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FOREST PRODUCTS EXPORTS AND ECONOMIC GROWTH: EVIDENCE FROM RICH COUNTRIES

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ABSTRACT

The forest sector contributes a significant share to national income. The existence and magnitude of causal relationships between forest product exports and economic growth is thus important to understand not least for policy issues. It has vital implications for policy-makers enacting proper development strategies. This causality is usually analyzed using the export-led economic growth hypothesis. International trade is affecting economic growth through enhanced competition and specialization. Export, more specifically, foster economic growth via the accumulation of foreign exchange, by stimulating efficient investments in the right sectors and by allowing for improved economies of scale. Surprisingly, practically no studies have been done analyzing the forest products export-led economic growth hypothesis. Thus, the current study fills an important gap in the literature. The study attempts to test the forest product export-led growth hypothesis for 22 economies over the period 1970 to 2011. Various generations of panel unit root and cointegration tests are applied. The time frame and the selection of countries are purely dictated by the availability of data and the amount of existing productive forest area. The econometric tests are based on augmented Dickey-Fuller unit root and cointegration tests. These tests are necessary before assessing the impact of forest product exports on GDP. The connection between economic growth and forest products exportation is analyzed using an error correction model (ECM) based panel causality test structure. The ECM is subsequently used to estimate short- and long-run elasticities. The series are found to be integrated of order one and cointegrated, especially when applying the third-generation tests. Uni-directional causality running from forest product exports to economic growth is uncovered in the both the short-run and the long-run. Moreover, forest products exportation is found to positively affect economic growth. The short-run elasticity reveals positive and significant income elasticity. A 1% increase in forest product exports will lead to a 0.022% increase in economic growth in the short-run and 0.002% in the long-run. The regional dummy is also significant and positive, implying that countries with significant forest land coverage are bound to experience higher economic growth. The findings will help policymakers in their projections and implementing natural resource and forest policies. Unidirectional causality implies forest product exports can be used to predict economic growth in both short-run and long-run but not vice versa. In general, the results support the ELG hypothesis. Promotion of forest product exportations can lead to a multiplier effect.

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INTRODUCTION

Forest resources have been a major source of economic development for many states. They do not only cater for wood, wild foods, medicines, soil conservation, carbon dioxide storage, and landscape beauty but additionally contribute in stimulating foreign exchange earnings, employment and economic growth. Forests indeed epitomize a productive asset which can be employed as a means for attaining national development goals, including equity, stability, investment and growth (Food and Agricultural Organization, FAO 2005). According to the FAO (2009), the forest sector contributes about US$ 468 billion to national income, representing about 1% of global GDP in 2006. The question of whether there is a causal relationship between forest product exports and economic growth has vital implications for policy-makers in enacting proper development strategies.

This paper presents the first study of the link between forest products exports and economic growth using panel data from 22 rich countries over the period of 1970-2011. The export-led growth (ELG) can be employed to investigate whether a particular sector such as the forestry has contributed significantly to the economic growth for those rich economies. 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. Forest product exports are found to Granger-cause economic growth. In addition, these have a positive impact on economic growth. Overall, these findings lend support to the ELG hypothesis for the rich economies. It is essential for these economics to preserve their forest and sustain their forestry.

REVIEW OF LITERATURE

The theoretical foundation of the export-led growth (ELG) hypothesis debate goes back to the pioneering works of the classical economists such as Adam Smith (1776) and David Ricardo (1817). They demonstrate the crucial role of international trade on economic growth and the economic gains through enhanced competition and specialization according to comparative advantage. Kernal *et al*. (2002) put forward several arguments in which exports can foster economic growth. First, accumulation of foreign exchange allows the possibility of high-tech imports which could enlarge production possibilities. Second, exports can cause investments to be concentrated in the most efficient sectors. Third, the association of the international and domestic markets allows for greater scope of economies of scale. Finally, greater trade can lower allocative inefficiencies through enhanced competition.

