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THE IMPACT OF TRADE OPENNESS ON INCOME INEQUALITY IN THE FORMER CENTRALLY PLANNED MIDDLE–INCOME ECONOMIES

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ABSTRACT

This study explores the association between income inequality and trade openness in the former centrally planned middle–income economies over the period 1994–2019. Apart from the full sample, the European Union member or candidate middle–income Balkan economies, and the middle–income Commonwealth of Independent States subsamples were used in estimations as benchmark. Foreign direct investment, inflation, government expenditures and gross fixed capital formation control variables are incorporated into the PMG–ARDL model used in the study. The results for the sample including the middle–income Balkan economies reveal that within 1% increase of trade openness, income inequality worsens by 0.27. We suggest these countries strengthen their income redistribution strategies while pursuing policies that promote integration into the global market.

Keywords: *Transition Economies, Trade Openness, Income Inequality, Panel ARDL.*

JEL Codes: *C33, D31, D63, F10, F60, O15, R11.*

1. INTRODUCTION

The influence of trade openness (*TROP*) on income inequality (*ININ*) is an important topic of discussion among scholars in the era of contemporary globalization especially in developing economies. In this context, these countries seek trade strategies and policies that might increase economic growth rates and reduce *ININ* simultaneously.

Trade is considered as a critical factor that contributes to the improvement of economic growth by increasing competitiveness and productivity. It is predicted that the economy might be more efficient due to the increase in trade openness and competition through international trade, and consumers shall benefit from the having a wide variety of goods and services at different quality and price levels. Trade openness also has potential effects on income distribution, and job gains or losses. While the findings of empirical investigations broadly support a positive association amid trade and growth, there are situations where growth is accompanied by worsening poverty and income inequality. Whether *TROP* is an instrument contributing to the rise of *ININ* is among the debates in the economics literature. Although it has been found by some researchers that trade openness alleviates poverty on average around

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the world, it appears to bring about distributional changes that can *ININ* in some economies (Winters et al., 2004). However, there are disagreements among researchers on the influence of *TROP* on *ININ*. During transformation and transition, risks and costs arising from trade openness might have significant impacts on the emerging economies and cause unequal distribution of the benefits of economic growth. One of the issues related to income inequality is the high trade level between countries partially activated by the technology enhancement (Norris et al., 2015). The use of new technologies brings the skilled and educated workforce to the forefront rather than unskilled laborers and causes a wage gap between them resulting in income inequality. However, it might be the case that trade openness in developing countries decline inequality as a consequence of increasing demand for unskilled labor (Ravallion, 2004).

The direction and magnitude of the impacts of *TROP* on income distribution are shaped by countries' global economic policies and growth models. Among these, it is discussed whether the liberalization in the economic field in the countries that switched to market economy in the 1990s caused the emergence or deepening of imbalances between regions, as well as other economic issues. It was observed that transitioning to a market economy had resulted welfare losses in the former centrally planned economies in the short–run. Welfare losses were evident in terms of traditional income–based indicators such as wage rate, per capita income and consumption, poverty rates as well as demographic variables (Cornia, 1996; Cornia et al., 1996). Nevertheless, there has been acceleration in the speed of globalization in the last two decades. While some Central and Southeastern European, and Baltic transition economies were included into the enlargement of the European Union (EU), a part of them has been continuing this process within the Commonwealth of Independent States (CIS). Although the paradigm shift resulted in some of the countries integrated into the EU rising to the high–income countries, the most of transition countries remained among the middle–income economies.

The intention of this article is to explore the liaison amid income inequality and trade openness with a particular focus on the former centrally planned economies. Furthermore, due to the emergence of different trade orientations in the transition process, these economies will be examined in subpanels in terms of *TROP* – *ININ* association according to their status as EU member – candidate and CIS countries. Although some individual country or a few panel studies targeted these economies, as far as we know, none have examined these issues focusing on the middle–income post–communist economies and its subgroups particularly.

