WILL INCOME INEQUALITY DIFFERENCES AMONG THE EURO AREA AND FORMER SOCIALIST EUROPEAN COUNTRIES WIDEN? A PANEL DATA ANALYSIS

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ABSTRACT

Unfair distribution of income and poverty has been one of the most serious problems in the world economy since wind of globalization has become prevalent. In order to develop polices improving income distribution reducing income differences among countries, it is critically important to incorporate the role of economic, social and demographic factors. In this regard, we estimate the possible effects of economic, demographic and social indicators on income differences between Euro area and the former socialist countries of Europe with panel GMM model. Our results emphasize the importance of demographic factors to examine the dynamics of income equalities. It is also implied that monetary economic policies should be coordinated to foreign trade and development to sustain economic development and reduce income difference with respect to the Euro area in former socialist European countries.

Keywords: Income inequality, Panel Data Analysis, Former Socialist European countries.
JEL Codes: O1, O5.

1. INTRODUCTION

The Central and Eastern European countries (CEE) have undergone a number of economic, political and institutional changes during the transition period from centrally planned economy to market economy since 1990s. Most of these countries tended to implement adaptation policies regarding currency stabilization, monetary measures and fiscal discipline as well as privatization programmes towards restructuring the economy. These attempts were followed by extensive shocks to macroeconomic fundamentals which also affected the process of economic development. Transition economies then went through the process of the access to the European Union (EU) and European Monetary Union (EMU) which require the convergence of inflation, interest rates, exchange rates and government deficits to the average levels of the EU countries. Accordingly, candidate CEE countries implemented a number of reform programmes in order to capture the extensive benefits from catching up the developed EU countries. However, during the catching up and adaptation process, CEE countries have been expected to face with fluctuations and macroeconomic shocks which could be severe in less developed countries and also lead to disparities in growth performances and income equality among these countries and the EU countries. The issue whether the transition to market economies from centrally planned economy and adaptation to the EU have led the economies of these countries to diverge or converge has been extensively analyzed by many researchers.

Economic convergence occurs when macroeconomic conditions and also other development indicators of country groups under estimation close up so as to reduce disparities in the growth levels and per capita income levels among countries. The issue is generally examined with its two main aspects, (1) income equalization and convergence of development levels of countries, and (2) convergence of business cycles among countries. The literature on economic convergence dates back to the arguments on traditional international trade theories which stress upon the convergence effect of international trade incorporated in factor returns and relative price equalization. However, the theoretical explanation on economic convergence has mainly been developed within the neoclassical (Solow, 1956) and then endogenous growth theories. The main assertion of economic growth theories is that less developed countries with a lower GDP per capita will grow faster than developed ones due to higher return to capital in less developed countries which attract foreign capital inducing faster growth (Solow, 1956; Mankiw et al., 1992). In the models, the economic convergence is measured by β and σ convergence. β convergence, unconditional and conditional, measures the correlation between per capita output levels and growth rates in countries under estimation. Unconditional β convergence measures the explanatory power of initial per capita income over growth rates while the indicators such as education, health and other policy variables included as explanatory variables into the growth regression indicate conditional β convergence. The other measure is σ convergence which signals the reduction in income difference among countries (Barro and Sala-i-Martin, 1992; Quah, 1993; Sala-i-Martin, 1996). However, endogenous growth models put forward that due to different initial conditions and other peculiar factors of different countries, income levels cannot converge.

Also, Romer (1986) and Lucas (1988) support the inequality effect of integration on economic convergence since the research and development and brain drain. The newer theories imply that new opportunities brought by greater economic integration also contribute to divergence in growth and income among countries (Krugman, 1991).

From the empirical side of the issue, the factors affecting the distribution of income and the economic implications of income inequality have been the major topics in development economics. Many researchers have been applying quantitative models to expose dynamics of income equalities and thus shed light to policy makers. Within this context, labor markets, supply responses of workers, macroeconomic policies, redistribution policies, minimum wage legislation and policy jurisdiction have been analyzed. This paper aims to contribute to the ongoing argument on convergence outcome of economic integration regarding income equalization among CEE countries which have joined the EU. Income convergence will be analyzed by economic, social and demographic factors since these factors regarding the economic convergence of CEE countries are greatly related to EU markets, namely to trade and capital inflows, transfer of technology and labor markets.

The paper is organized as follows: Section 2 outlines the empirical studies briefly. Section 3 describes the materials and econometric methodology used in testing for income inequality convergence. Section 4 presents the empirical results and Section 5 concludes the paper.

