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FDI AND NATURAL RESOURCE RENTS: EVIDENCE FROM EIGHT POST-COMMUNIST COUNTRIES

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This study examines the long-run relationship between foreign direct investment (FDI), trade openness, and natural resource rents in Transcaucasian and Central Asian post-communist countries. The paper deals with cross-sectional dependence, homogeneity restrictions, and panel unit root tests. The departure from earlier studies of the role of natural resources in attracting FDI is in the asymptotic theory of likelihood-based panel cointegration allowing for multiple cointegrating vectors. The results reveal that the variables cointegrate in six countries out of eight under review. This implies that natural resource rents and trade openness promote FDI. Policy implications and legal aspects are discussed.

Keywords: Foreign Direct Investment, Natural Resource Rents, Likelihood-Based Panel
JEL Classification: F23, F41, C20

1. INTRODUCTION

Theoretical literature has long emphasized the importance of natural resources in foreign direct investment (FDI) context. On the one hand, Dunning's (1980) OLI paradigm states that firms engage in foreign investment, when the three advantages: ownership, location, and internationalization, are achieved. He argues that international firms can reduce transaction costs and avoid trade barriers by engaging in foreign investment and gaining control over critical resources that can be used as leverage in the host country.

On the other hand, resource dependence theory discusses that resources are important for organizational success and that continuous access and control over resources boost competitive advantage (Gaffney et al., 2013). Market constraints and uncertainties in the stable flow of natural resources, such as raw materials, are vital factors influencing FDI location decisions of MNEs (Pfeffer and Salancik, 2003). The Heckscher-Ohlin trade theory also explains that comparative advantages in factor endowments are important for trade and investment.

However, it should be stated the relevance of the Prebisch-Singer Thesis (PST), i.e., the price of primary commodities declines relative to the price of manufactured goods over the long term, which causes the terms of trade of primary-product-based economies to deteriorate. Natural resources are significant components of the economy, especially in developing countries where the resource extractive sector generates a considerable portion of the gross domestic product (Shapiro et al., 2018).

There are various channels by which conventional tradable sectors may be crowded out by a booming resource sector and the non-tradable sector including: i) increased productivity in the resource sector drives wages up, bidding labor out of the production of the manufacturing sector, additionally, since natural resource sectors are likely to offer higher returns on investment (by exploiting the resource rent), investment and thus economic growth would tend to be biased towards the resource sector; ii) amplified incomes shift demand from the lagging tradable sectors to non-tradable, where wages will also be pushed up. This spending effect will further drain factors of production out of the non-resource tradable sector (Dutch Disease).

Resource extraction is a capital-intensive activity that requires high levels of capital investment. FDI is perceived to have spillover effects, such as job creation, productivity boost, competitive enhancement, and technology transfer to other industries which do not necessarily occur because natural-resource-rich countries tend to devote resources to that same industry.

The problem is that the above argument could create an appreciation of the currency and make the non-resource and manufacturing sectors less competitive (Dutch Disease). This diversion of resources to the natural resource sector leads to a “crowding out” effect relative to the other sectors and may generate a contraction of other tradable activities. Kojo (2015) argues that, in long-run economic models, capital and labor are assumed to be perfectly mobile internationally, so the real exchange rate is not affected by an export boom.

Poelhekke and Van der Ploeg (2013) study the effects of natural resources on FDI to the resource sector and the non-resource sector. They reveal that natural resources attract FDI in the resource sector but crowd out FDI in the non-resource sector. This crowding out effect is stronger for countries that were not resource producers in the past, but in general, the contractions of non-resource FDI outweigh the gains from resource FDI.

Asiedu (2013) examines the effects of oil exports and oil rents on aggregate FDI inflows. The author shows that natural resources have an adverse effect on FDI. Bokpin et al. (2015) study the effect of natural resources on FDI in developing countries. They use decomposed measures of natural resource rents, identifying between oil rents, mineral rents, and forest rents. They conclude that different measures of natural resource rents can have different impacts on FDI inflows. Gonchar and Marek (2014) show that resource FDI does not crowd out non-resource FDI and argue that the difference between positive and negative outcomes depends on the measurement criteria chosen for available resources.

Although the natural resources and FDI nexus has received considerable attention (e.g., Asiedu, 2013; Doytch, 2015; Kang, 2018; Lu et al., 2020), however, there is no common consensus over the impact of natural resources on inward FDI. This work aims at answering the following question: how do natural resource rents affect FDI? Thus, we choose Transcaucasian and Central Asian post-communist countries where most of the FDI can be explained by the abundance of oil, natural gas, minerals, and the gravity of natural resources in attracting FDI inflows to the region.