2. LITERATURE REVIEW

2.1. The Linkage between Trade Openness and Income Inequality

There is an ongoing debate in theoretical literature whether trade openness has income inequality exacerbating or narrowing impacts. The classical theoretical framework for exploring the liaisons amid *TROP* and *ININ* is the Heckscher–Ohlin (Ohlin, 1967), which was raised on the theory of comparative advantage (Ricardo, 1817). The Heckscher–Ohlin (H–O) model asserts that economies' specialization in production within the relatively abundant factor causes them to focus on the export of these goods. In other words, the goods which use abundant factors of production intensely are exported, while the goods that make intense use of scarce factors of production are imported. In this model, the inequality impact of trade openness and the extent to which individuals are dependent on capital or labor income are explained according to the productivity differences and relative factor endowments of economies (Dorn et.al., 2022). One of the most important consequences of the H–O is the Stolper–Samuelson theorem (Stolper and Samuelson, 1941) which describes the association between factor rewards and commodity prices. The theorem states that the subsequent relative product price changes due to the *TROP* increase the real return of the factor(s) which is used most intensively in the production of factor–abundant exports goods, while causing a decrease in other factor(s). The size and magnitude of the impacts of *TROP* on *ININ* depends on the degree of development of an economy which indicates that developing countries shell export more labor–intensive products while they are importing more skill– and capital–intensive types of goods (Meschi and Vivarelli, 2009). Some of the papers published in recent years revealed that the Stolper Samuelson theorem in H–O framework is consistent for developing countries. Among these research, Lin and Fu (2016) concluded that *TROP* had been decreasing *ININ* in autocracies. Yang and Greaney's (2017) Engle–Granger two-step ECM estimation results suggested that *TROP* had a decreasing impact on *ININ* in China. Ponce et al. (2023) affirmed that *TROP* led to a significant decrease in *ININ* in small developing countries. Findings from the quantile regression analysis by Tufaner and Özen (2023) suggested that *TROP* was beneficial to income distribution for MIST countries including Mexico, Indonesia, South Korea and Türkiye. However, empirical evidence remains inconclusive to validate the H–O and Stolper Samuelson predictions for all developing countries. In contrast, a large body of recent research provided significant empirical evidence that *TROP* is detrimental to *ININ* for many developing countries. Ezcurra and Andrés Rodríguez-Pose (2014) identified a robust exacerbating impact of *TROP* on *ININ* in 22 emerging economies. Empirical results obtained by Lee at al. (2020) revealed that *TROP* had been exacerbating *ININ* in lower income and non-OECD countries. The findings of Xu et al. (2021) also showed that *ININ* was positively associated with *TROP* in sub-Saharan Africa. Naanwaab (2022) noted that the free trade between North–South caused *ININ* to increase in developing countries. On the other hand, some published studies do not provide any clear–cut evidence on the relationship between *ININ* and *TROP* in developing countries. By applying an ARDL bounds testing approach, Bayraktar et al. (2019) found an insignificant association between *ININ* and *TROP* for Türkiye. The findings of the panel data analysis conducted by Çelik and Erkişi (2022) did not provide significant evidence on the relationship between *ININ* and *TROP* for nine upper–middle and lower–middle income countries. Similar results were obtained for low–income countries from a meta–analysis by Huang et al. (2022).

According to the Kuznets curve theory (Kuznets, 1955), inequality would follow an inverted–U shape due to the transfer of excessive labor in the agricultural sector to other industries during economic growth. Although the Kuznets Curve and the Stolper and Samuelson theorem are based on growth and foreign trade respectively, the development in income distribution is associated with long–run structural changes in both cases. Given this, it might be predicted that *ININ* would increase in the initial stages of *TROP*. Due to the positive effects of the globalization such as the technological developments and the deepening of the financial system, *ININ* would begin to reduce after a certain level. Even though the Kuznets curve considers economic development as an independent variable, there are also some papers in the literature that explore the presence of Kuznets curve based on a number of variables linked to the globalization such as *TROP* and foreign direct investment (*FDI*). Dobson and Ramlogan (2009) examined the validity of the Kuznets curve hypothesis using data for Latin America. The evidence showed that the association between *TROP* and *ININ* followed an inverted U–shaped pattern. In contrast, Topuz and Dağdemir (2020) detected a U–shaped association amid *TROP* and *ININ* in Türkiye. Using a panel dataset for 59 countries, Gerni et al. (2018) explored the connection between per capita income and per capita net foreign direct capital investments. The authors affirmed that the Kuznets curve appeared in the period under investigation.

The internal growth problems that emerged with globalization brought the need to look beyond the aforementioned framework. Early endogenous growth (or trade) models which maintain that primary cause of economic growth is internal forces, rather than external ones were introduced to economics literature by Romer (1986) and Lucas (1988). The endogenous growth theory predicts that while there is no increase in income inequality in developed countries because of the increasing use of technology, the majority of developing countries do not benefit from globalization due to their fundamental macroeconomic problems (Feenstra and Hanson, 1999).

2.2. Empirical Literature: Trade Openness – Income Inequality Nexus in Transition Economies

The number of empirical studies focusing on the linkage between *TROP* and *ININ* in transition economies is relatively small. The findings of these studies differ depending on countries or country groups examined, time intervals focused, and the empirical models used. Hypothesizing that the increasing level of *FDI* and *TROP* since 1989 could be significant determinants of *ININ* in transition economies, Franco and Gerussi (2013) applied a panel estimation model which observed over the period 1990–2006 and consisting of 17 transition economies. The findings indicated that *FDI* and *TROP* had insignificant impacts on *ININ*. In the next stage, they deepened their study by disaggregating trade in exports and imports. The findings revealed the presence of a significant association between imports and exports, and inequality. However, the capacity of these variables to affect inequality were divergent. Dorn et al. (2022) examined the influence of *TROP* on *ININ* comparatively for developed and developing countries within the framework of the predictions of Stolper–Samuelson theorem. The subgroups of the study, in which they estimated the basic panel model by ordinary least squares techniques, included transition economies. The findings of their study spanning from 1970 to 2014 implied strong impact of *TROP* on *ININ* within transition countries. A recent study by Badur and Sohag (2023) explored whether trade integration impacted *ININ* for the panel of 12 post–Soviet states over the 1991–2019 period. The outputs of the quantile regressions via method of moments showed that openness to trade contributed to *ININ* monotonically.