2. PREVIOUS RESEARCH

One of the earlier contributions to the literature was by Dollar and Kraay (2004) who implied that globalization caused to faster growth in poor countries and also led to poverty reduction in these countries. However, Winters, McCulloch and McKay (2004) revealed that in the long-run trade liberalization was likely to be strongly poverty alleviating and it would increase overall poverty. Bergh and Nilsson (2010) applied panel GMM models to analyze whether the KOF Index of Globalization and the Economic Freedom Index of the Fraser institute were related to each other within country income inequality using panel GMM modeling for 80 countries. They exposed that freedom to foreign trade, social globalization and deregulation might be related to inequality in rich and less developed countries, whereas monetary reforms, legal reforms and political globalization did not have distorting effects on inequality. On the other hand, tariffs to trade in terms of reduction on final quality-differentiated goods and a reduction on intermediate goods may have different impacts on wage inequality according to Ma and Dei (2009) who found that reducing the tariff on final quality differentiated goods have opposite effects on welfare inequality and wage inequality. Relationships between openness and income inequality may also differ; Jalil (2011) investigated the long-run relationship between these two variables in openness Kuznets curve framework with the Auto Regressive Distributed Lag (ARDL) estimator. He obtained parallel results to the Kuznets hypothesis implying that income inequality rises with the increase of openness and then started to fall.

Most recently, Asteriou, Dimelis, Moudatsou (2014) investigated the relationship between globalization and income inequality with panel data analysis for EU-27 countries sub grouped as the Core, Periphery, High Technology, and the New EU Member. Their results showed that trade openness had a positive impact on income inequality in EU-27 countries, whereas financial globalization, capital account openness and stock market capitalization were major factors increasing income inequality in these countries. Thus, it can be inferred that expansion of foreign trade, along with the increase in external competitiveness may be the driving force of fixing the inequality in income distribution in EU-27 countries. Wu and Hsu (2012) provided different outcomes for the effects of FDIs on income inequality in their study using the endogenous threshold regression model for 54 countries. Their results showed that FDI was likely to be harmful to the income distribution of those host countries with low levels of absorptive capacity, whereas FDI had limited impacts on income inequality in the case of countries with better absorptive capacity. On the other hand, findings of Asteriou, Dimelis, Moudatsou (2014) imply that monetary policies improving financial development and reduction of financial fragility may have effects improving income inequality. Lessmann (2013) obtained parallel results to Asteriou, Dimelis, Moudatsou (2014) with panel data techniques, showing that foreign direct investments might increase regional inequality in low and middle income countries, while they had no negative redistributional effects in high income economies. Lessmann (2013) also indicated that foreign direct investments might increase regional inequality since many different regions of a country usually could not receive FDI in equal measure.

Feenstra and Hanson (1995) stated that growth in FDI was a factor leading to a demand push to the skilled labor and thus increase in the skilled labor share of total wages. Alderson and Nielsen (1999) employed random-effects regression models incorporating the role of unmeasured country heterogeneity to study effects of foreign capital penetration on inequality. They suggested that the relationship between income inequality and investment dependence should be examined within the context of inflow and outflow of foreign capital to economic development. Chiquiar (2008) revealed that overall wage levels increased in the regions of a country that were more exposed to globalization, moreover it is stressed that trade liberalization had a spatial dimension. Zhao (2001) put forward that FDI could increase relative wages of skilled labor even without bringing in skill-biased technology. On the other hand, Ravallion (2003) exposed that within-country income inequalities had been slowly converging since the 1980s; inequality had the trend of fall (rise) in countries with initially high (low) inequality. Bleaney and Nishiyama (2003) revealed that inequality convergence was slower amongst developing countries than amongst OECD countries. Ezcurra and Pascual (2005) detected that there was a process of convergence in regional inequality levels in among the European regions and they asserted the importance of the national component in explaining the dynamics of regional inequality distribution. Similarly, Gomes (2007) showed that Brazilian municipalities are converging to an inequality level greater than the year 2000 level, while Lin and Huang (2012) provided empirical evidence in support of convergence in income inequality from Panel LM unit-root tests for the 48 contiguous states in the US over the 1916–2005 period.