Giles and Williams (2000) and Medina-Smith (2001) provide a thorough review of the economic growth-exports. The empirical literature in connection to the ELG paradigm can be segmented into three groups such as: (i) cross-sectional (e.g. Lussier, 1993) (ii) country-specific time-series (e.g. Siddique and Selvanathan, 1999) and (iii) panel data (e.g. Parida and Sahoo, 2007) studies. Practically no studies have been done to analyze the impact of forest trade on economic development. Econometric tests such as cointegration and causality tests have been used quite scantly in the forest economics literature.

Practically no studies have been done explicitly on the forest products export-led economic growth hypothesis. On the contrary, some studies have focused on export-led production growth hypothesis. For instance, Chao and Buongiomo (1999) examine the hypothesis of 15 major countries of wood pulp. They found weak supporting evidence in both the short-run and long-run. Chao and Buongiomo (2002) study the export-led production growth hypothesis for pulp and paper for 15 major exporting countries and find feedback effects between exports and production. These studies make use of Granger-causality test. Oluwatoyese and Applanaidu (2014) analyze the effects of the agricultural sector on economic growth for Nigeria. Using time-series analysis, they find no evidence of a significant contribution of forestry. In general, those studies fail to account for structural breaks in the data and also cross-sectional dependence. Ignoring these effects can lead to erroneous conclusions.

DATA AND METHODOLOGY

Data

Forest data are obtained from the FAO of the United Nations website and those of real gross domestic product (GDP, at constant 2000) are compiled from the 2012 World Development Indicators of The World Bank. The time frame and the selection of countries are purely dictated by the availability of data and the amount of existing forest area. The 22 countries are: Australia (AUS), Austria (AUT), Canada (CAN), Denmark (DEN), Finland (FIN), France (FRA), Germany (DEU), Greece (GRC), Hungary (HUN), Ireland (IRL), Israel (ISR), Italy (ITA), Japan (JPN), Netherlands (NLD), Norway (NOR), Portugal (PRT), South Korea (KOR), Spain (ESP), Sweden (SWE), Trinidad & Tobago (TTO), UK (GBR) and United States (USA).

Methodology

To investigate whether forest products ELG hypothesis, the following basic reduced-form equation can be estimated (Siddique and Selvanathan, 1999):

$$LGDP_{it} = g_0 + g_1 LFOR_{it} + \varepsilon_{it} \qquad (1)$$

where $LGDP_{it}$ captures economic growth and denotes the natural logarithm of GDP (at constant 2000) for country *i* and year *t*. $LFOR_{it}$ denotes the natural logarithm of forest products exportation (at constant 2000) for country *i* over year *t*. Finally, g_0 is the constant term and ε_{it} represents the error term.

The coefficient g_1 illustrates the responsiveness of income to changes in forest products exportation. A statistically significant positive value provides support to the export-led growth hypothesis. However, it is also important to analyse the magnitude of the estimated coefficient to understand the economic significance of the hypothesis.

Table 1 shows the mean statistics of $LGPD_{it}$ and $LFOR_{it}$ over the period 1970-2011 as well as some country characteristics relating to the forest sector. The share of forest product exports to GDP is rather relatively large for Canada, Denmark, Finland and

Sweden. All but Denmark is relatively with forest resources which might explain part of the large share of forest product exports to GDP.

