Income Inequality and FDI: Evidence with Turkish Data

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11 June 2014

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Abstract

This paper explores how foreign direct investment (FDI) and other determinants impact income inequality in Turkey in the short- and long-run. We apply the ARDL (Auto-Regressive Distributed Lag) modelling approach, which is suitable for small samples. The data for the study cover the years from 1970 to 2008. The empirical results indicate the existence of a cointegration relationship among the variables. The positive impact of the FDI growth rate on income inequality, worsening inequality, is shown to be significant in the short-run, though at the 10% significance level only and with a quantitatively small impact, and insignificant in the long-run. In other words, FDI increases income inequality initially somewhat but this effect disappears in the long run. The literacy rate clearly reduces inequality in the long run, but also in the short run. On the other hand, population growth worsens inequality in the long run, and the effect is quite large, though it has no statistically significant effect on inequality in the short run. Also, an increase in GDP growth reduces inequality especially in the short run (at a 5% level of significance) but also in the long run (though only at the 10% level).

1. Introduction

In recent decades, there have been numerous investigations into the relationship between income inequality and other variables. The literature indicates that income and wage inequality have been rising in many countries since the 1970s. There is supporting evidence, for both developed and developing countries, for an increase in inequality. In fact, Caselli (1999) states that "income and wage inequality has been rising in the United States, as well as in several other countries." Furthermore, Bernstein and Mishel (1997) and McDonald and Yao (2003) report that, starting in the early 1970s, income and wage inequality has increased quite sharply in the United States. There are some studies on developing countries examining the issue of income and wage inequality. Recent studies from developing countries indicate a rise in income and wage inequality as well. Miles and Rossi (2001) claim that "wage dispersion had increased significantly in developing countries, despite the openness to trade of these economies". In particular, Diwan and Walton (1997), Dev (2000), among others, state that income and wage inequality has increased in developing countries like Mexico and several other countries in Latin America.

The number of studies examining income inequality has increased in line with the rise of inequality. Many previous studies have investigated the relationship between income inequality and varied factors, which influence the overall distribution of income. Economists have been interested in how other factors than foreign direct investment (FDI) affect income inequality. For instance, Rapanos (2004) examines the effects of a change in the minimum wage on the income distribution and employment in a developing economy. Saunders (2005) investigates the recent trends in wage income inequality in Australia. The author reports that full-time earnings inequality has increased since the mid-1970s for both men and women. His findings show that further labor market deregulation created more inequality of wage outcomes. Furthermore, Kijima (2006) analyses how and why inequality has accelerated in India. The author argues that wage inequality in urban India started increasing before 1991.

Related to the issue of FDI and income inequality is the relationship between trade liberalization and inequality that has received considerable attention in recent years. In this context, Wood (1997) examines the relationship between openness and wage inequality in developing countries. He states that the experience of East Asia indicates that "greater openness to trade tends to narrow the wage gap between skilled and unskilled workers in

developing countries". However, in Latin America, increased openness affected the wage differential upwards. Additionally, Munshi (2008) provides panel data evidence on trade liberalization and wage inequality in Bangladesh. His results indicate some weak evidence that openness contributes to a reduction in wage inequality between skilled and unskilled workers. Cornia (2005) analyses the relationship between within-country income inequality and policies of domestic liberalization and external globalization. The author argues that inequality often rose with the introduction of such reforms. Gourdon (2007) presents new results on the sources of wage inequalities in manufacturing taking into account South-South trade. The author observes increasing wage inequality is more due to the South-South trade liberalization than to the classical trade liberalization with northern countries.