Although not much research has been conducted particularly focusing on the association amid *TROP* and *ININ* in transition economies, *TROP* is used as a control variable in some of the papers exploring the determinants of *ININ*. While some of these studies provide outcomes for the relationship between *TROP* and *ININ* in panel data from transition economies or from those including former centrally planned economies as well as other countries, others explore the possible factors impacting income inequality in individual countries. Neagu et al. (2016) investigated the impacts of a number of variables on *ININ* in ten transition economies over the period 2000–2014. The findings of panel data analysis significantly indicated that *TRO*P had an increasing effect on *ININ*. Alili and Adnett (2018) explored the influence of *FDI* on wage inequality in 19 selected transition economies by employing panel data techniques. Trade openness was among the control variables in the model they estimated. The findings did not provide evidence that increased *TROP* reduced wage inequality. Acaravcı et al. (2018) explored the causal links between *TROP*, *ININ* and democracy in the Balkan countries over a period of 1996 to 2010. A significant causality relationship amid *TROP* to *ININ* was revealed for Bulgaria and Romania. However, such a causality link was not found for Croatia and Slovenia by the authors. Kumo et al. (2018) investigated (i) whether the increasing trade openness in Russia in the 1990s and 2000s favors the poor at the regional level, and (ii) whether the distributional effect of increasing free trade on incomes is positive or not in the regions investigated. The findings showed that the influence of *TROP* on *ININ* was significantly negative. The findings also indicated that the export of resources tended to be related to greater inequalities. Braha–Vokshi et.al. (2021) explored the influence of multinational enterprises on *ININ* in 6 former Yugoslavian countries from 2007 to 2019. The results provided evidence that trade openness, which was one of the control variables, had a decelerating effect on inequalities. Tsaurai (2021) utilized a panel data from 2009 to 2019 for 11 Central and Eastern European countries and suggested mixed results on the *TROP–ININ* nexus. While the findings of the Dynamic GMM revealed positive but insignificant impact of *TROP* on the Gini coefficient (*GINI*), the results based on the pooled OLS techniques showed that *TROP* significantly enhanced *ININ* in Central and Eastern European countries. However, trade liberalization resulted in a reduction in *ININ* in these countries according to the outputs of the fixed– and random–effects models. Sadiku et al. (2023) used Least Squares Dummy model to analyze panel data for Central and Eastern European countries over the period 2003–2020. The results of the study focusing on the macroeconomic determinants of *ININ* showed that *TROP* was not among the factors significantly impacting *GINI*. Saputro and Aisyah (2023) explored the determinants of *ININ* in six lower middle–income countries two of which — Kyrgyzstan and Ukraine — are former centrally planned economies. Using panel data techniques for the period of

2017–2021, the authors obtained evidence that trade openness was among the significant factors adversely affecting income distribution.

A group of studies examining the determinants of *ININ* focuses on the linkage between economic growth and distribution of income. Velkovska et al. (2021) employed a panel data regression formed from 2001 to 2012 for the countries in the EU and the EU candidate countries. The findings related to the EU candidate Balkan countries subsample were consistent with the Kuznets curve hypothesis, implying that economic growth initially decreased *ININ*, but as rising incomes reached a turning point, *ININ* began to increase. Also, Vezentan and Neagu (2022) obtained evidence confirming the existence of Kuznets curve for Romania by applying ordinary least squares regression. Fawaz and Rahnama (2022) scrutinized the economic growth – income inequality linkage in low– and high–income transition economies by utilizing fixed–effects and dynamic panel techniques. The results suggested that *ININ* was positively associated with economic growth in high–income transition economies. In contrast, the authors provided evidence that economic growth reduced *ININ* in low–income transition economies. A more recent study validating the Kuznets curve hypothesis was conducted by Badur et al. (2023). The outcomes of the panel data regression performed by the authors for 12 post–communist countries and 24 years (1996–2019) highlighted the presence of a U–inverted curve in terms of income inequality.

Previous studies provide evidence that some factors other than trade openness and economic growth have impacts on income distribution in the former centrally planned economies. The results obtained by Roy–Mukherjee and Udeogu (2021) via multiple linear regression using data from 1991 to 2017 confirmed that economic or export complexity level and labor unionization degree reduced *ININ* in western Balkan countries. Using Ordinary Least Squares regression techniques, Özparlak and Özhan (2021) explored the role of financial inclusion in reducing *ININ* in 11 countries in Central Asia, Caucasia and Balkans, all of which are former centrally planned economies except Türkiye. Analysis results showed that a rise in financial inclusion caused a reduction in *ININ*. Recepoglu (2022) explored the connection between public expenditures, economic growth and *ININ* by utilizing the dataset of the CIS economies spanning from 1998 to 2019. The outputs of the Bootstrap Panel Granger Causality test suggested a unilateral causality association running from public expenditures to *ININ* for Kazakhstan and Belarus. Tavadyan and Ghazaryan (2022) investigated the association amid FDI and Palma ratio in Armenia. The empirical findings indicated the existence of an inverted U–shaped association between the variables. Tushaj et. al. (2023) examined the factors impacting income inequality in Albania, Kosovo, Montenegro and Serbia. The results obtained from the Generalized Method of Moments estimator indicated that financial development adversely impacted *ININ*. Yuldashev et al. (2023) explored the *FDI* and *ININ* connection for ten countries in Asia for the 1999–2020 period. The authors presented empirical evidence that *FDI*s reduced *ININ* in Kazakhstan and Uzbekistan, even more effectively in the existence of human capital.

3. MODEL, DATA, AND METHODOLOGY

3.1. Sample

The countries remained in the centrally planned system for at least fifty, some approximately seventy years, followed different transition paths. In this process, while most of the middle–income Central European, Baltic and Balkan transition economies became a member or candidate of the EU in which economic convergence is among the main objectives for the achievement of prosperity (Albu et al, 2019), almost all transition economies in Central Asia and the Caucasus remained within the CIS. On the other hand, Ukraine and Georgia stayed outside of both groups, despite their recent rapprochement with western economies. Since *TROP* may have had different effects on *ININ* in these economies, we have divided them into subsamples depending on their macroeconomic indicators' similarity and enhanced economic relations. The subsamples include the former centrally planned middle–income economies which are members or candidates of the EU, and the CIS countries (Table 1).