TABLE 1: COUNTRY STATISTICS

Country	Forest Area (sq. km in 2010)	Forest Area (% of land area in 2010)	Forest Products Export Dependency (%)	*LGDP* (Mean 1970-2011)	*LFOR* (Mean 1970-2011)
AUS	1,493,000	19.434	0.156	26.429	19.887
AUT	38,870	47.155	1.800	25.714	21.787
CAN	3,101,340	34.105	1.308	27.005	23.510
DEN	5,440	12.821	0.197	25.556	19.607
FIN	221,570	72.909	5.354	25.254	22.722
FRA	159,540	29.131	0.298	27.677	21.832
DEU	110,760	31.772	0.608	28.035	22.447
GRC	39,030	30.279	0.044	25.366	17.889
HUN	20,290	22.412	0.774	24.478	19.269
IRL	7,390	10.727	0.241	24.729	18.796
ISR	1,540	7.116	0.028	25.057	17.443
ITA	91,490	31.104	0.225	27.498	21.201
JPN	249,790	68.529	0.055	28.908	21.358
NLD	3,650	10.821	0.618	26.376	21.248
NOR	100,650	32.949	0.405	25.511	21.020
PRT	34,560	37.783	0.965	25.129	20.681
KOR	62,220	64.078	0.235	26.323	20.697
ESP	181,730	36.433	0.340	26.802	20.621
SWE	282,030	68.731	3.214	26.024	22.821
TTO	2,260	44.055	0.055	22.728	14.248
GBR	28,810	11.908	0.123	27.774	21.233
USA	3,040,220	33.236	0.174	29.583	23.269

Note: Forest products export dependency is measured as real forest products exports as a percentage of real GDP in 2011. Forest area data are compiled from the World Development Indicators 2012.

Unit Root and Cointegration Tests

Econometric tests such as unit root and cointegration tests are necessary before assessing the impact of forest product exports on GDP. Most of the unit root tests are based on an augmented Dickey-Fuller (ADF) unit root test type. Let *y* denotes a variable, thus:

$$\Delta y_{it} = r_i + \kappa_i t + \delta y_{it-1} + Z_{im}\sum_{j=1}^{\rho_i}\Delta y_{it-j} + e_{it} \qquad (2)$$

where $\Delta y_{it} = y_{it} - y_{it-1}$, *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: $\hat{\delta} = 0$), then the series can be considered to be non-stationary. y_{it} is integrated of order of *d*, i.e. $y_{it} \sim I(d)$, if it were to be differenced by *d* times to become stationary. Initially, time series unit root tests such as the ADF, Zivot and Andrews (1992) and Narayan and Popp (2010) will be first computed for each country. Several generations of panel unit root tests are next employed. First generation panel unit root tests include those of Levin, Lin and Chu (2002, LLC), Im, Pesaran, and Shin (2003, IPS) and Im, Lee and Tieslau (ILT, 2005). But, these tests assume independence of individual cross-sections and this is very 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 test, Chang and Song (2009) propose a test which tackles any forms of dependence whether short-run or long-run.

Unit root tests are commonly computed via two different regressions. One regression includes a constant term only and one which includes both a constant term and a time trend. Time-series data tend to be non-stationary and display a trend over time. In such cases, it is more suitable to apply 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 then 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, several panel cointegration tests can be performed. Pedroni (1999, 2004) was among the firsts to propose testing the panel cointegration. Such first generation panel cointegration test assumes cross-sectional independence across individuals. With regard to second generation panel cointegration tests, Westerlund (2008) and Westerlund and Edgerton (2008) suggest some panel cointegration tests which can effectively deal with cross-sectional dependence. The latter also allows for unknown structural breaks in both the intercept and slope of the panel cointegrating regression, which can be located at different periods for different countries. A third generation panel cointegration test which is robust to short-run and long-run dependence across countries is devised by Di Iorio and Fachin (2012). As a whole, none of the above unit root and cointegration tests is devoid from statistical shortcoming in terms of size and power properties. It is consequently more convenient to apply various tests in order to come to a conclusion about the properties of the series.

Panel Causality Test

The connection between economic growth and forest products exportation can consequently be investigated. An ECM-based panel causality test structure is as follows:

$$\Delta LGDP_{it} = \varpi_{it} + \sum_{q=1}^{p} \varphi_q \Delta LFOR_{it-k} + \sum_{q=1}^{p} \psi_q \Delta LGDP_{it-k} + \lambda_{2i} ECM_{it-1} + f_{2i} + \upsilon_{it} \tag{3a}$$

$$\Delta LFOR_{it} = \omega_{it} + \sum_{q=1}^{p} \theta_q \Delta LFOR_{it-k} + \sum_{q=1}^{p} \phi_q \Delta LGDP_{it-k} + \lambda_{1i} ECM_{it-1} + f_{1i} + \mu_{it} \tag{3b}$$