Anderson *et al*. (2006) investigate the relationship between globalization, co-operation costs, and wage inequalities. The authors report that globalization "tends to narrow the gap between developed and developing countries in the wages of less-skilled workers, but to widen the wage gap within developed countries between highly-skilled and less-skilled workers." Miles and Rossi (2001) investigate the effects of market forces or government intervention on wage inequality. They find that "in Uruguay most of the increase in wage dispersion could be explained by a significant increase in public wages and a decrease of the minimum wage". Moreover, Cortez (2001) evaluates the impact of the educational expansion and changes in labor market institutions on wage inequality among Mexican workers using a simulation technique. The author concludes that "while increases in the relative rate of return of higher education would have induced an increase in wage inequality, changes in the composition of the educational distribution would have led to a stronger decline in wage inequality".

There is a growing interest in recent years in examining the relationship between FDI and income inequality. Basu and Guariglia (2007), Bircan (2007), Jensen and Rosas (2007), Sun (2007), Choi (2006), Stringer (2006), Tang and Selvanathan (2005), and teVelde (2003), among others, examine how FDI affects income inequality. In this paper, we attempt to investigate the relationship between FDI and income inequality in Turkey. Our major motivation for this paper is that there has been a significant increase in FDI inflows to Turkey during the past decade. In fact, FDI inflows to Turkey reached about 10 billion dollars in 2005, compared to only 2.8 billion dollars in 2004. This figure increased to around 20 billion dollars in 2006 and about 22 billion dollars in 2007. However, between the years 1980 and

2000, the total amount over the entire period was only around 15 billion dollars. Another motivation is the rising income inequality in Turkey. In fact, as Bircan (2007) states, income and wage inequality is high in Turkey.

The main purpose of this paper is to analysis the relationship between income inequality and FDI in Turkey. We investigate how FDI inflows affect domestic income inequality by using the Auto-Regressive Distributed Lag (ARDL) modeling approach to cointegration. The ARDL method can be applied regardless of whether variables have a unit-root or are covariance stationary. Furthermore, the methods corrects for endogeneity and serial correlation. The remaining sections are organized as follows: Section 2 presents the literature review and an overview of previous studies. Section 3 explains the econometric methodology and data used for examining the relationship between FDI inflows and domestic income inequality in Turkey. Section 4 analyses the relationship for the long-run and the short-run by using ARDL modeling and presents empirical results. Section 5 evaluates our findings.

2. Literature Review

As mentioned above, there is a growing interest in examining the relationship between FDI and income inequality lately. Choi (2006) states that, with the recent increase in FDI, concerns about the effects of FDI on income inequality have heightened. However, there are very few studies that examine this issue in Turkey. In this section, we present the results of recent studies which analyze the relationship between income inequality and FDI. We should mention that theories regarding the impact of FDI show that FDI may increase or decrease income inequality. The issue cannot be settled theoretically. However, empirical findings on the effects of FDI on income distributions are mixed as well.

Choi (2006) analyses the relationship between FDI and income inequality within countries using pooled Gini coefficients for 119 countries from 1993 to 2002. The author attempts to determine whether FDI affects domestic income inequality. Choi (2006) finds that income inequality increases as FDI stocks (as a percentage of GDP) increase. Furthermore, Figini and Görg (2011) analyze the relationship between FDI and wage inequality by using a panel of more than 100 countries for the period 1980 to 2002. The authors argue that the effects of

FDI differ according to the level of development. The results are that wage inequality decreases with FDI stocks in developed countries, however for developing countries, "wage inequality increases with FDI stocks but this effect diminishes with further increases in FDI." Moreover, Stringer (2006) examines the effects of FDI on income inequality in developing countries. In the paper, the author uses industry level data in an attempt to further the understanding of the causal mechanisms behind the relationship of FDI and income inequality.

Herzer and Nunnenkamp (2013) examine the effects of inward and outward FDI on income inequality in Europe by using panel cointegration techniques and unbalanced panel regressions. The results show that, on average, both inward and outward FDI have a negative long-run effect on income inequality. Furthermore, Bhandari (2007) empirically tests the link between FDI and income inequality for transitional countries in Eastern Europe and Central Asia for the period 1990 to 2002. The statistical evidence of the paper suggests that FDI inward stocks exacerbated wage income inequality, while reducing capital income inequality[1].

