, Austria, Canada, Netherlands, Portugal, Sweden and USA in the long-run.

With regard to the overall panel, the system GMM estimation reveals a unidirectional causality running from economic growth to copper consumption in the long-run. The absence of feedback effects in the long-run reveals the importance of growth of real GDP in stimulating the metal consumption. These results are consistent with those of Ghosh (2006), Huh (2011) and Jaunky (2012).

Conclusion and policy implications

The paper examines the relationship between copper consumption and economic growth for 16 rich countries over the period of 1966–2010. Three generations of panel unit root and cointegration tests are applied. Both series are found to be I(1) cointegrated especially after controlling for cross-sectional cointegration. Time-series and panel causality tests are both performed. Unidirectional causality running from economic

Table 6
VECM-based causality test.
Source: Computed.

Country	ρ	*LGDP* ⇒ *LCOP*			*LCOP* ⇒ *LGDP*		
		Break	Short-Run Causality	Long-Run Causality	Break	Short-Run Causality	Long-Run Causality
Australia	1	0.58 [0.451]	0.93 [0.341]	2.04 [0.162]	0.94 [0.339]	7.04 [0.012]+	0.87 [0.357]
Austria	2	0.75 [0.393]	2.07 [0.141]	0.41 [0.524]	0.06 [0.801]	0.96 [0.391]	0.47 [0.496]
Belgium	1	0.15 [0.697]	1.02 [0.320]	10.21 [0.003]*	0.93 [0.340]	0.06 [0.802]	10.53 [0.003]*
Canada	1	1.99 [0.167]	0.13 [0.718]	1.28 [0.264]	0.03 [0.867]	0.07 [0.795]	0.91 [0.345]
Finland	1	0.64 [0.427]	0.02 [0.890]	4.24 [0.046]+	1.55 [0.221]	2.88 [0.098]‡	0.63 [0.432]
France	1	0.73 [0.397]	0.47 [0.497]	3.93 [0.055]‡	0.06 [0.814]	0.00 [0.981]	2.57 [0.118]
Greece	1	1.22 [0.276]	0.02 [0.881]	7.98 [0.008]*	7.90 [0.001]*	0.08 [0.784]	2.86 [0.099]‡
Italy	2	0.24 [0.624]	0.86 [0.431]	3.29 [0.078]‡	0.89 [0.351]	5.01[0.012]+	25.63 [0.000]*
Japan	1	2.01 [0.164]	0.03 [0.871]	3.01 [0.091]‡	1.15 [0.289]	0.37 [0.541]	12.38 [0.001]*
Netherlands	2	1.67 [0.204]	0.10 [0.902]	0.13 [0.724]	0.02 [0.902]	1.60 [0.215]	0.51 [0.481]
Portugal	1	2.64 [0.113]	0.07 [0.789]	0.02 [0.879]	0.44 [0.513]	0.31 [0.578]	1.42 [0.241]
South Korea	2	0.79 [0.380]	4.25 [0.022]+	16.64 [0.000]*	4.01 [0.053]‡	0.25 [0.776]	2.88 [0.099]‡
Spain	1	11.40 [0.002]*	0.71 [0.405]	1.64 [0.208]	10.86 [0.002]*	0.09 [0.767]	4.74 [0.036]+
Sweden	1	1.52 [0.224]	3.08 [0.087]‡	0.71 [0.404]	4.81 [0.035]+	0.00 [0.969]	0.01 [0.915]
UK	3	0.58 [0.452]	2.56 [0.073]‡	12.19 [0.001]+	3.86 [0.058]‡	0.34 [0.796]	0.78 [0.394]
USA	1	1.76 [0.192]	3.30 [0.032]+	1.65 [0.207]	0.04 [0.836]	2.23 [0.144]	1.68 [0.202]
Panel	1	0.57 [0.451]	0.03 [0.865]	6.14 [0.013]+	5.23 [0.022]+	1.12 [0.290]	1.70 [0.193]

Note: *LGDP* ⇒ *LCOP* and *LCOP* ⇒ *LGDP* denote the causality tests of whether economic growth Granger-causes copper consumption and vice-versa respectively. The null of no causation is tested. The order of lags is determined according to the AIC. The F-statistics and χ^2-statistics are computed for the time-series and panel causality tests correspondingly. The statistics are computed via the Wald test. Robust two-step GMM estimations are obtained by making use of the finite-sample correction to the covariance matrix as derived by Windmeijer (2005). The explanatory variables are assumed to be endogenous and their lags are instrumented in GMM-style (Roodman, 2006). The number of instruments is set to 15. No AR(2) correlation are to be found while both the Sarjan and Hansen tests do not reject to the null of valid instruments.

growth to copper consumption is detected for South Korea, Sweden, UK and USA in the short-run and for Finland, France and UK in the long-run. Unidirectional causality is found running from copper consumption to economic growth for Australia, Finland and Italy in the short-run and for Spain in the long-run. Bi-directionality between economic growth and copper consumption prevails for Belgium, Greece, Italy, Japan and South Korea in the long-run. No causality is found for Australia, Austria, Canada, Netherlands, Portugal, Sweden and USA in the long-run. Long-run panel unidirectional causality running from economic growth to copper consumption is also uncovered.

These findings have important implications for policymakers in assisting them to make long-run projections of copper demand and growth. In line with the panel causality result, economic growth can be used to predict copper consumption but not vice versa in the long-run. At individual levels, the same result is obtained for Finland, France and UK in the long-run. As such, mineral conservation policies could be implemented by these countries, without much concern of a long-run impact on economic growth. Copper consumption is not a good indicator for predicting future long-run economic growth outcomes for Australia, Austria, Canada, Netherlands, Portugal, Sweden and USA and the reverse also holds true. Likewise, similar impact of those policies on economic growth can be expected. However, copper consumption can be used to predict economic growth in Spain, but the opposite is not true. In this case, mineral conservation policies can affect the Spanish long-run economic growth rate. Furthermore, in the case of Belgium, Greece, Italy, Japan and South Korea, economic growth can be used to forecast copper consumption and vice versa. As a result, mineral conservation policies could well impede economic growth in those countries. The impact of those policies may be more pronounced for those economies whose growth is based on manufacturing such as Japan and South Korea.

In sum, economic growth will continue to affect the demand for copper for the rich countries even in the far future. Despite the availability of a wide variety of copper substitutes, the copper industry has a vital role to play in the economic sphere of those countries. For instance, in case policymakers forecast a rise in economic growth, it is imperative they look at their mineral policies to ensure an adequate and efficient supply of copper in order to sustain their long-term development goals.

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