According to Wei and Wu (2001) openness is also critical factor for the urban-rural income inequality. In their study, they found empirical evidence from Chinese cities for the fact that cities witha greater degree of openness in trade might exhibit a greater decline in urban-rural income inequality. However, Wei and Wu (2001) emphasized that a negative association between openness and inequality might be detected when a geography-based instrumental variable to correct for possible endogeneity of a region's trade. Dobson and Ramlogan (2008) found a non-linear relationship between trade openness and inequality, showing that inequality increased until a critical level of openness was reached after which inequality began to fall. Thus, Dobson and Ramlogan (2008) suggested that redistribution policies and trade liberalization measures can be conducted to overcome the negative effects of trade liberalization. Song (2013) suggested that fiscal policies (spending decentralization, revenue decentralization and autonomy power) might have a determinative role on regional income inequality. Liu et.al (2014) considered the role of social fiscal policies focusing on rural households of China. More specifically, effects key priority forestry programs on rural households' inequality are explored with fixed-effects model panel data model. They stressed the fact that overall effects of social programs to rural population might differ from these policies' direct effects because of the rural households using more capital inputs for their farmland. In addition, education and local road quality can be considered as key factors increase non-agricultural labor supply and income in remote areas, and thus improve inequatility (Yamauchi, 2011). Inequality may increase as financial sector development increases at very low levels of financial sector development; however inequality was less when financial development was greater in the long-run (Clarke, Xu and Zou, 2006). Beck, Demirgüç-Kunt and Levine (2007) emphasized that financial system had a major impact on either through the reductions in income inequality or through the impact of financial development on aggregate economic growth. On the contrary, Li and Zou (1998) found that inequality was positively and significantly associated with economic growth from baseline estimations and a sensitivity analysis show income for developed and developing countries.

As for the empirical analysis of convergence among CEE countries, Kocenda (2001) explored that there has been a considerable tendency to convergence in macroeconomic Fundamentals of 11 CEE countries under estimation (Czech Republic, Slovakia, Hungary, Poland, Slovenia, Romania, Bulgaria, Albania, Estonia, Latvia, and Lithuania) in a panel setting. Kutan and Yigit (2005) expand the study of Kocenda (2001) using panel unit root techniques in an attempt to examine the convergence of the new EU members to EU standards. Their results show that for the Baltic states, there is a strong monetary policy and price-level convergence while CEE5 countries do not have convergence mainly due to the lack of fiscal discipline. However, in a study testing growth specifications of transition economies developed by Barro (1991) and Levine and Renelt (1992), Campos (2001) found that the former centrally planned economies differ from market economies at similar levels of per capita income. Matkowski and Prochniak (2004) analyzed income convergence among 8 CEE countries regressing GDP growth rates on GDP per capita levels and also cyclical convergence using industrial production indexes and industrial confidence indicators over the period 1993-2004. Their results reveal that there is an income convergence among CEE countries and between CEE countries and the EU. Besides, these countries exhibit a strong cyclical synchronization with the EU. Amplatz (2003), testing both β and σ convergence types for the period 1996-2000, found that most of the CEE accession candidates showed all types of economic convergence among themselves but there was not convergence between these countries and Western Europe.

3. MATERIALS AND METHODS

In this study, we attempted to analyze the role of economic, social and demographic factors that may influence the Gini coefficient in the former socialist countries of Europe (Bulgaria, Czech Republic, Croatia, Hungary, Macedonia, Poland, Romania, Serbia, Slovakia and Slovenia) with panel data analysis. We applied to the statistical database of the World Bank for the needed data, however, Gini coefficent is proxied by the GINI Index have missing values for Bulgaria, Czech Republic, Croatia, Hungary, Macedonia, Poland, Romania, Serbia, Slovakia and Slovenia, thus the empirical analysis is carried by using Gross National Income (GNI) per capita[1]. Due to the availability of data for all the countries, effects of unemployment -as percentage share of total employment- ($unemp$), male unemployment -% of male labor force- ($unempml$), female unemployment -% of female labor force- ($unempfm$), consumer price inflation (cpi), GDP growth rate ($gdpg$), tax rate -as a percentage share of total profits- ($taxr$), health expenditure -as a share of GDP- (hea), degree of openness ($open$)[2], foreign direct investments (fdi), population of aged between 15 and 64 -as a percentage share of the total- (pop), rural population -as a percentage share of the total- ($rpop$), urban population -as a percentage share of the total- ($upop$), adolescent fertility rate -births per 1.000 women ages 15 and 19- ($fert$) and internet users -per 100 people- ($inet$) on gross national income per capita differences between Euro area and former socialist countries of Europe ($dgni$)[3] are studied. The time series of the variables of related

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[1]GNI per capita is the gross national income, converted to U.S. dollars using the World Bank Atlas method, divided by the midyear population. Atlas method uses a a conversion factor that averages the exchange rate for a given year and the two preceding years, adjusted for differences in rates of inflation between the country, and through 2000, the G-5 countries, see (World Bank Statistical Database).

[2] We compute the degree of openness to trade as; (exports+imports)/GDP.

[3] We compute the gross national income per capita differences between Euro area and former socialist countries of Europe as; $dgni = gni_t^{eur} - gni_t^{fsce}$; where gni_t^{eur} denotes the gross national income per capita of the Euro area and gni_t^{fsce} refers to the gross national income per capita of each former socialist countries of Europe. gni_t^{eur} and gni_t^{fsce} are in logarithms.

countries are for the period from 2005 to 2011 and they are pooled to estimate the impacts on income differences.