Furthermore, teVelde (2003) analyses FDI and income inequality for Latin America experiences and argues that income inequality is persistent and relatively high in almost all Latin American countries. The author reviews results with different data sources and states that "all findings support the conclusion that in most countries the relative position of skilled workers has improved over much of the late 1980s and early 1990s". Moreover, teVelde (2003) mentions that not all types of workers necessarily gain from FDI to the same extent. The author argues that a review of micro and macro evidence shows that, at a minimum, FDI is likely to perpetuate inequalities. In another study on Latin America, Herzer, Hühne and Nunnenkamp (2014) investigate the long-run impact of FDI on income inequality in five Latin American host countries, namely Bolivia, Chile, Colombia, Mexico and Uruguay by applying country-specific and panel cointegration techniques. According to their results,

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[1]Hanousek *et al*. (2011) survey the literature on direct and spill-over effects of FDI and conduct a meta-analysis for transition economies going from a command to a market system in central and eastern Europe, the Balkans and the Commonwealth of Independent States. They find that the weakening of FDI effects over time found in several studies is due to a publication bias in these studies. See also Herzer *et al.* (2008) on FDI and economic growth in general.

except for Uruguay, FDI contributes to widening income gaps in all individual sample countries.

On the other hand, Jensen and Rosas (2007) examine the relationship between investments of multinational corporations (foreign direct investment) and income inequality in Mexico. They use an instrumental variables approach and find that increased FDI inflows are associated with a decrease in income inequality within Mexico's thirty-two states. Furthermore, Tang and Selvanathan (2005) examine the relationship between FDI inflows and regional income inequality using data for the period 1978 to 2002 at national, rural and urban levels. They find that FDI inflows are one of the main factors that have led to increasing regional income inequality at the national level, as well as in rural and urban regions of China.

In Bircan's paper (2007), where the author investigates the effects of FDI on the manufacturing sector in terms of wages and productivity, models are estimated in order to demonstrate the impact of plant-level foreign equity participation on wages. The results indicate that "foreign plants pay on average higher wages to their workers, and both production and non-production workers benefit from foreign ownership," which might be interpreted as more FDI participation increasing the wage inequality within the plants, as well as across them.

Tsai (1995) investigates the relationship between FDI and income inequality by comparing models with and without geographical dummies. The study shows that the statistically significant correlation between FDI and income inequality is widely prevalent in earlier studies. Moreover, Vijaya and Kaltani (2007) examine the impact of FDI on manufacturing wages by using a cross-country analysis. According to the results, "the FDI-flows have a negative impact on overall wages in the manufacturing sector and this impact is stronger for female wages."

One group of the studies examines the relationship between FDI, growth and income inequality. For instance, Sun (2007) investigates the relationship between FDI, economic growth, and income inequality in a pooled time-series cross-section statistical model with 68 countries from 1970 to 2000. The author finds that there is no effect of FDI stocks on income inequality while the effect of FDI inflows on income inequality is non-linear. Additionally,

Basu and Guariglia (2007) examine the interactions between FDI, inequality, and growth, both from an empirical and a theoretical point of view. They use a panel of 119 developing countries and observe that FDI promotes both inequality and growth. Furthermore, Kuştepeli (2006) explores the relationship between income inequality and economic growth in the context of EU enlargement. The paper evaluates how the latest enlargement of the EU affects the relationship between income inequality and growth, for both original EU member countries and for countries in the enlarged EU region. The results show that there is no evidence of a significant effect for any of the groups of countries in the paper. Moreover, Giovannetti and Ricchiuti (2005) analyze the effects of new patterns of FDI on growth and inequality, with particular attention to the Mediterranean Partner Countries. A recent study that includes data for the recent global financial crisis by Asteriou *et al.* (2014) looks at a panel of 27 EU countries, including sub-groups, and finds that the highest contribution to income inequality comes from FDI. Also, the financial crisis significantly increased inequality in the EU-periphery and the new member states.