The main contribution of this study stems from its used methodology which is a likelihood-based panel cointegration under assumptions of cross-sectional dependence and slope homogeneity restrictions. This is an extension of the Johansen (1995) multivariate maximum likelihood developed by Larsson and Lyhagen (1999) and Larsson et al. (2001). They have developed a likelihood-based panel test of the cointegrating rank and a general likelihood-based framework for inference in panel-VAR models with cointegration restriction, allowing for multiple cointegrating vectors. By using this method, the assumption of a unique cointegrating vector and the problem of normalization is relaxed. This is not the case with the usual residual-based tests of cointegration (e.g., Kao, 1999; Pedroni, 1999). However, to the best of the author's knowledge, this study is the first attempt to test the impact of natural resource rents on FDI in the developing countries using panel cointegration techniques based on likelihood inference of cointegrating vectors.

The remainder of this paper is organized as follows: Section 2 introduces our model, data, and methodology. In Section 3, the empirical evidence is presented. Finally, Section 4 offers conclusion, discussion, and policy implications.

2. MODEL, DATA AND METHODOLOGY

Given the above background, we can use the following model:

$$FDI_{it} = \delta_{0i} + \delta_{1i}NRR_{it} + \delta_{2i}TOP_{it} + \varepsilon_{it}, i = 1, \cdots, N, \ t = 1, \cdots, T, \tag{1}$$

where FDI, NRR, and TOP are net inflows of foreign direct investment, natural resource rents, and trade openness, respectively. ε is a well-behaved disturbance term. The sample consists of all ex-communist Transcaucasian and Central Asian economies (Azerbaijan, Armenia, Georgia, Tajikistan, Turkmenistan, Kazakhstan, Uzbekistan, and Kyrgyzstan) and covers the period 1990-2016. The choice of the time period is dictated by data availability and the reason for choosing these countries is the fact that the countries under review are important recipients of FDI after the collapse of Soviet.

The dependent variable examined is annual net foreign direct investment (FDI) inflows as a percent of GDP. While it would have been ideal to distinguish between resource and non-resource FDI such data is not generally available. The explanatory variables are total natural resource rents as a share of GDP (measured as the sum of oil rents, natural gas rents, coal rents, mineral rents, and forest rents) and trade openness (measured as total trade as percent of GDP) which represents locational advantages like reduction of trade cost and avoidance of trade barriers. The variables are extracted from the World Bank Development Indicator (WDI) and the United Nation Conference on Trade and Development (UNCTAD). Descriptive statistics for the variables under analysis is reported in Table A1, Appendix A. The process is estimated by implementing likelihood-based panel framework developed by Larsson and Lyhagen (1999) and Larsson et al. (2001). By using this method, the assumption of a unique cointegrating vector and the problem of normalization is relaxed which is not the case with the usual residual-based tests of cointegration approach. Let LR denote the cross-section-specific likelihood-ratio (trace) statistic of the hypothesis that there are at most r cointegrating vectors in the system. The standardized LR-bar statistic is given by:

$$Y_{\overline{LR}} = \frac{\sqrt{N(\overline{LR} - \mu)}}{\sqrt{V}}, \tag{2}$$

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

Two steps should be followed before using any cointegration tests: testing the panel for cross-sectional dependence and testing for cross-country heterogeneity. The first issue means the transmission of shocks from one variable to others. In other words, all countries in the sample are affected by globalization and have common economic characteristics. The second issue shows that a significant economic connection in one country is not necessarily replicated by the others. A set of three tests is constructed in order to check the cross-sectional dependence assumption: the Breusch and Pagan (1980) cross-sectional dependence (CDBP) test, the Pesaran (2004) cross-sectional dependence (CDP) test, and the Pesaran et al. (2008) bias-adjusted LM test (LMadj). Regarding the country-specific heterogeneity assumption, the slope homogeneity tests ($\bar{\Delta}$ and $\bar{\Delta}_{adj}$) of Pesaran and Yamagata (2008) are used (Appendix B provides more information about these tests). The traditional panel unit root tests do not consider cross-sectional dependence of the contemporaneous error terms. Failing to take into account cross-sectional dependence may lead to misleading results. Thus, to eliminate this problem, we use the cross-sectionally augmented panel unit root test (CIPS) which allows for parameter heterogeneity and serial correlation between the cross-sections (Pesaran, 2007).2

Finally, we check diagnostic tests, i.e., if the residuals are normally distributed and

[1] This methodology is also used in Irandoust and Ericsson (2005), and Irandoust (2020).

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

there is no autocorrelation. The normality test stems from a multivariate extension of the Bowman–Shenton test developed by Doornik and Hansen (1994) and the test for autocorrelation is the Ljung-Box test statistics.