Table 1. Panel Structure of Sample

	Countries
Panel–1	Albania, Armenia, Belarus, Bulgaria, Georgia, Kazakhstan, Kyrgyz Republic, North Macedonia, Moldova, Romania, Russia, Ukraine
Panel–2	Albania, Bulgaria, North Macedonia, Romania
Panel–3	Armenia, Belarus, Kazakhstan, Kyrgyz Republic, Moldova, Russia

3.2. Model and Data

Due to the unavailability of full data set for all former centrally planned middle–income economies in Europe, Caucasus and Central Asia, this empirical study examines annual time serial data on 12 of them over the period 1994–2019. The World Bank Country Classifications, created by using the Atlas method which depends on the Gross National Income Levels of the economies, was used to determine the former centrally planned middle–income economies (Worldbank, 2021a). The data other than the income inequality were derived from the databases of the World Bank (Worldbank, 2021b).

Based on the theoretical framework and empirical literature, the following equation is set out to examine the liaison amid *TROP* and *ININ*.

$$GINI_{i,t} = TROP_{i,t} + FDI_{i,t} + INF_{i,t} + GOV_{i,t} + CAP_{i,t} + \lambda_t + \mu_{i,t} \quad \textbf{(1)}$$

where *i* indicate countries, and *t* stands for the time spanning from 1994 to 2019. λ and μ are the time dummies and the error term, respectively.

The Gini coefficient (*GINI*), based on the Lorenz curve as a measure of *ININ*, is the target variable of the data set. Income inequality can be measured by means of the ratios like Decile dispersion ratio and Palma ratio or indices such as Atkinson's index, Schutz index, Theil index and Gini index. Among these measurements, *GINI* is the most widely used proxy of inequality in the literature. The higher *GINI* indicates that the distribution of income in an economy has become much more unequal. The *GINI* data used in the study were gathered from the Standardized World Income Inequality Database (Solt, 2016).

Trade openness, measured as the ratio of the sum of nominal exports and imports to gross domestic product (GDP), is the independent variable of interest in the model. Trade openness is an appropriate indicator which has been extensively used in the literature for exploring the liaisons amid free trade and income distribution. If the Stolper–Samuelson Theorem, based on the H–O model, is valid for the former centrally middle–income economies, *TROP* would reduce *ININ* due to the increase in demand for unskilled labor and the decline in demand for skilled labor. However, the former centrally planned middle–income countries face huge challenges in the process of transition from a planned to a market economy which caused free trade to be frequently expressed as one of the driving forces of inequality for these countries. If the endogenous growth theory asserting that most of the developing countries cannot benefit from free trade due to their fundamental macroeconomic problems is valid, *TROP* would result in increased *ININ* in the examined economies. Furthermore, *TROP* might have an interaction effect on the association amid economic growth and *ININ*. If there is such an effect for the former centrally planned economies under examination, evidence can be obtained regarding the existence of the Kuznets curve.

The controlling variable ''*FDI*'' is the net inflows of foreign direct investments expressed as the percentage of GDP. Impact of *TROP* on *ININ* may come through *FDI*. In the H–O and Stolper Samuelson framework, the influences of *FDI* and TROP on income distribution resemble each other. Therefore, *FDI* is expected to reduce *ININ* in developing countries if the Stolper Samuelson theorem is valid. The re–evaluation of the macroeconomic consequences of *FDI* has appeared as an important phenomenon especially for the Central and Eastern Europe transition economies under the effect of international investment (Săvoiu et al., 2013). Many developing countries encourage foreign investors to capitalize on their country by upgrading doing–business indicators. Reducing the minimum wage and taming the labor union are among the approaches in this direction. However, if foreign investors terminate their activities and leave the host country for various reasons, unemployment rates may increase (Salvatore, 2013). Furthermore, Stiglitz (2002, 2012) asserts that *FDI* might weaken the social security system and thus increase *ININ*. The author claims that, in order to attract foreign investment, developing countries put downward pressure on tax rates, and remove or reduce the restrictions on the factors of production, particularly the capital. Consequently, multinational companies are likely to invest in the countries with low wage levels, poor working conditions and low environmental regulations. Moreover, the economies with a relatively high level of welfare confront the influx of low–skilled immigrants. In both cases, income distribution pattern would be distorted in favor of the corporate sector.

Therefore, to account for distributive *TROP*, it is pertinent to explore the influence of changes in *FDI* flows on the *GINI* (Mah, 2003).

Inflation (*INF*), one of the control variables used in the model, is an important phenomenon particularly for the developing countries characterized by macroeconomic and political instability. A common belief is that inflation impacts the households with different types of income streams including labor income, capital income, private transfers, and state transfers, to different degrees. Inflation has a high potential to erode real wages and disproportionately affect those within the bottom percentiles of the income distribution (Fischer and Modigliani, 1978; Lundberg and Squire, 2003; Gourdon et al., 2008; Cassette et al, 2012; Roser and Cuaresma, 2016; Barusman and Barusman, 2017). On the other hand, unanticipated inflation might redistribute income to some extent. According to the debtor–creditor hypothesis (Kessel and Alchian, 1962), unanticipated inflation reduces the real value of existing debts, and redistributes income between the lenders and creditors in favor of the debtors.