3. Empirical Modelling and Econometric

3.1 Theoretical Aspects of Modelling Income Inequality

The conventional Heckscher-Ohlin model of international trade considers two countries that are identical, except for their resource endowments. If emerging countries are deemed relatively abundant in unskilled labor, and the opposite is true for developed countries, then FDI should be concentrated in activities that use less-skilled labor intensively in emerging economies, according to standard trade theory[2]. Then, FDI should lead to an increase in the demand for low-skilled labor and drive up wages of the low-skilled workers relative to the wages of the skilled workers in the emerging economy. Therefore, income inequality will decline in the emerging economy as FDI increases. However, when the restrictive assumptions of the Heckscher-Ohlin type model are relaxed, the effects of FDI on the income distribution can be negative, leading to more inequality. For example, Feenstra and Hanson (1996, 1997) present a model, along with empirical evidence to support it, where FDI increases the relative wage of the skilled workers in the emerging economy (Mexico) as well

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[2]The Stolper-Samuelson theorem predicts that trade (and FDI) would take advantage of the relatively abundant factor of production, which is low-skilled labor in the emerging economy (see, for example, Lee and Vivarelli, 2006).

as in the developed economy (United States). The activities related to FDI in their model employ relatively large amounts of unskilled labor from the perspective of the developed country. However, from the perspective of the emerging country, the labor used in FDI activities in relatively large amounts is skilled labor and not unskilled labor, comparing skilled and unskilled labor within the Mexican labor market.

Another example of relaxing the assumptions of the standard Heckscher-Ohlin type model is to allow for production functions (technologies) that differ across countries (e.g., Grossman and Helpman, 1991). FDI can have adverse effects on income inequality in such a model. Further, technological change may be skill-biased (Wang and Bloomstrom, 1992) and increase the relative wage of skilled workers. Also, FDI can be seen as a vehicle for bringing new technologies into a country, with spill-over effects when imitation by local firms occurs (Piva, 2003). FDI can also lead to intra- and inter-industry technology upgrading (Kinoshita, 2000). If these new technologies require relatively more skilled than unskilled labor, relative wages of skilled labor increase along with FDI (teVelde, 2003). Figini and Görg (2011) also consider FDI as a vehicle to introduce new technology into a country, such as FDI carried out by multinational firms.[3] They use the endogenous growth model of Aghion and Howitt (1998). A new technological innovation in that model leads to increases in wage inequality at the early stage because firms use skilled labor to implemnt the new technology. However, at later stages less skilled labor is used when the new technology has been implemented and more wage equality is the result[4].

Various other theoretical models and explantions of the relatiosnhip between FDI and income inequality have been proposed in the literature. For example, FDI can cause crowding-out of domestic production (Aitken and Harrison, 1999) and investment (Berg and Taylor, 2001). Moreover, the employment effects of FDI may be country- and sector-specific (Lee and Vivarelli, 2004). Here, FDI affects the income distribution via realtive wages. Overall, on a theoretical level the direct and indirect effects of FDI could improve or worsen income

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[3] See also Saglam and Sayek (2011) on the role of productivity spill-overs and imperfect labor markets for domestic wages when multinational firms are active.

[4] A related literature, surveyed by Clark and Higfill (2011) and Ostry *et al*. (2014), debates at what point inequality becomes harmful to economic growth and swamps any positive effects of inequality on growth that stem from providing rewards for effort and innovation.

inequality. The issue cannot be settled on a theoretical level. An answer has to come form empirical investigation.