3. ESTIMATION RESULTS

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

Table 1. Cross-Sectional Dependence and Slope Homogeneity Tests

Method	Test statistic
Cross-sectional → dependence test	
CD_{BP}	191.376*** (0.000)
CD_P	33.528*** (0.000)
LM_{adj}	44.021*** (0.000)
Slope homogeneity test	
$\bar{\Delta}$	15.319*** (0.000)
$\bar{\Delta}_{adj}$	12.560*** (0.000)

Notes: 1) *** indicate significance for 0.01 levels. The numbers within parentheses show *p*-values. 2) CD_{BP} test, CD_P test, and LM_{adj} test show the cross-sectional dependence tests of Breusch and Pagan (1980), Pesaran (2004), and Pesaran et al. (2008), respectively. 3) $\bar{\Delta}$ and $\bar{\Delta}_{adj}$ tests show the slope homogeneity tests proposed by Pesaran and Yamagata (2008).

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

Table 2. Panel Unit Root Test

Variable	CIPS Statistic
Cross-sectional → dependence test	
NRR	-2.146
TOP	-2.007
FDI	-1.925

Note: Critical values for the CIPS test are -2.57 (1%), -2.33 (5%), and -2.21 (10%), Pesaran (2007).

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

Table 3. Test for Cointegrating Rank

H_0	ACV^a	BCV^b	$-2logQ_T$
$R = 0$	428.18	595.62	566.33
$R \leq 0$	220.21	453.29	342.41
$R \leq 2$	96.28	215.13	139.92

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

Since the panel cointegration tests show that the common cointegrating rank is one, thus, it is interesting to estimate the cointegrated vectors. The estimated cointegrating vectors, normalized with respect to FDI are presented in Table 4.

According to Table 4, we can assert that FDI is positively associated with NRR and TOP for almost all countries in the sample. Exceptions are Armenia and Tajikistan. In these countries the coefficients for NRR and TOP have a very low value and they are not significant. Once again, it seems that Armenia and Tajikistan are different from the rest six countries. However, the magnitude of parameters varies from country to country. There is cointegration in Azerbaijan, Georgia, Turkmenistan, Kyrgyzstan, Kazakhstan, Uzbekistan. We could not find any cointegration in Armenia and Tajikistan. This may

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

stem from three reasons (i) a rather low share of natural resource rents, (ii) a rather low degree of trade openness compared to the other countries in the sample, and (iii) regulatory restrictions on FDI and low level of governance.

Table 4. Cointegrating Vectors Normalized on FDI

	Armenia	Azerbaijan	Georgia	Kazakhstan	Kyrgyzstan	Turkmenistan	Tajikistans	Uzbekistan
FDI	-1.000	-1.000	-1.000	-1.000	-1.000	-1.000	-1.000	-1.000
TOP	0.167	2.553	1.369	1.470	2.508	3.792	0.137	1.125
NRR	0.086	3.729	1.475	2.541	2.631	4.865	0.062	1.392

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

Table 5. Diagnostic Tests[a]

Normality	Autocorrelation
0.040	0.659

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

4. DISCUSSION, CONCLUSION AND POLICY IMPLICATIONS

In this paper, the effect of natural resource rents and trade openness on FDI inflows is examined by using a panel data set over the period 1990-2016. The countries under review are: Armenia, Azerbaijan, Georgia, Turkmenistan, Kazakhstan, Kyrgyzstan, Tajikistan, and Uzbekistan. The departure from earlier studies of the effect of natural resource rents on FDI is in the asymptotic theory of likelihood-based panel cointegration under assumptions of cross-sectional dependence and homogeneity restrictions. This method allows for multiple cointegrating vectors, which is not the case with the usual residual-based tests of panel cointegration. Hence, the assumption of a unique cointegrating vector and the problem of normalization are relaxed.

The tests for panel cointegration reveal one cointegrating vector. However, the findings show that natural resource rents and trade openness positively associated with foreign direct investment inflows in all countries in the sample except in Armenia and Tajikistan. This result is consistent with the economic theory of FDI and factor endowments trade theory which assert that natural resource rents enhance FDI inflows by supplementing the domestic capital formation. In other words, natural resource rents and trade openness contribute to FDI inflows.

Regarding the lack of cointegration in Armenia and Tajikistan, it is worth to mentioning that the transition from planned to market economy requires restructuring institutional framework and economic policy reforms. Laws, regulations, and public institutions that protect the security of private property rights, the competence of the civil service in performing state activities, and the transparency of the legal system are crucial factors in attracting FDI.

Mariotti and Marzano (2021) reveal that the effectiveness of competition policy enforcement is a significant factor in encouraging FDI, but only in host countries characterized by institutional configurations where the lack of trust is accompanying with a high-quality regulatory institutional environment. They assert that the high quality of the regulatory institutional environment, and the country effectiveness in implementing strategies to boost competitive market mechanisms will actually hinder discriminatory competition policies and market manipulations. They conclude that the effectiveness of competition policy enforcement, together with other regulatory policies, depends on important complementarities between the pillars of the national institutional setting.