Another control variable ‘‘*GOV*’’ is the total expenditures of government for purchases of goods and services expressed as a percentage of GDP. According to many economists, choices about the types and structure of government expenditures are important factors in reducing *ININ* and poverty rate (Anderson et al., 2017). Keynesian theory emphasizes three paths in which government expenditures reduce *ININ* and poverty: (i) the living standards of low–income households can be improved by easing restrictions on certain types of government expenditures; (ii) the demand for labor can be increased through public projects which results in less unemployment, and lower *ININ* level; (iii) the multiplier influences of public projects may increase economic activities and private sector investments, which will lead to new labor demands (Stack, 1978). In the long–run, even if the total public expenditures remain constant, changing the composition so as to increase the weights of education and health expenditures particularly will reduce *ININ* (Doumbia and Kinda, 2019).

Investment rate (*CAP*), represented by gross fixed capital formation as a percentage of GDP, is the last control variable. Gross fixed capital formation is theoretically considered to be a very important factor in increasing employment and economic growth, ensuring socioeconomic development, improving income equality, and reducing poverty (Bekele and Merid, 2020; Ullah et al., 2021). According to the Kuznets curve, the allocation of investment in the economic sector would cause a rise in *ININ* during the initial stages of economic growth until a turning point beyond which *ININ* reduces with economic growth. Nevertheless, Lewis (1954) asserts that as long as the supply of labor comes from the traditional sector that can work for a fixed wage, economic growth occurs as a result of investments in the modern sector. This would ensure capital accumulation in the modern sector and cause income to be distributed unevenly between capital and labor favoring capital income.

3.3. Methodology

This study employs pooled mean group – autoregressive distributed lag model (PMG–ARDL). Cointegration tests based on the ARDL technique have several econometric advantages in comparison to the traditional approaches. The ARDL technique permits heterogeneity in the slope parameters and the cross–sectional dependency in relation to variables in the panels, to capture the long– and short–run associations between the target and independent variables. Furthermore, this approach is able to evaluate the presence of short– and long–run associations between the variables with both I(0) and I(1) or in their combinations, and deal with small sample bias.

The pervasive cross–section dependency problem may arise in panel data estimations if the units in a cross–section are correlated. In this context, before the design and analysis of PMG–ARDL, it is essential to check the cross–sectional dependence of the panel data to avoid spurious results. The existence of cross–sectional dependency is examined by applying CD test as suggested by Pesaran (2004). The CD test is a two–tailed technique following a standard normal distribution asymptotically under the null hypothesis that N and T tend to go to infinity in any order (Jensen and Schmidt, 2011). This test considers the mean pair–wise correlation coefficients from the residuals of the ordinary least squares regressions of Augmented Dickey–Fuller equations for each unit used in the analysis (Pesaran, 2007).

Checking the stationarity properties of the data series is the second step before estimating the econometric model, as misleading results may occur if non–stationary time series variables are used. Nevertheless, first–generation panel unit root tests may create inconsistent and biased results if units in a cross–section are correlated. For this reason, CIPS panel unit root test (Pesaran, 2007), which allows for the existence of cross–unit correlation, is conducted for the units determined to have cross–sectional dependency problem. The reasons for selecting this method are its flexibility in terms of allowable N and T values, convincing performance for small samples and being robust to cross–section dependence. Furthermore, since the order of integration of the variables is not important for the panel ARDL estimations, ADF (Dickey and Fuller, 1981), IPS (Im et al., 2003), LLC (Levin et al., 2002) and PP (Phillips and Perron, 1988) first generation panel unit root tests are performed for the units that passed the CD test to see whether all variables are I(0) and/or (1).

In panel data estimations, Pesaran and Smith (1995) are the first to propose applying a separate regression for each unit and averaging the results simply. This method, known as the Mean Group (MG), allows the coefficients to be heterogeneous and vary in the short– and long–run. However, the effects between units may converge in the long–run depending on the reasons such as the implementation of similar policies and strategies. In this case, utilizing the MG estimator may cause the long–run impacts to vary with individual differences. Pesaran et al. (1999) increased the GM to a general Autoregressive Distributed Lag model (ARDL) and subsequently remodeled it as a vector error correction form in order to overcome this issue. Consequently, the long–run coefficients are involved in the regression equation, the coefficients are pooled over the cross–section units and averaged to form a robust PMG estimator. The PMG estimator permits for heterogeneity in short–run coefficients and error variances as in MG but imposes homogeneity in long–run coefficients. The significance of the variations between PMG and MG estimators is examined by employing the Hausman test (Hausman, 1978). The findings supporting the h_0 hypothesis of this test poses the non–significant difference between the estimators and requires the PMG estimator to be preferred because of its reliability. Alternatively, the rejection of h_0 hypothesis indicates a significant difference between the estimators and MG is preferred in this case.

As reiterated above, this study employs PMG–ARDL technique. Within the framework of this approach, Equation 2 and Equation 3 are modelled for long– and short–run estimations, respectively.

$$lnGINI_{it} = \beta_1 + \sum_{i=1}^{k} \alpha_{i1} lnGINI_{j,t-i} + \sum_{i=0}^{k} \varphi_{i1} lnTROP_{j,t-i} + \sum_{i=0}^{k} \psi_{i1} lnFDI_{j,t-i} + \sum_{i=0}^{k} \omega_{i1} lnINF_{j,t-i} + \sum_{i=0}^{k} \theta_{i1} lnGOV_{j,t-i} + \sum_{i=0}^{k} \eta_{i1} lnCAP_{j,t-i} + \mu_{it1} \quad \mathbf{(2)}$$

$$\Delta lnGINI_{it} = \beta_2 + \sum_{i=1}^{k} \alpha_{i2} \Delta lnGINI_{j,t-i} + \sum_{i=0}^{k} \varphi_{i2} \Delta lnTROP_{j,t-i} + \sum_{i=0}^{k} \psi_{i2} \Delta lnFDI_{j,t-i} + \sum_{i=0}^{k} \omega_{i2} \Delta lnINF_{j,t-i} + \sum_{i=0}^{k} \theta_{i2} \Delta lnGOV_{j,t-i} + \sum_{i=0}^{k} \eta_{i2} \Delta lnCAP_{j,t-i} + \lambda \mathrm{ECT}_{j,t-i} + \mu_{it2} \quad \mathbf{(3)}$$

where ECT is the correction term and λ represents the coefficient of the ECT. The variables have gone through a logarithmic transformation.