3.2 The Empirical Model

Income inequality is relatively high in Turkey, depending on the countries Turkey is compared to, of course[5]. The links between income inequality and FDI are multifaceted; however, we attempt to examine the relationship in Turkey. In the econometric analysis, we do not only use FDI as a determinant of income inequality. A linear model will be used to test the hypothesis of causality and study the long-run relationship. We explore the effects of the following variables on income inequality: FDI, the population growth rate (*POPGR)*, the inflation rate (*INF*), the GDP growth rate (*GDPGR*), and the literacy rate (*LR*).

$$GINI = f(FDIGFC, \ INF, \ LR, \ POPGR, \ GDPGR) \qquad (1)$$

Inequality is measured by the Gini index (*GINI*) and *FDIGFC* is the inward annual FDI flow into Turkey, expressed as a percentage of gross fixed capital formation. *INF* is the annual inflation rate, based on the GDP-deflator and *GDPGR* is the annual growth rate of GDP. *LR* is in annual percentage change in the adult literacy rate and *POPGR* is the annual population growth rate. In this study, we take into account only the macroeconomic factors that affect the Gini coefficient, with a particular emphasis on FDI.

Herzer and Nunnenkamp (2013) state that Gini coefficients cannot strictly be a pure unit-root process because Gini indices are bounded from below and above and a true unit-root process would cross any bound with probability 1. However, in the relevant range in small samples, unit-root behavior may approximate the unknown true data generating process much better than a near-unit-root process with very high persistence. Gini coefficients are likely affected by permanent shocks to factors such as tastes, time preferences and government policies, which lead to unit-root behavior. In a unit-root process, shocks have permanent effects, in contrast to, say, a mean-reverting stationary process where they have only temporary effects.

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[5]The Gini index of inequality (from the SWIID data base) in 2005 was 43.92 in Turkey. It reached a value of 67.76 in that year for South Africa, at the higher end. The index in 2005 is 45.04 for Russia, 37.05 for the United States, 34.64 for the United Kingdom, 28.11 for Germany, 23.70 for Sweden and 25.34 for Norway, to give just a few examples for comparison purposes.

Guest and Swift (2008) found that the Gini coefficients are stationary in first differences and are therefore I(1) for all countries in their study. Similarly, Chintrakarn *et al.* (2012) state that Gini coefficients are integrated and cointegrated with other variables (determinants) for the United States.

3.3 The Econometric Methodology

First, we focus on examining the time-series properties of our data before estimating the model of inequality in equation (1). Previous studies have found it difficult or impossible to find data (especially quarterly long-term time series) for Turkey due to the late declaration of the Gini coefficient and FDI data problems and our study is no exception. Therefore, we use annual data that are available from 1970 to 2008 for all the variables included in the subsequent estimations.

When considering stationarity of the macroeconomic time series data from 1970 to 2008, we analyze the data for a unit root in the levels and also for a unit root in the first differences, i.e., we test for I(1) and I(2). Next, we examine the long-run relationship of FDI with its determinants. The residual-based co-integration tests are sensitive to the specification of the test regression and the tests can lead to conflicting results, especially when there are more than two I(1)variables in the analysis. The model of income inequality is estimated within the context of recent developments in econometric methodologies, particularly with respect to cointegration analysis and error correction models that allow estimation of both the short-run and long-run dynamics.[6] In this regard we use two different methodologies, the Auto-Regressive Distributed Lag (ARDL) approach to cointegration (Pesaran and Shin, 1998) and the fully modified ordinary least squares (FM-OLS) method of Phillips and Hansen (1990), in order to calculate the long-run coefficients. Both methods correct for endogeneity and serial correlation in cointegrating regressions, thereby providing unbiased estimates of the cointegrating coefficients. These methodologies have proven to produce reliable estimates in small samples and provide a cross-check for the robustness of the results[7]. The advantage of

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[6]The concepts of the long-run and the short-run do not determine a specific period of time such as 10 years or 5 months.