3.1. Panel Unit Root Analysis

Panel unit root test have a theoretical structure parallel but not identical to unit root tests of single time *series* data. Restrictions that can be placed on the autoregressive process across cross-sections or series determine the specification of the panel unit testing. In this context, panel data having AR(1) process is expressed as below;

$$y_{it} = \rho_i y_{it-1} + X_{it}\delta_i + \varepsilon_{it} \tag{1}$$

where $i = 1, 2, ..., N$ cross-section units or series for the periods $t = 1, 2, ..., T_i$. X_{it} denotes exogenous variables in the model with any fixed effects or individual trends. ρ_i are the autoregressive coefficients of the model and ε_{it} are the error terms that are assumed to be mutually independent idiosyncratic disturbance. If $|\rho_i| < 1$ then y_{it} is weakly (trend-) stationary, while y_{it} has a unit root if $|\rho_i| = 1$. The Levin, Lin, and Chu (LLC), Breitung, and Hadri tests all employ the panel unit root test by assuming that the persistence parameters are common across cross-sections, $\rho_i = \rho$ for all i. On the other hand, the Im, Pesaran, and Shin (IPS), and Fisher-ADF and Fisher-PP tests assume that ρ_i varies across cross-sections (E-Views 8 User Guide II, 2013: 487). In order to specify the appropriate type of the model, we applied panel unit root tests with different assumptions and their results are reflected in Table 1.

Table 1: Panel Unit Root Tests Results

	Levin, Lin, and Chu		Im, Pesaran, and Shin		Fisher-ADF		Fisher-PP	
	Statistic	Prob	Statistic	Prob	Statistic	Prob	Statistic	Prob

unemp	-2,74	0,00	0.42	0.66	15,85	0,72	23,57	0,26
Δ*unemp*	-2,62	0,00	0,68	0,75	11,68	0,75	9,67	0,97
unempml	-2,77	0,00	0,50	0,69	15,28	0,75	21,48	0,36
Δ*unempml*	-2,91	0,00	0,48	0,68	13,23	0,86	11,51	0,93
unempfm	-2,04	0,02	0,27	0,60	17,62	0,61	22,27	0,32
Δ*unempfm*	-2,76	0,00	0,54	0,70	13,04	0,87	11,02	0,94
cpi	-6,89	0,00	-1,64	0,04	33,65	0,02	50,10	0,00
Δ*cpi*	-8,43	0,00	-2,13	0,01	39,31	0,00	59,64	0,00
gdpg	-4,32	0,00	-0,34	0,36	19,65	0,47	18,50	0,55
Δ*gdpg*	-8,21	0,00	-1,93	0,02	37,32	0,01	52,88	0,00
taxr	-7,00	0,00	-0,65	0,25	28,65	0,09	40,69	0,00
Δ*taxr*	-7,93	0,00	-2,01	0,02	35,25	0,00	51,51	0,00
hea	-3,40	0,00	0,26	0,60	15,29	0,75	24,66	0,21
Δ*hea*	-4,61	0,00	-0,52	0,29	22,75	0,30	29,60	0,07
open	-4,32	0,00	-0,34	0,36	19,65	0,47	18,50	0,55
Δ*open*	-8,21	0,00	-1,938	0,0263	37,32	0,10	52,88	0,00
fdi	-3,253	0,000	-0,298	0,382	20,25	0,44	20,53	0,42
Δ*fdi*	-6,893	0,000	-1,593	0,05	33,23	0,31	45,38	0,00
pop	0,32	0,62	3,04	0,99	9,53	0,97	28,60	0,09
Δ*pop*	9,82	1,00	1,22	0,88	10,18	0,85	10,24	0,85
rpop	-3,50	0,00	-0,03	0,48	32,13	0,04	41,48	0,00
Δ*rpop*	0,04	0,51	1,78	0,96	11,31	0,93	16,59	0,67
upop	0,41	0,65	2,98	0,99	10,56	0,95	28,21	0,10
Δ*upop*	-2,12	0,01	-2,31	0,01	28,75	0,02	28,29	0,02
fert	-1,86	0,03	3,37	0,99	19,82	0,46	43,68	0,00
Δ*fert*	-4,06	0,00	0,41	0,66	13,13	0,87	19,83	0,46
inet	-1,98	0,02	0,83	0,79	15,22	0,76	40,61	0,00
Δ*inet*	-5,88	0,00	-1,31	0,09	30,56	0,06	34,47	0,02
dgni	-7,42	0,00	-1,08	0,13	28,75	0,09	55,83	0,00
Δ*dgni*	-1,84	0,03	0,77	0,78	12,06	0,91	11,86	0,92

According to Table 1, all variables can be treated as either stationary or non-stationary even at the 10 percent significance level. Thus, we did not explore the possibility of cointegration relationships among the using panel cointegration tests. We ignored the stationarity of the variables and a dynamic panel data model (Arellano–Bond Dynamic Panel GMM Estimators) is used.