Given the uncertainties surrounding legal aspects in this region and in the context of national development objectives, the greatest value of a foreign investment statute may show its potential for perceiving this community of interest in explicit terms-first at the initial stage of negotiating a foreign investment-development agreement, and subsequently in obtaining the continuing contribution of the foreign investment to the development goals of the host country. Thus, a cost-benefit analysis and the economic realities of the development process should be of higher importance for the protection of private foreign investment than are the legal doctrines that have dominated and artificially limited the approach to this problem (Meier, 1966).

However, the economic implication of our results is that natural resource rents have a key role in boosting developing countries' FDI. That is, there is a positive relationship between the variables because FDI not only augments domestic resources, but also assists to close the foreign exchange gap, creates access to modern technology and managerial skills, and allows easier access to foreign markets.

The results confirm the presence of a long-run relationship between all variables considered (except in two countries out of eight under review). Thus, if policymakers wish to boost long-run economic growth and FDI, our general findings augment the case for assisting the development of the natural resources alongside the encouragement of trade openness or freer trade. A more stable macroeconomic environment can promote all macroeconomic variables associated with FDI. The main message from our study for researchers and policymakers is that inferences drawn from research that do not consider the multiple cointegration vectors, the cross-sectional dependence and homogeneity restrictions, and the dynamic interrelation of all the macroeconomic and policy variables in our study will be unreliable. It is the conjoint interplay between natural resource rents, foreign direct investment, and the openness to trade that distinguishes our study and guides future research on this topic.

There are a few limitations of this study. First, the data used here was not sectorial FDI inflows and second, the model did not take into account nonlinearity. A natural extension to this paper would carry a similar analysis using sectorial FDI inflows and a nonlinear estimation methodology.

APPENDIX

Appendix A.

Table A1. Descriptive statistics of the variables, 1990-2016, n = 27 for each individual country

Country	Mean	S.D.	Skewness	Kurtosis
Armenia				
ARMNRR	1.440808	1.505095	0.914399	2.290295
ARMTOP	76.67688	14.52359	0.934851	3 520586
ARMFDI	4.313554	3.107431	0.431716	2.779965
Azerbaijan				
AZENRR	24.24890	10.25813	0.053819	2.036780
AZETOP	88.77896	20.17088	1.082028	3.611447
AZEFDI	14.35884	15.10754	1.471270	4.480714
Georgia				
GEONRR	0.863648	0.489872	1.147667	4.727247
GEOTOP	83.14309	24.03237	1.419033	6.639413
GEOFDI	6.279560	4.930789	0.532391	2.801008
Kyrgyzstan				
KGZNRR	4.106061	3.715982	0.675691	2.220722
KGZTOP	102.2601	25.05388	0.398392	1 640728
KGZFDI	4.703217	4.127813	1.028183	4.095166
Tajikistan				
TJKNRR	1.195084	1.099728	2.061982	7.409573
TJKTOP	107.0206	39.30774	0.219624	1.749463
TJKFDI	3.382884	3.539890	1.627804	4.527571
Kazakhstan				
KAZNRR	18.06836	8.571159	-0.024681	1.653781
KAZTOP	88.54500	25.75008	1.358093	4 330 381
KAZFDI	6.846881	3.961268	0.105743	1.979662
Turkmenistan				
TURNRR	45.19871	19.69925	-0.123183	2.303861
TURTOP	94.39742	33.51897	0.834264	3 011903
TURFDI	6.617787	4.489147	1.966517	7.359648
Uzbekistan				
UZBNRR	16.57372	9.334682	0.291593	1.789059
UZBTOP	53.19445	13.94941	0.250868	2.166317
UZBFDI	1.167915	0.976941	0.952494	3.043865

Appendix B.

Cross-sectional Dependence Tests

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

$$y_{it} = \alpha_i + \beta_{it} x_{it} + \mu_{it},\ i = 1, 2, \cdots, N,\ t = 1,2, \cdots T, \tag{1A}$$

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

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

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

$$CD_{BP} = \sqrt{\frac{1}{N(N-1)}} \sum_{i=1}^{N-1} \sum_{j=i+1}^{N} (T\hat{\rho}_{ij}^2 - 1). \tag{3A}$$

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

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

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

Slope Homogeneity Tests

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

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

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

$$\bar{\Delta}_{adj}= \sqrt{N}\left(\frac{N^{-1}S^{\approx}-E(z_{it}^{\approx})}{\sqrt{var(z_{it}^{\approx})}}\right), \tag{6A}$$

where $E(z_{it}^{\approx}) = k, var(z_{it}^{\approx}) = 2k(T-k-1)/(T+1)$.

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