[7]The time period that we look at is not very long but these methodologies are the best available in this case. See Pesaran and Shin (1998) for more information. In particular, the ARDL method uses simulations for proper inference in small samples.

the ARDL method is that it can be applied regardless of whether variables are I(0) or I(1), whereas FM-OLS relies on variables that are I(1). Moreover, it is generally the case that the time span of the period considered for the empirical analysis is of crucial importance when studying long-run relationships, such as cointegration. The frequency of observation is of lesser importance (Haug, 2002). In other words, moving from annual to quarterly data would certainly increase the sample size but it would likely not help much in terms of getting better estimates for the long-run coefficients that we are most interested in.

In what follows we briefly explain these two methodologies. Assume that the long-run formulation of the cointegration regression is

$$y_t = \mu + \delta\, t + \theta' x_t + v_t \quad (2),$$

where $\Delta x_t = e_t$ and $\xi_t = (v_t, e_t')'$ follows a general linear stationary process. In this case the ordinary least squares (OLS) estimators of δ and θ are consistent, but in general the asymptotic distribution of the OLS estimator of θ involves the unit-root distribution as well as a second-order bias in the presence of the contemporaneous correlation that may exist between v_t and e_t. Therefore, the finite sample performance of the OLS estimator is poor, and in addition nuisance parameter dependences make inference on θ using the usual t-test in the OLS regression of (2) invalid. To overcome these problems, Phillips and Hansen have suggested the fully modified OLS (FM-OLS) estimation procedure that asymptotically takes account of these correlations in a semi-parametric manner. FM-OLS assumes that v_t and e_t in (2) follow a general correlated linear-stationary process:

$$v_t = A_1(L)u_t \quad \text{and} \quad e_t = A_2(L)\varepsilon_t \quad (3)$$

where $\xi_t = (v_t, e_t')'$ are serially uncorrelated random variables with zero means and a constant variance. Assuming $A_1(L)$ and $A_2(L)$ are invertible, FM-OLS takes into account the presence of a constant term and possible correlation between the error term and the differences of the regressors.

The use of the ARDL estimation procedure is directly comparable to the semi-parametric FM-OLS approach to the estimation of cointegrating relationships. Pesaran and Shin proved that OLS estimators of the ARDL model lead to consistent short-run parameter estimates and to super-consistent long-run parameter estimates. Therefore, standard asymptotic normal

theory applies for carrying out statistical inference with the OLS parameter estimates, i.e., usual F- and t-tests can be used. However, due to asymptotic nature of the model, it is necessary to explore how well the ARDL methodology performs in typical small samples. Pesaran and Shin have carried out a Monte Carlo study with samples of size T= (50, 100, 250). They also compare the ARDL approach to the FM-OLS method of Phillips and Hansen, which is the closest competitor for inference with I(1) variables. They compare bias of the two estimators and size- and power properties of associated t-statistics in the Monte Carlo simulations. They find that the bias is generally smaller for the ARDL estimates than for the FM-OLS estimates. Similarly, empirical test sizes are much closer to their nominal values for the ARDL method as compared to the FM-OLS method. In addition, the ARDL method leads to test with higher power that the FM-OLS method, as far as power comparisons are feasible. However, Pesaran and Shin point out that their Monte Carlo comparison of the two methods is not "comprehensive" because the data generating process used by them favors the ARDL method. For this reason, we apply both methods, ARDL and FM-OLS, in our empirical analysis. If the results of the two methods are close to each other, we can be quite confident that the results are fairly reliable and robust.

We consider the following general ARDL (p, m) model:

$$\Delta y_i = \beta_0 + \pi_{yy} y_{t-1} + \pi_{yx} x_{t-1} + \sum_{i=1}^{p} \vartheta_i \Delta y_{t-i} + \sum_{j=0}^{m-1} \phi \Delta x_{t-j} + \theta w_t + \mu_t \quad (4)$$