3.2. Panel GMM Model

The point of departure of our analysis in this study is the linear panel data regression model specified as below;

$$Y_{it} = \alpha + \beta X_{it} + \varepsilon_{it} \tag{2}$$

where Y_{it} and X_{it} refer to the dependent and independent variables of the model, respectively. Y_{it} and X_{it} have both i and t subscripts for $i = 1, 2, ..., N$ sections and $t = 1, 2, ..., T$ time periods. α and β are coefficients of the model with no subscripts, pointing that they will be same for all unit and samples. Finally, ε_{it} denotes the error term of the model. Estimation of the common constant method (pooled OLS method) as in (2) presents results under the assumption that there are no differences among the data matrices of the cross-sectional dimension N; more precisely the model estimates a common constant α for all cross-sections (Asteriou and Hall, 2007: 345).

On the other hand, the error term ε_{it} determine whether the model may have fixed effects or random effects. Similar to a dummy variable model in one dimension, it is assumed that the error term ε_{it} varies non-stochastically over i and t in fixed effects model. In a random effects model, the error term is assumed to be varying stochastically. Within this framework, type of models as in (3) can be estimated using a pool object.

$$Y_{it} = \alpha + X'_{it}\beta_{it} + \delta_i + \gamma_t + \varepsilon_{it} \tag{3}$$

In (3), Y_{it} refer to the dependent variable, while X_{it} is a vector of k regressors, and ε_{it} are the error terms for cross-sectional units, $t = 1,2,...,T$. α is the constant term and cross-section or period specific effects (random or fixed) are denoted by δ_i and γ_t. In order to identify the panel data model, restrictions placed on β coefficients [common (across cross-section and periods), cross-section specific, and period specific regressor parameters] are needed (E-Views 7 User Guide, 2010: 601). For instance, M cross-sectional equations each with T observations stacked on top of one another can be expressed as below.

$$Y_{it} = \alpha I_T + X_{it}^{'}\beta_{it} + \delta_i I_T + I_T\gamma_t + \varepsilon_t \quad \text{for } i = 1,2,...,M \tag{4}$$

where I_T refers to the $T-$ element identity matrix and all the period effects $\gamma^{'} = (\gamma_1, \gamma_2, ..., y_T)$ are included in vector γ (E-Views 7 User Guide, 2010: 602). Similar to (4), we can specify as a set of T period specific equations, each with M observations stacked on top of one another as in (5);

$$Y_{it} = \alpha I_T + X_t^{'}\beta_{it} + I_M\delta_i + \gamma_t I_M + \varepsilon_t \quad \text{for } i = 1,2,...,M \tag{5}$$

where I_M is the $M-$ element identity matrix, vector-δ has all of the cross-section effects $\delta^{'} = (\delta_1, \delta_2, ..., \delta_T)$ (E-Views 7 User Guide, 2010: 602).

Accordingly, if all of the β_{it} are common across cross-sections and periods model (3) can be written as; $Y_{it} = \alpha + X_{it}^{'}\beta + \delta_i + \gamma_t + \varepsilon_{it}$, and thus there are k coefficients in each β corresponding to an element of x. On the other hand, when all of the β_{it} coefficients are cross-section specific, we can specify the model (3) as; $Y_{it} = \alpha + X_{it}^{'}\beta_i + \delta_i + \gamma_t + \varepsilon_{it}$ and thus there are k in each β_i for a total of Mk slope coefficients[4]. In the third case, all of the β_{it} coefficients are period specific so that model (3) takes the form $Y_{it} = \alpha + X_{it}^{'}\beta_i + \delta_i + \gamma_t + \varepsilon_{it}$ for a total of Tk slope coefficients (E-Views 7 User Guide, 2010: 603).

[4] The fixed effects estimator is known as the least-squares dummy variables (LSDV) estimator since it includes a dummy variable for each group to allow for different constants.