where $i = 1, ..., N$, $t = 1, …., T$, Δ denotes first differences, ω_{it} and ϖ_{it} are the intercept terms, f_{1i} and f_{2i} are the fixed effects components while θ_{it}, ϕ_{it}, φ_{it}, ψ_{it},, λ_{1i} and λ_{2i} are that parameters which are required to be estimated. The ECM_{it-1} variable represents the error-correction term and is lagged by one period. It is derived from the cointegrating vector of equation (1) and the error terms μ_{it} and υ_{it} are independent and identically distributed (*i.i.d.*).

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 specifying whether previous changes in one variable contribute significantly to the prediction of the future value of the other variable. Per se, forest products exportation does not Granger-cause economic growth if and only if all of the coefficients $\varphi_q; \forall =1, \ldots, p$ are not significantly different from zero in equation (3a). The dependent variable reacts only to short-term shocks. Likewise, economic growth does not Granger-cause forest product exportation in the short run if and if all of the coefficients $\phi_q; \forall =1, \ldots, p$ are not significantly different from zero. These can be referred to as the "*short-run Granger causality*" tests. The coefficients on the *ECM*s represent how fast deviations from the long- run equilibrium are eliminated. Another channel of causality can be explored by testing the significance of the *ECM*s. This test can be denoted as the "*long-run Granger causality*" tests.

Conventional ordinary least squares (OLS), fixed-effects or random-effects models tend to yield biased results due to the correlation between the lagged dependent variables and the error terms. To remedy the correlation and endogeneity problems, Blundell and Bond (1998) suggest a two-step system GMM estimator which has superior finite-sample properties. Robust estimates can be obtained by making use of the finite-sample correction to the two-step covariance matrix as derived by Windmeijer (2005). For instruments to be valid there should not be serial correlation in μ_{it} and υ_{it}. The optimal lag length, ρ, is selected when no serial correlation is obtained in the residuals.

The direction of causality between economic growth and forest product exportations has significant policy implications. If there is no causality, then adopting a conservative resource policy measures to limit the exportation of forest products can be implemented, without the concern of negatively impacting on economic growth. This can eventually cause a reduction in the exploitation of natural resources and environmental degradation. If causality runs from economic growth to forest products exportation, environmental and resource policies can still be implemented. For instance, environmental taxes and tariffs can be imposed on the forest industries. These policies will have no impact on economic growth. However, if a unidirectional causality running from forest product exportations to economic growth exists, then resource conservation policies will adversely affect the growth rate of the economy.

Short-Run and Long-Run Elasticities

The short-run elasticity is estimated by running the following reduced-form regression:

$$\Delta LGDP_{it} = \beta_1 \Delta LFOR_{it} + \lambda ECM_{it-1} + \beta_2 D_{itk} + \varepsilon_{it} \tag{4a}$$

where ε_{it} is the error terms, β_1 is the short-run elasticity and similar to the ϕ's in equation (4), λ measures the speed of adjustment towards the long-run equilibrium. The Engle and Granger (1987) ECM model is augmented by a dummy variable D_{itk} which is defined as:

$$D_{itk} = \begin{cases} 0 \; if \;\; t \le k \\ 1 \; if \;\; t > k \end{cases} \tag{4b}$$

where k denotes the point at which a country's forest area is above its regional average. That is, regional averages are calculated for five continents (Africa, Asia including Oceania, Europe and South and North America) and the dummy variable is set to unity for countries which has as a larger forest area than their corresponding regional average. Countries for D_{itk} = 1 are Austria, Canada, Finland, Japan, South Korea, Sweden and USA. In effect, while the regional dummy variable accounts for any shifts in the dependent variable, it also offers a means to minimize any misspecification bias which could arise during the calculation of the short-run elasticities.