Here, π_{yy} and π_{yx} are long-run multipliers. β_0 is the drift and w_t is a vector of exogenous components, e.g., dummy variables. Lagged values of Δy_t and current and lagged values of Δx_t are used to model the short-run dynamic system. As a starting point for the ARDL approach, we estimate equation (4) in order to examine first if there is a long-run relationship among the variables by carrying out an F-test. We denote the test, normalized on inequality, by F(*GINI| FDIGFC, INF, LR, POPGR, GDPGR*). In cases where independent variables are I(d), two asymptotic bounds for critical values provide a test for cointegration. No matter what order of integration "d" the time series are, the null hypothesis of no long-run relationship can be rejected if the F-statistic exceeds the upper critical value. Conversely, it cannot be rejected when the test statistic is below the lower critical value. In the second step, when there is a long-run relationship between variables, there is an error correction representation. In the next step of the analysis, the error correction model (ECM) is estimated. The error correction

model estimation result shows the speed of the adjustment back to the long-run equilibrium after a short-run shock.

4. Empirical Inference

This section presents empirical results on the relationship of income inequality (the Gini coefficient) and *FDIGFC, INF, LR, POPGR* and *GDPGR* in Turkey. As our focus is on Turkey in particular, for which data availability is somewhat limited, we undertake a time series analysis for annual data for a period of 38 years, from 1970 to 2008. The logarithmic form is not appropriate for modeling because of percentage growth rate values that we use (some negative values occur).

The Gini index (*GINI*) of inequality in equalized household disposable income, post-tax and post-transfers, was obtained from SWIIDv3.0 (Standardized World Income Inequality Database, version 3.0) and TURKSTAT. The FDI flow into Turkey was retrieved from UNCTAD. The GDP-deflator is from the World Bank, as is *GDPGR* and *POPGR*. The adult literacy rate was retrieved from the World Bank and TURKSTAT.

The ARDL approach has the advantage that it does not require pre-testing of the regressors for the presence of unit roots, a problem that afflicts other approaches to estimation of long-run relations, such as the FM-OLS approach of Phillips and Hansen (Pesaran, 1997). This can be particularly an issue when the unit-root test results are mixed, as they will turn out to be in our case. In any event, we study first the integrating order of all the variables by applying standard unit-root tests. Unit-root tests allow us to classify each series as being stationary or having one or more unit roots. The Augmented Dickey-Fuller (ADF) and Phillips-Peron (PP) tests are tests for the null hypothesis of a unit root against the alternative hypothesis of a stationary process around a constant mean or deterministic time trend. The Kwiatkowski–Phillips–Schmidt–Shin (KPSS) test considers instead the null hypothesis of stationarity versus the alternative hypothesis of a unit root. The ADF and PP test results in Table A1 in the Appendix show that all variables are non-stationary in levels and stationary in first differences (i.e., have a unit root), except for possibly GDP growth (*GDPGR*). Both tests indicate that GDP growth is likely stationary in levels, i.e., is I(0). The KPSS results corroborate these findings for all variables with the following exceptions. The KPSS test

indicates that *GINI , INF* and *LR* are possibly stationary in levels for a 5% significance level, which is contradicted by the ADF and PP tests. However, the KPSS test results are a borderline case and at the 10% level of significance the null of stationarity is rejected in all three cases in favor of a unit root. The only other cases of test conflict are for the first differences of *FDIGFC* and *POPGR*, where the ADF test does not rejected the null hypothesis, indicating I(2). The ADF test may lack power. Also, it is possible that the presence of structural breaks leads to a spurious finding of either I(0), I(1), or I(2) behavior, depending on where in the sample the break occurs (Leybourne *et al*., 1998). For this reason, we employ next a unit-root test that considers up to two breaks, both under the null hypothesis of a unit root and under the alternative hypothesis of stationarity around a trend. Due to events in the Turkish economy, the potential presence of structural breaks is a main concern.