On the other hand, depending on the panel model specification in (3), GMM panel estimators can be specified as below;

$$g(\beta) = \sum_{i=1}^{M} g_i(\beta) = \sum_{i=1}^{M} Z_i^{'} \varepsilon_i(\beta) \tag{6}$$

where Z_i is $T_i \times p$ matrix containing instruments for cross-section i, and $\varepsilon_i(\beta) = (Y_i - f(X_{it}, \beta))$. Panel GMM model estimation objects to minimize the quadratic form; $S(\beta) = g(\beta)^{'} Hg(\beta)$, with respect to β and $p \times p$ weighing matrix H. $\hat{\beta}$ represents the estimates of the coefficient vector and he coefficient covariance matrix of the panel GMM model can be computed as below;

$$V(\hat{\beta}) = (G^{'}HG)^{-1}(G^{'}H\Lambda HG)(G^{'}HG)^{-1} \tag{7}$$

In (7), Λ is an estimator of $E(g_i(\beta)g_i(\beta)^{'}) = E(Z_i^{'}\varepsilon_i(\beta)\varepsilon_i(\beta)^{'}Z_i)$, while G is a $T_i \times k$ derivative matrix. GMM estimation requires specifying the instruments, choosing the weighting matrix H and determining an estimator for Λ (E-Views 8 User Guide II, 2013: 792-793).

4. EMPIRICAL RESULTS

In this study, we estimated a panel GMM model based on Arellano and Bond (1991) to analyze effects of unemployment, GDP per capita, inflation, tax rate, openness to trade, foreign direct investments, rural population, population and health expenditures on the gross national income per capita differences between Euro area and former socialist countries of Europe. We used the first and second lags of the independent variables of the model as instruments to make the endogenous variables pre-determined and thus not correlated with the error term in equation. In order to eliminate the fixed effect in the model, orthogonal deviations are employed as a transformation method and Period SUR instrument is used as the weighing matrix. According to Table 2, over identifying restrictions of the model are also valid and satisfied since J-statistic has a p-value as 0,32. Estimation results are presented in Table 2.

Table 2: GMM Model Estimation Results

Method: Orthogonal Deviations
Estimation weighting matrix: Period SUR instrument
Instrument specification: c, $dgni(-2)$, $unemp(-1)$, $unempml(-1)$, $unempfm(-1)$, $gdpg(-1)$, $cpi(-1)$, $taxr(-1)$, $open(-1)$, $fdi(-1)$, $hea(-1)$, $pop(-1)$, $rpop(-1)$, $upop(-1)$, $fert(-1)$, $inet(-1)$, $unemp(-2)$, $unempml(-2)$, $unempfm(-2)$, $gdpg(-2)$, $cpi(-2)$, $taxr(-2)$, $open(-2)$, $fdi(-2)$, $hea(-2)$, $pop(-2)$, $rpop(-2)$, $upop(-1)$, $fert(-1)$, $inet(-1)$.
J**-Statistic:** 19,06
Instrument Rank: 32

Variable	Coefficient	Prob.
$dgni(-1)$	0,6279	0,0000
$unemp$	0,0093	0,0285
$unempml$	-0,0039	0,2915
$unempfm$	-0,0014	0,7560
cpi	-0,0037	0,0096
$gdpg$	-0,0026	0,0036
$taxr$	-0,0012	0,3669
hea	0,0148	0,0710
$open$	0,0015	0,0002
fdi	$-1{,}56\times10^{-13}$	0,1434
pop	-0,7848	0,3138
$rpop$	$4{,}97\times10^{-5}$	0,7050
$upop$	-0,0007	0,3091
$fert$	0,0108	0,0021
$inet$	-0,0003	0,5902

Table 2 shows that $gdpg$ and cpi has negative and statistically significant coefficients at the 10% level as; -0,0030 and -0,0029 implying that expansionary economic policies may reduce the income gap between the Euro area and the former socialist countries of Europe.

Thus, it can be inferred that macroeconomic development may be sustained and unemployment rate may be reduced in the long-run with expansionary policies. Our policy implications for increasing the GNI per capita in the former socialist countries of Europe are also consistent to the Keynesian Theory. The statistically significant coefficient of *unemp* (0,0093) is supporting our inferences; an increase in the unemployment rate in the former socialist countries of Europe deteriorates the total demand which in turn may increase the income equality with respect to the Euro area. The statistically insignificant coefficients of male and female unemployment rates does not allow us to make interpretations on the effects of decreases in male and female unemployment on the GNI per capita differences between the Euro areas and the former socialist countries of Europe.