Long-run estimates can be computed via the dynamic OLS (DOLS) panel technique. which control for both endogenous and serially correlated regressors. The long-run regression is augmented by lead and lagged difference of the dependent and explanatory variables to control for serial correlation and endogenous feedback effects. The within-dimension-based DOLS model as per Kao and Chiang (2000) can be represented as follows:

$$LGDP_{it} = \alpha_i + \beta LFOR_{it} + \sum_{q=-\ell}^{\ell} \gamma_j \Delta LFOR_{it-q} + \zeta_{it} \qquad (5)$$

where, α_i denotes the individual fixed effects, β is the homogenenous coefficient across the rich countries, ℓ is the number of leads and lads for the first differenced of the $LGDP_{it}$ series and ζ_{it} is the error terms for country *i* and year *t*.

EMPIRICAL RESULTS

The ADF, Zivot-Andrews and Narayan-Popp unit root tests are performed both at levels and at first differences. The individual country results are not presented in the study but are available upon request. For the ADF unit root statistics for individual countries the results suggest that the order of integration, both the $LCOP_t$ and $LGDP_t$ series are found to be I(1) for Canada, Denmark, France, Netherlands, Norway, Portugal, Singapore, South Korea and USA. The ADF test ignores the presence of structural breaks in the series. As per Perron (1989), this can lead to the unit root test to be biased towards the non-rejection of the null hypothesis. The Zivot-Andrews test can control for one endogenous structural break in the series. Both its series are found to be I(1) for Canada, France, Germany, Greece, Hungary, Netherlands, Singapore, South Korea, UK and USA. The Narayan-Popp test can account for the presence of two endogenous structural breaks and adds more power to the testing framework. 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 series for Austria and Ireland only while no series are found to simultaneously to follow this process when computing the $M2_{B,L}$ test.

On average, when referring to the Narayan-Popp model, the first break in either series tends to fall around end 1970s to mid 1980s. These periods coincide with the 1974-1975 and 1980-1981 oil price shocks, following the Yom Kippur War and Iranian Revolution respectively. The second break tends to occur around the early 1990's and 2000's. These periods once again match oil price shocks following the Gulf War in 1990, the 1997 Asian Financial crisis and the 2000-2001 international recessions respective to the Middle East tensions owing to the Second Intifada. Moreover, the habitat conversation rules in the early 1990s in Pacific Northwest did cause supply shock in the international wood product markets (Perez-Garcia and Barr, 2005).

Toda (1995) issues a caveat about the poor performance and low power of time-series tests even in the presence of 100 observations. These tests could produce spurious results. Therefore this raises need to exploit panel data techniques. Table 2(a) reports the LLC test statistics for both series, where $LGDP_{it}$ ~ I(1) and $LFOR_{it}$ ~ I(1). The assumptions of the LLC test are quite restrictive. The test ignores the presence structural breaks and assumes cross-sectional independence. Cross-sectional dependence can bias the panel data unit root tests towards the alternative hypothesis (Banerjee et al., 2004).

TABLE 2(A): LLC PANEL UNIT ROOT TEST STATISTICS

Variable	Deterministics	Level Form		First-Difference	
		***t*-value**	***t****	***t*-value**	***t****
$LGDP_{it}$	Constant	-3.946	-0.489 [0.313]	-20.103	-13.955 [0.000]*
	Constant + Trend	-8.456	-0.317 [0.376]	-20.924	-12.217 [0.000]*
$LFOR_{it}$	Constant	-7.494	-1.625 [0.052]‡	-26.446	-21.299 [0.000]*
	Constant + Trend	-20.924	-12.217 [0.000]*	-25.771	-16.759 [0.000]*

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.*

The IPS test controls for heterogeneity between groups and control for cross-sectional dependence using demeaned data. As reported in Table 2(b), the IPS test provides similar processes to the LLC test for the $LGDP_{it}$ and $LFOR_{it}$ series.