4. EMPIRICAL FINDINGS

The outputs of the CD tests show that the h_0 of no cross–sectional independence for all parameters was not accepted for all panel models. Thus, it can be confidently expressed that there is cross–sectional dependency in each panel. Furthermore, CD tests were conducted for each of the variable in the panels individually to make it clear whether the cross–sectional dependency comes from residuals or not. The findings reveal the existence of cross–sectional dependency in all units except *lnGINI* in Panel–1, *lnTRO* in Panel–3, and *lnGOV* in Panel–1 and Panel–2 (Table 2).

Table 2. CD Test Results

Variable	Panel–1		Panel–2		Panel–3	
	CD–test	*p*–value	CD–test	*p*–value	CD–test	*p*–value
Panel	1.743**	0.045	–1.856**	0.043	3.924*	0.000
lnGINI	2.943	0.072	9.362*	0.000	12.838*	0.000
lnTROP	2.804*	0.005	9.574*	0.000	0.269	0.788
lnFDI	12.953*	0.000	5.336*	0.000	5.659*	0.000
lnINF	15.416*	0.000	7.483*	0.000	9.512*	0.000
lnGOV	–0.415	0.678	0.690	0.490	1.968**	0.049
lnCAP	14.271*	0.000	6.866*	0.000	8.872*	0.000

*and ** represent the levels of significance at 1% and 5%, respectively.

CIPS unit root test for the cross–sectionally dependent variables, and first–generation ADF, IPS, LLC and PP unit root tests for those cross–sectionally independent were conducted to determine stationarity properties. The findings of the unit root tests ran at the 0.05 significance level showed that the order of integration of the variables are mixed as I(0) and I(1). (Table 3, Table 4). In this case, applying static panel models might bring out invalid results. For this reason, Panel ARDL technique is preferred to investigate the short– and long–run relationships between the variables in the models.

Table 3. CIPS Panel Unit Root Test Results

	Variable	M1	*p*–value	M2	*p*–value
Panel – 1	*lnGINI*	–	–	–	–
	lnTROP	–2.426**	<0.05	–2.820	<0.10
	lnFDI	–2.925*	<0.01	–3.178*	<0.01
	lnINF	–3.698*	<0.01	–3.880*	<0.01
	lnGOV	–	–	–	–
	lnCAP	–2.586*	<0.01	–4.045*	<0.01
Panel – 2	*lnGINI*	–2.913*	<0.01	–2.611	>0.10
	lnTROP	–3.438*	<0.01	–3.388*	<0.01
	lnFDI	–2.565*	<0.05	–3.395*	<0.01
	lnINF	–4.263*	<0.01	–4.517*	<0.01
	lnGOV	–	–	–	–
	lnCAP	–2.851*	<0.01	–3.261*	<0.01
Panel – 3	*lnGINI*	–	–	–	–
	lnTROP	–3.134*	<0.01	–4.325*	<0.01
	lnFDI	–3.419*	<0.01	–3.548*	<0.01
	lnINF	–2.464**	<0.05	–3.317*	<0.01
	lnGOV	–2.387	<0.05	–3.225*	<0.01
	lnCAP	–3.082*	<0.01	–2.980*	<0.01

M1 and M2 refers to constant, and constant and trend forms, respectively.
* and ** imply the significance levels at 1% and 5%, respectively.
The optimal lag length selection is based on Schwarz information criterion.

Table 4. First Generation Panel Unit Root Tests Results

	Variable	ADF		IPS		LLC		PP	
		Statistic	*p*–value	Statistic	*p*–value	Statistic	*p*–value	Statistic	*p*–value
Panel–1	*lnGINI*	23.853	0.470	–0.025	0.490	–2.209**	0.014	22.961	0.522
	Δ*lnGINI*	53.686*	0.001	–3.482*	0.000	–2.786*	0.003	73.691*	0.000
	lnGOV	65.327*	0.000	–4.212*	0.000	–3.135*	0.001	52.168*	0.001
	Δ*lnGOV*	154.491*	0.000	–11.224*	0.000	–11.452*	0.000	239.635*	0.000
Panel–2	*lnGOV*	37.503*	0.000	–4.430*	0.000	–3.495*	0.000	26.212*	0.001
	Δ*lnGOV*	67.054*	0.000	–8.447*	0.000	–10.775*	0.000	75.597*	0.000
Panel–3	*lnTROP*	23.557**	0.023	–1.449	0.074	–0.713	0.238	20.691	0.055
	Δ*lnTROP*	95.634*	0.000	–6.468*	0.000	–5.824*	0.000	95.634*	0.000

Δ implies the first difference operator.
* and** stand for statistical significance at 1% and 5% levels, respectively.
Schwarz information criterion is used to identify the optimum lag length.

After determining that the unit root tests provided the stationarity properties of ARDL approach, the Hausman test was applied to select the most appropriate estimator for the ARDL model between MG and PMG. The results indicate that the χ^2 statistical coefficients are statistically significant at 0.05 level in Panel–1, and at 0.01 level in Panel–2 and Panel–3 (Table 5). In this case, the h_0, which predicts the homogeneity condition on the regressors in the long–run cannot be dismissed. Therefore, it is concluded that the PMG will be more efficient than the MG for Panel ARDL estimations. In other words, using the panel ARDL approach based on PMG to explore the relationship between the variables is found to be robust and appropriate.