The negative coefficient of *taxr* exposes that increasing tax rates may affect the GNI per capita in the former socialist countries of Europe positively and thus reduce the income difference with the Euro area, however *taxr* has a statistically insignificant coefficient. Thus, it is difficult to make inferences about the outcome of tax policies on income inequality in these countries. On the other hand, our estimation results reveal that health expenditures increase the GNI per capita inequality between the Euro area and the former socialist countries of Europe. Despite it is regarded that social expenditures have importance for sustaining the development in the long-run, we assert health expenditures affect that total demand negatively in the former socialist countries of Europe. Considering the fact that health expenditures are being financed by the government in these countries, increases in these kinds of expenditures reduce the expenditures boosting the aggregate demand under the governments' budget constraint. Empirical results also indicate that increase in the population of aged between 15 and 64 of the former socialist countries of Europe can be accepted as a raise in the labor participation rate. This phenomenon may influence the aggregate demand positively and become a factor reducing the income gap between the Euro area and the former socialist countries of Europe. However, we can't make inferences about the outcome of population of aged between 15 and 64, rural population and urban population on GNI per capita differences since they have statistically insignificant coefficient. Fertility rate has a increasing impact on GNI per capita differences between the Euro area and the former socialist countries of Europe since the statistically significant coefficient is found as 0,0021. Accordingly, it can be inferred that increase in the birth rates had a negative on the GNI per capita differences, whereas effects of increase in the birth rates becomes positive in the long-run.

Openness to trade has statistically significant and positive coefficient according to our GMM estimation results. Thus, it can be inferred that liberalizing foreign trade policies may have a negative impact on GNI per capita in the former socialist countries of Europe. In this regard, we emphasize that structural policies and reforms on micro basis should be conducted for increasing the competitiveness of these countries. On the other hand, we do not obtain precise results for the effects of foreign direct investments on GNI per capita in the former socialist countries of Europe since has statistically insignificant coefficient. Structural policies and reforms on micro basis can also become factors causing positive effects on GNI per capita as a result GNI per capita. In this respect, our results suggested that an increase in the number of internet users may have a reducing impact on GNI per capita between the Euro area and the former socialist countries of Europe. However, the coefficient of is *inet* statistically insignificant.

5. CONCLUSIONS

Panel GMM model based on Arellano and Bond (1991) is used as an estimation strategy to analyze the effects of unemployment, GDP per capita, inflation, tax rate, openness to trade, foreign direct investments, rural population, population and health expenditures on the gross national income per capita differences between Euro area and former socialist countries of Europe. Our empirical results imply that expansionary macroeconomic policies in the former socialist countries of Europe can be implemented to reduce the income difference with respect to the Euro area. However, the sign and the p-values of the coefficients regarding to fiscal policy (hea and $taxr$), did not support the suggestion that expansionary fiscal policies should be implemented. Thus, the importance of expansionary monetary policies for the former socialist countries of Europe is stressed for reducing the income differences with the Euro area. For reducing the risks that may occur in the financial system as a result of the expansionary monetary policies, monetary authorities of these countries should implement policies aiming financial stability in parallel.

On the other hand, our findings may imply that the major role of increasing labor force and employment for boosting the aggregate demand and thus reducing the GNI per capita gap with respect to the Euro area. Along with the increase in the population, it is important to finance investments generating employment. Thus, comprehensive development plans including all sectors of economy should be adopted. These plans should incorporate measures to increasing the competitiveness of the economy. In this respect, conduction of innovation and technology polices have critical importance. Implications about the competitiveness of the economy are also supported by the sign and the p -values of $open$ and fdi.

REFERENCES

Alderson, A.S., Nielsen, F., 1999. Income inequality, development, and dependence: A reconsideration, American Sociological Review. 64, 606-631.

Amplatz, C., 2003, The economic convergence performance of central and eastern european countries, Economics of Planning. 36, 273-295.

Arellano, M., Bond, S., 1991, Some tests of specification for panel data: Monte carlo evidence and an application to employment equations, The Review of Economic Studies. 58, 277-297.

Asteriou, D., Hall, S.G., 2007, Applied econometrics A modern approach using microfit and eviews revised edition, Palgrave MacMillan, New York.

Asteriou, D., Dimelis, S., Moudatsou, A., 2014. Globalization and income inequality: A panel data econometric approach for the EU27 countries, Economic Modelling. 36, 592–599.

Barro, R., 1991, Economic growth in a cross section of countries, The Quarterly Journal of Economics. 106, 407-443.

Barro R., Sala-i-Martin X., 1992, Convergence, Journal of Political Economy. 100, 223-251.

Beck, T., Demirgüç-Kunt, A., Levine, R., 2007, Finance, inequality and the poor, Journal of Economic Growth. 12, 27–49.

Bergh, A., Nilsson, T., 2010, Do economic liberalization and globalization increase income inequality?, European Journal of Political Economy. 26, 488-505.

Bleaney, M., Nishiyama, A., 2003, Convergence in income inequality: differences between advanced and developing countries, Economics Bulletin. 4, 1-10.

Campos, N. F., 2001, Will the future be better tomorrow? The growth prospects of transition economies revisited, Journal of Comparative Economics. 29, 663-676.