TABLE 2(B): IPS PANEL UNIT ROOT TEST STATISTICS

Variable	Data	Deterministics	Level Form		First-Difference	
			***t*-bar**	**Ψ_t**	***t*-bar**	**Ψ_t**
$LGDP_{it}$	Raw	Constant	-1.636	-0.650 [0.258]	-4.183	-14.212 [0.000]*
		Constant + Trend	-2.140	0.166 [0.566]	-4.409	-13.035 [0.000]*
	Demeaned	Constant	-1.249	1.402 [0.920]	-4.294	-14.802 [0.000]*
		Constant + Trend	-2.047	0.709 [0.761]	-4.332	-12.584 [0.000]*
$LFOR_{it}$	Raw	Constant	-1.848	-1.813 [0.035]†	-6.025	-23.969 [0.000]*
		Constant + Trend	-2.743	-3.342 [0.000]*	-5.594	-19.836 [0.000]*
	Demeaned	Constant	-1.802	-1.566 [0.059]‡	-5.682	-22.151 [0.000]*
		Constant + Trend	-2.423	-1.478 [0.070]‡	-5.249	-17.846 [0.000]*

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 -1.930, -1.810 and -1.750 while with inclusion of a time trend, the critical values are -2.550, -2.440 and -2.380 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.

Referring to Table 2(c), the ILT test which can specifically account for endogenous breaks reveals a stationary process for both series.

TABLE 2(C): ILT PANEL LM UNIT ROOT TEST STATISTICS

Variable	With One Break	With Two Breaks
$LGDP_{it}$	-5.223*	-6.164*
$LFOR_{it}$	-11.241*	-14.078*

Notes: 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.

As a second generation test, following Pesaran (2007), 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). As can be deduced from Table 2(d), both $LCOP_{it}$ and $LGDP_{it}$ series follow an I(1) process.

TABLE 2(D): PESARAN CADF PANEL UNIT ROOT TEST STATISTICS

Variable	Deterministics	Level Form		First-Difference	
		***t*-bar**	***Z***	***t*-bar**	***Z***
$LGDP_{it}$	Constant	-1.710	0.298 [0.617]	-4.319	-12.720 [0.000]*
	Constant + Trend	-2.060	1.492 [0.932]	-4.376	-10.851 [0.000]*
$LFOR_{it}$	Constant	-1.959	-0.944 [0.173]	-5.011	-16.171 [0.000]*
	Constant + Trend	-2.349	-0.049 [0.480]	-4.758	-12.886 [0.000]*

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 -2.300, -2.160 and -2.080; and -2.780, -2.650 and -2.580 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.

A third generation panel unit test accounts for cross-sectional cointegration which arises when there are short-run and long-run co-movements among the variables. Chang and Song (2009) suggest a test which employs of a set of orthogonal functions as instrument generating function (IGF) for controlling any long-run dependence. As illustrated in Table 2(e), the tm_c panel statistic confirm an an I(1) process for $LGDP_{it}$ while with the exception of tm_c, all test statistics confirm similar process for $LFOR_{it}$.

TABLE 2(E): CHANG-SONG PANEL UNIT ROOT TEST STATISTICS

Statistics	$LGDP_{it}$		$LFOR_{it}$	
	Level Form	**First-Difference**	**Level Form**	**First-Difference**
ta_c	-2.243$^+$	-5.066*	-0.056	-11.330*
ta_h	1.112	-0.800	-1.030	-3.123*
ta_a	-0.333	-0.646	-0.010	-2.542*
tm_c	-1.219	-3.057$^+$	-3.042$^+$	-4.515*
tm_h	-0.901	-1.954	-1.058	-4.285*
tm_a	-0.790	-2.328	-1.379	-3.689*

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. 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 .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 -3.351, -2.870 and -2.635 for minimum test (N=25) respectively.

In general, the panel unit root tests give support to the a-priori expectation of an I(1) process. Panel cointegration tests are subsequently computed. As presented in Table 3(a), only the panel-ν statistic with trend rejects the null. Weak evidence of a cointegrating relationship is found. Second-generation panel cointegration tests which can effectively deal with cross-sectional dependence are next considered.