Table 5. Hausman Test Results

	Panel–1		Panel–2		Panel–3	
	χ^2	p	χ^2	p	χ^2	p
Hausman	9.588**	0.048	15.917*	0.007	21.101*	0.001*

* and ** imply the statistical significances at 1% and 5% levels, respectively.

ECT coefficients ranging between –1 and 0 at 0,10 significance level (Table 6) indicate that all models correct themselves and converge back to equilibrium in the long–run in case of a short–run disequilibrium. However, although the coefficients are negative (Panel–1: ECT = –0.066; Panel–2: ECT = –0.103; Panel–3: ECT = –0.665), the evidence is relatively weak (Panel–1: p = 0.076; Panel–2: p = 0.094; Panel–3: p = 0.081).The findings of the PMG estimation results presented in Table 6 indicate a positive but weak relationship at 0.10 significance level (p = 0.062) between *GINI* and *TROP* for three of the panels in the short–run. Although the findings show that rises in *TROP* increase *ININ* in the former centrally planned middle–income economies in the short–run, the coefficients are statistically insignificant. In the long–run, the impact of *TROP* on *GINI* is negative for Panel–1 and Panel–3 indicating that increases in *TROP* reduce *ININ* in the middle–income CIS economies, and in the full sample. Nevertheless, the coefficients are still statistically insignificant. Unlike Panel–1 and Panel–3, there is a significant positive association at 0.01 level (p = 0.000) between *GINI* and *TROP* in Panel–2 in the long–run. In the EU member or candidate middle–income economies, 1% rise in the ratio of the sum of nominal imports and exports to GDP intensifies the *ININ* by increasing *GINI* by 0.268% in the long–run, which emphasizes that the benefits of *TROP* were not evenly distributed among the households. While a rapid process of *TROP* emerged in the EU member or candidate economies during the 1994–2019 period, it is likely that the development of welfare states and labor market institutions did not occur at the same pace in some of these countries (Dorn et al., 2022), which may be a driver of the worsening impact of *TROP* on *ININ*.

Table 6. PMG Estimation Results

	Panel–1 (2,1,1,1,1,1)		Panel–2 (2,1,1,1,1,1)		Panel–3 (2,1,1,1,1,1)	
	Coefficient	p–value	Coefficient	p–value	Coefficient	p–value
Short–run						
Δ*lnTROP*	0.002	0.450	0.012	0.094	0.642	0.522
Δ*lnFDI*	0.001	0.334	0.003	0.183	–0.274	0.784
Δ*lnINF*	0.000	0.236	5.002	0.260	–1.889	0.062***
Δ*lnGOV*	0.004	0.357	0.014	0.865	–0.181	0.857
Δ*lnCAP*	0.004	0.369	0.016	0.383	0.355	0.724
ECT	–0.066***	0.076	–0.103***	0.094	–0.665	0.081***
Long–run						
lnTROP	–20.672	0.929	0.268*	0.000	–6.633	0.845
lnFDI	–0.549	0.928	0.005*	0.014	–0.168	0.878
lnINF	–0.225	0.929	–0.003*	0.000	–0.248	0.859
lnGOV	–8.096	0.930	0.228*	0.000	2.537	0.844
lnCAP	16.167	0.928	0.179*	0.000	3.007	0.856

The optimal lag lengths of panels shown in parentheses rely on Schwarz information criterion.
*,**, and *** implies the statistical significance at 1%, 5%, and 10% levels, respectively.
ECT is the error correction term.

All panel model findings at 0.10 significance level show that the changes in the *FDI*, *GOV* and *CAP* insignificantly impact the *GINI* in the short–run. Although the results highlight that climbing inflation increases *ININ* in the short–run in Panel–1 and Panel–2, coefficients are statistically insignificant at 0.10 level. However, negative and statistically significant weak association (p = 0.062) between *GINI* and *INF* is revealed for Panel–3, whereby 1% increase in inflation brings along 1.889% decrease in *ININ* in the middle–income CIS economies in the short–run.

The findings of Panel–1 and Panel–3 imply that there is not any statistically significant liaison at 0.10 significance level amid *GINI* and the independent variables in the long–run. On the contrary, long–run PMG–ARDL regression coefficients of all independent variables are significant in Panel–2. The findings show that the association between *GINI* and *FDI* is positive at 0.05 significance level (p = 0.014) in the Balkan economies subsample. The PMG–ARDL long run estimations suggest that if the *FDI* increases by 1%, *ININ* rises 0.05%. It has also been found by Franco and Gerussi (2013), and Alili and Adnett (2018) for the EU transition economies; and by Neagu et al. (2016) for the Central and Eastern European transition economies that increases in *FDI* led to greater *ININ*. The coefficient on *INF* has a negative sign at 0.01 significance level for the former centrally planned middle–income Balkan economies. *ININ* decreases by 0.003%, by an additional 1% rise in inflation rate in these countries. Finally, government expenditures and gross fixed capital formation have significant impacts at 0.01 level on *ININ* in the sample of the middle–income Balkan economies in the long–run. Within 1% increase of the *GOV* and *CAP*, *ININ* worsens by 0.228% and 0.179%, respectively.