Chiquiar, D., 2008. Globalization, regional wage differentials and the Stolper–Samuelson Theorem: Evidence from Mexico, Journal of International Economics. 74, 70–93.

Clarke, G.R.G., Xu, L.C., Zou, H-f., 2006. Finance and income inequality: What do the data tell us? Southern Economic Journal. 72, 578-596.

Dobson, S., Ramlogan, C., 2008. Is there an openness kurnets curve? Evidence from Latin America. Notthingham Trent Universtiy Working Paper. 2008/10, 1-22.

Dollar, D., Kraay, A., 2004. Trade, growth and poverty, The Economic Journal. 114, F22–F49.

Eviews 7.0. User Guide , 2010, Quantitative Micro Software, LLC, California.

Eviews 8.0. User Guide , 2013, IHS Global Inc, California.

Ezcurra, R., Pascual, P., 2005, Is there convergence in income inequality levels among the European regions. Applied Economics Letters. 12, 763-767.

Feenstra, R.C., Hanson, G.H., 1995. Foreign direct investments and relative wages: Evidence from Mexico's Maquiladoras, NBER Working Paper. 5122, 1-34.

Gomes, F., 2007, Convergence in income inequality: The case of Brazilian municipalities, Economics Bulletin. 15, 1-9.

Jalil, A., 2012. Modeling income inequality and openness in the framework of Kuznets curve: New evidence from China, Economic Modelling. 29, 309–315.

Krugman P., 1991, Geography and Trade, MIT Press, Cambridge, Mass.

Kocenda, E., 2001, Macroeconomic convergence in transition economies, Journal of Comparative Economics. 29, 1-23.

Kutan, A.M., Taner M. Y., 2005, Real and nominal stochastic convergence: Are the new EU members ready to join the Euro zone?, Journal of Comparative Economics. 33, 387-400.

Lessmann, C., 2013, Foreign direct investment and regional inequality: A panel data analysis, China Economic Review. 24, 129-149.

Levine, R., Renelt, D., 1992, A sensitivity analysis of cross-country growth regressions, American Economic Review. 82, 942-963.

Lin, P-C., Huang, H-C., 2012, Convergence in income inequality? Evidence from panel unit root tests with structural breaks Empirical Economics. 43, 153-174.

Li, H., Zou, H-fu, 1998, Income inequality is not harmful for Growth: Theory and evidence, Review of Development Economics. 2, 318-34.

Liu, T., Liu, C., Liu, H., Wang, S., Rong, Q., Zhu, W., 2014, Did the key priority forestry programs affect income inequality in rural China? Land Use Policy. 38, 264-275.

Lucas R.E., 1988, On the mechanics of economic development, Journal of Monetary Economics. 22, 3-42.

Ma, Y., Dei, F., 2009. Product quality, wage inequality, and trade liberalization, Review of International Economics. 17, 244-260.

Mankiw N.G., Romer D., Weil D.N., 1992, A contribution to the empirics of economic growth, Quarterly Journal of Economics. 107, 407-437.

Matkowski, Z., Prochniak, M., 2004, Real economic convergence in the EU accession countries, International Journal of Applied Econometrics and Quantitative Studies, 1, 5-38.

Quah, D., 1993, Galton's fallacy and tests of the convergence hypothesis, Scandinavian Journal of Economics. 95, 427-443.

Ravallion, M., 2003, Inequality convergence, Economics Letters. 80, 351–356.

Romer P.M., 1986, Increasing returns and long run growth, Journal of Political Economy. 94, 1002-1037.

Sala-i-Martin, X., 1996, The classical approach to convergence of growth analysis. Economic Journal, 106, 1019-1036.

Solow R., 1956, A contribution to the theory of economic growth, Quarterly Journal of Economics. 70, 65-94.

Song, Y., 2013, Rising Chinese regional income inequality: The role of fiscal decentralization, China Economic Review. 27, 294-309.

Wei, S-J., Wu, Y., 2001, Globalization and inequality: Evidence from within China, NBER Working Paper. 8611, 1-34.

Winters A.L., McCulloch, N., McKay, A., 2004, Trade liberalization and poverty: The evidence so far. Journal of Economic Literature. 42, 72-115.

Wu, J-Y., Hsu, C-C., 2012. Foreign direct investment and income inequality: Does the relationship vary with absorptive capacity? Economic Modelling. 29, 2183-2189.

Yamauchi, F., 2011, Are schooling and roads complementary? Evidence from income dynamics in rural Indonesia, World Development. 39, 2232–2244.

Zhao, Y., 2001. Foreign direct investments and relative wages: The case of China, China Economic Review. 12, 40-57.