TABLE 3(A): PEDRONI PANEL COINTEGRATION TEST STATISTICS

Statistics	Without Trend	With Trend
Panel ν-statistic	-3.209	15.352*
Panel ρ-statistic	0.553	1.662
Panel pp-statistic	-0.297	1.215
Panel adf-statistic	0.289	0.643
Group ρ-statistic	-0.134	2.260
Group pp-statistic	-0.995	1.365
Group adf-statistic	-0.160	0.100

Note: The panel and group statistics are within-dimension statistics while group statistics are between-dimension ones. Critical values of one-sided tests for 1%, 5% and 10% significance levels are -2.326, -1.645 and -1.282 respectively. The panel ν-statistic is compared to 2.326, 1.645 and 1.282 at 1%, 5% and 10% significance level respectively. The H_0 of no cointegration is tested.

Based on the Durbin-Hausman principle, Westerlund (2008) puts forward two sets of tests such as the DH_g and DH_p which are robust against the presence of stationary regressors. As reveals by DH_g statistics in Table 3(b), the variables are found to be cointegrated.

TABLE 3(B): WESTERLUND PANEL COINTEGRATION TEST

Statistics	Value
DHg	1.629‡
DHp	-0.285

Note: All these statistics are distributed standard normally. The H_0 of no cointegration is tested. Critical values of one-sided tests for 1%, 5% and 10% significance levels are 2.326, 1.645 and 1.282 respectively.

Westerlund and Edgerton (2008) cointegration tests are reported in Table 3(c). $Z\tau(N)$ and $Z\phi(N)$ and cointegration tests also allow for unknown structural breaks in both the intercept and slope of the cointegrating regression, which may be located at different dates for different countries. The $Z\phi(N)$ test statistic confirm a cointegrating relationship when controlling for breaks in both the intercept and slope. As a third-generation panel cointegration test, Di Iorio and Fachin (2012) devise some residual-based Stationary Bootstrap (RSB) tests which are robust to short- and long-run dependence across countries.

TABLE 3(C): WESTERLUND AND EDGERTON PANEL COINTEGRATION TEST STATISTICS

Statistics	Without Trend	With Trend
$Z_\tau(N)$	-0.844 [0.199]	-1.249 [0.106]
$Z_\phi(N)$	-0.963 [0.168]	-1.678 [0.047]+

Note: The maximum lag length is based on the Bartlett kernel. The trimming parameter is set to 0.25. 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.

As shown in Table 3(d), all the mean, median and maximum ADF test statistics reject the H0 of no cointegration. In essence, the panel data techniques supply ample evidence of an I(1) process for both $LGDP_{it}$ and $LFOR_{it}$ series which are also found to be cointegrated.

TABLE 3(D): DI IORIO AND FACHIN BOOTSTRAP PANEL COINTEGRATION TEST WITH LONG-RUN CROSS-SECTION DEPENDENCE

Statistics	Without Trend	With Trend
Median ADF	-6.302 [0.000]*	-6.126 [0.000]*
Mean ADF	-6.545 [0.000]*	-5.823 [0.000]*
Maximum ADF	-3.874 [0.000]*	-3.317 [0.000]*

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.

Moreover, if the $LGDP_{it}$ and $LFOR_{it}$ series are cointegrated, then causality should run in at least one direction (Engle and Granger, 1987). As presented in Table 4, the computed Hansen test statistics cannot reject the H0 of valid instruments. The number of instruments used is equal to 12. No second order serial correlation is also established. These results imply the absence of autocorrelation among disturbances. The lag order ρ of the panel ECM based causality tests is accordingly found to be 1. Unidirectional causality running from forest product exports to economic growth is uncovered in both the short-run and the long-run. A change in forest product exports has an impact on economic growth but not vice versa for individual high-income countries such as Australia, Austria, Canada, etc. as well as for the whole.