The findings of the PMG–ARDL model are not consistent with the H–O model and the Stolper Samuelson predictions. However, Braha–Vokshi et.al. (2021) provided evidence that *TROP* decelerated *ININ* in 6 former Yugoslavian states. On the other side, the outputs from the Balkan countries subsample are in line with the sizable body of previous literature which concluded the significant impact of *TROP* on *ININ* in transition economies (e.g., Franco and Gerussi, 2013; Neagu et al., 2016; Alili and Adnett, 2018; Cevik and Correa–Caro, 2020; Braha–Vokshi et al, 2021; Dorn et al., 2022). Nonetheless, considering the multivariate findings of the long–run ARDL estimations, it might be claimed that the Balkan countries included in Panel–2 have been at an early stage of the Kuznets curve. Likewise, Velkovska et al. (2021) argued that both the EU member former centrally planned Balkan countries and the EU candidate former Yugoslavian states had been experiencing the early stages of the Kuznets Curve. Sadiku et al. (2023) investigated the macroeconomic determinants of *ININ* by employing a least–squares dummy–variable model. The authors concluded that the Central and Eastern European countries had not reached the peak of Kuznets curve.

5. CONCLUSION

This study explored the association between trade openness and income inequality in the 12 former centrally planned middle–income economies by employing Panel ARDL technique. Foreign direct investment, inflation, government expenditures and gross fixed capital control variables were incorporated into the model. Apart from the full sample, the EU member or candidate middle–income economies, and the middle–income CIS countries subsamples were used in estimations as benchmark.

Initially, a CD test was employed to examine whether all units in the same cross–section were correlated or not. The results validate that the determinants in the model impact each other. CD test findings are consistent with the paths Panel–1 and Panel–2 countries took to achieve globalization. Although the countries examined in this study are all former centrally planned middle–income economies, the countries included in Panel–2 and Panel–3 differ from each other depending on various factors such as geographical proximities, historical connections and political ties. EU member and candidate countries in Panel–2, which are all located in Balkans, should adopt the criteria of the Community Acquis and ensure the good neighborly relations conditions of participation in order to comply with the decisions taken by the EU and the regulations it implemented. The integration of these countries into the EU criteria for membership enabled them to deal with the challenges of the transition from a centrally planned economy to a market–based system much better than the CIS economies. On the other hand, the CIS countries in Panel–3, which are all located in Eurasia, appeared poor to implement institutional reforms and progress toward creating a market economy. The CIS countries faced various challenges during transition to market based economy, such as the termination of large fiscal transfers, the increase of monopoly power, low–skilled and unskilled labor, weak trade–related institutions, and the ongoing political and economic influences of Russia.

Subsequently, first and second–generation unit root tests were conducted and the order of integration of the variables were found to be a mix of I(0) and I(1) which allowed Panel ARDL technique to be used in this study. The results of the Hausman test performed prior to the application of the Panel

ARDL specified that PMG estimator could offer robust outputs. The PMG estimator allowed to identify the short– and long–run impacts of trade openness and control variables on income inequality in transition economies.

The estimation results did not confirm that trade openness impact income inequality in the former centrally planned middle–income economies in the short–run. For the subsample of the middle–income CIS countries and the full sample, a similar result was obtained in the long–run revealing that there was no statistically significant association between trade openness and income inequality. However, the coefficient was positive, and significant at 0.01 level in the subsample including the middle–income EU member or candidate economies. In the examined period, a 1% increase in trade openness worsened income inequality by 0.268% in the long–run. The fact that trade openness has increased income inequality does not mean that the former centrally planned middle–income Balkan economies should shift to inward–oriented trade strategies to reduce income inequality. While following their strategies that promote integration into the global market, these countries should pursue income redistribution policies to reduce income inequality.

In terms of the effects of the control variables on income inequality, no significant coefficient at 0.10 level was obtained in the full sample, either in the short– or long–run. Nevertheless, it was determined that all control variables had significant impacts at 0.05 level on income inequality in the middle–income Balkan economies in the long–run. Increases in the ratios of net inflows of foreign direct investment, total expenditures of government, and gross fixed capital formation to GDP worsened income inequality. Although inflation had a coefficient with a significant negative sign, its effect on reducing income inequality was negligible. For the middle–income CIS economies, it was revealed that rise in inflation increased income inequality in the short–run whereas no significant relationship was observed between the control variables and income inequality in the long–run.

The long–run estimation coefficients related to the panel of the former centrally planned middle–income Balkan economies revealed that income inequality was significantly shaped by all independent variables included in the model. On the contrary, the long–run analysis outcomes of the panel including the CIS countries showed that none of the independent variables had significant influences on income inequality in the period under investigation. The contrasting test results are likely based on the dissimilarities between the Balkan countries in Panel–2 and the CIS countries in Panel–3 in terms of initial conditions, reform strategies and political frameworks, impacting transition.

This study has contributed to literature by identifying the liaisons amid income inequality and trade openness in the former centrally planned middle–income economies and as benchmark in its two subgroups. However, it has limitations in some respects, such as sample size and potential variables excluded. Firstly, some of the former centrally planned middle–income economies had to be excluded since the complete data set was unavailable. Moreover, the components of trade could not be included in the model, because of the non–availability of full data set once, and the estimations were solely based on the aggregate measure of trade openness as extensively used in the literature. Thirdly, it could not be detailed in which income groups and to what extent the impacts of trade openness on income inequality occurred. Furthermore, the model does not include technological progress since we could not reach the data of the variables measuring the development of technology. Incorporating the variables mentioned above into the regressions with a complete data set will provide more information about the nature of the link between trade openness and income inequality for the former centrally planned middle–income economies.

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