TABLE 4: PANEL ECM-BASED CAUSALITY TEST

Variable	$\Delta LGDP_{it}$	$\Delta LFOR_{it}$
$\Delta LGDP_{it-1}$	0.271	-0.257
	(0.301)	(1.348)
$\Delta LFOR_{it-1}$	-0.144	0.120
	(0.082)‡	(0.267)
ECT_{it-1}	-0.028	-0.319
	(0.009)*	(0.196)
Constant	0.027	0.074
	(0.016)‡	(0.129)
Observations	880	880
Number of Instruments	12	12
Sargan Test of Over-Identifying Restrictions	14.58	12.79
	[0.068]‡	[0.119]
Hansen Test of Over-Identifying Restrictions	9.36	13.11
	[0.313]	[0.108]
AR(1) Test of Serial Correlation	-1.19	-1.63
	[0.232]‡	[0.102]
AR(2) Test of Serial Correlation	-1.57	13.11
	[0.117]+	[0.108]
Short-Run Causality Test	3.10	0.04
	[0.078]‡	[0.849]
Long-Run Causality Test	10.60	2.65
	[0.001]*	[0.104]

Note: The standard errors are in parentheses. The lagged dependent variables are endogenous and thus instrumented in GMM-style (Roodman, 2006).

Finally, the short-run and long-run income elasticities of forest product exports are computed and reported in Table 5. The short-run elasticity reveals positive and significant income elasticity. A 1% increase in forest product exports will lead to a 0.022% increase in economic growth among the rich countries in the short-run. The regional dummy is also significant and positive implying that countries with significant forest land coverage are bound to experience higher economic growth. The panel DOLS estimator also shows a positive and lower long-run elasticity. A 1% increase in forest product exports will lead to only a 0.002% increase in economic growth among the rich countries.

TABLE 5: SHORT-RUN AND LONG-RUN COEFFICIENTS

Country	**Short-Run**			**Long-Run**
	$\Delta LFOR_{it}$	ECM_{it-1}	D_{it}	$LFOR_{it}$
Panel	0.022	-0.002	0.030	0.002
	(0.005)*	(0.001)+	(0.002)*	(0.001)+

Note: The ECM model is run using pooled OLS technique. The critical values of the two-tailed t-statistics test at 1%, 5% and 10% significance levels are -2.326, -1.645 and -1.282 respectively. For the DOLS, the leads and lags are set to 2 (Nelson and Donggyu, 2003). The Wooldridge's (2002) test statistic of F(1, 21) = 3054.48 (p-value=0.000) reveals autocorrelation. The DOLS therefore includes time dummies.

CONCLUSIONS AND POLICY IMPLICATIONS

The paper attempts to examine the forest product export-led growth hypothesis for 22 rich countries over the period of 1970-2011. Together with several time-series unit root tests, 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 dependence. A panel causality test is conducted and a unidirectional causality from forest product exports to economic growth is found in the short-run and a bi-directional causality is uncovered in the long-run. Moreover, a 1% rise in forest product exports causes a 0.022% and 0.002% rise in economic growth in the short-run and long-run respectively for the whole panel.

These findings have important implications for policymakers in assisting them to make projections and implementing natural resource and forest policies. Unidirectional causality implies forest product exports can be used to predict economic growth in both short-run and long-run but not vice versa. In general, we do find support for forest product ELG hypothesis. Promotion of forest product exportations can lead to a multiplier effect. However, a caveat arises as the magnitude of the exportation is much lower in the long-run.

It is also important to highlight that increases in forest product exportations might have adverse environmental effects, such as reduced biodiversity, carbon sequestration and recreation areas. Especially for countries which are relatively less endowed with forest areas or are currently utilizing are large share of it. As a result, overexploitation of forest resources can result in unaccounted environmental damages.

Natural forest management programmes which support the implementation of international standards and processes could guarantee the production of forest products without depleting the forest ecosystems. In addition, government schemes to promote innovative technologies at forest industry facilities will lead to higher-value mix of forest products and this will provide greater scope to expand the forest product market. In sum, it is crucial for the rich countries to preserve their forests and their exploitation should be done in a sustainable manner.

ENDNOTES

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