The standard unit-root tests that we used cannot identify structural breaks. Lee and Strazicich (2003) propose a unit-root test that is valid when there are possibly two structural breaks present in the sample. It is a two-break minimum Lagrange Multiplier (LM) unit-root test in which the alternative hypothesis definitely implies the series is trend stationary (Glyyn *et al*., 2007). A unique feature of this test is that it consider up to possibly two breaks under the null hypothesis of a unit root and under the alternative hypothesis of a trend-stationary process. In other words, a unit-root process with up to two breaks is tested against a trend-stationary process with up to two breaks. The null and alternative hypotheses are treated symmetrically in regards to breaks. This is an advantage over other break tests for unit roots that allow only a break under the alternative hypothesis. Lee and Strazicich show that the two-break LM unit-root test statistic, which is estimated according to the LM principle, will not spuriously reject the null hypothesis of a unit root.

Table A1 in the Appendix reports results for the unit-root t-statistics in the presence of breaks, along with the dates of breaks. We consider two models, one with two breaks in the constant term only, the other with two breaks each in the constant and trend. In the model with a trend, we report a significant break if at least one break is significant, either in the constant term or in the trend term. Once we allow for two breaks, the Gini index is still I(1) but *FDIGFC* and inflation seem to be I(0). The literacy rate, *LR*, is either I(0) or I(1), depending on whether the break is in the constant and trend or only in the constant,

respectively. The results for population growth and GDP growth remain unchanged when we allow for breaks. These mixed results illustrate the need for a method such as ARDL where it is unnecessary to pre-test for the order of integration.

We would like to emphasize that in regards to breaks, we are interested whether the linear ARDL function in equation (4) shows evidence of structural change, i.e., whether the relationship is stable over time, regardless of how the individual time series behave. It is possible that the co-movement of variables compensates for breaks in individual series when one models an error-correction process with a long-run equilibrium (the cointegrating relationship). In order to assess the structural stability of the ARDL model, we will examine the residuals from the ARDL regression with the CUSUM and CUSUMsq tests.

We start the ARDL analysis with testing for the existence of a long-run relationship. The ARDL approach to cointegration involves the comparison of the F-statistics against the appropriate critical values, as explained in Pesaran and Pesaran (1997) and Pesaran et al. (2001)[8]. They report two sets of critical values that provide critical value bounds for all classifications of the regressors into purely I(1), purely I(0) or mutually cointegrated. However, these critical values are generated for sample sizes of 500 and 1000 observations. Narayan (2005) argues that existing critical values cannot be used for small sample sizes because they are based on large sample simulations. He calculates two types of critical values, for a chosen significance level, with and without a time trend for small sample of between 30 to 80 observations. One set assumes that all variables are I(0) and the other set assumes they are all I(1). If the computed F-statistics is higher than the upper bound of the critical value then the null hypothesis of no cointegration is rejected. The F-statistic with income inequality as the dependent variable is: F(*GINI|FDIGFC, INF, LR, POPGR, GDPGR*) = 61.88. Based on 2000 replications, we calculated the upper bound critical value with stochastic simulations in Microfit (Pesaran and Pesaran, 1997) as 25.971 at the 5% level (and 28.692 at the 2.5% level). This leads us to conclude that the null hypothesis of no cointegration is rejected. For each model a maximum of one lag was used in the estimated ARDL model, chosen by the Schwarz Bayesian Criterion (SBC), which chooses the lag length consistently, for all possible combinations of values of *p* and *m* in equation (4).

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[8]The ambiguities in the order of integration of the variables lend support to the use of the ARDL bounds approach rather than one of the alternative co-integration tests.

Appropriate standard errors for the estimated ARDL regression coefficients in small samples are constructed with Bewley's approach, which produces the same results as the delta-method as they are numerically identical (Pesaran and Shin, 1998). In small samples, tests based on such standard errors perform much better than tests using standard errors based on asymptotic distributions (Pesaran and Shin, 1998). Results for the long-run model estimated by using ARDL and FM-OLS are presented in Table 1. The two methods provide similar results and have the expected signs, confirming the robustness of the long-run results.[9]

In this paper, long-term results are as expected on the basis of economic theory. Increased flows of FDI could have a positive effect on the distribution of income in developing countries. Our estimated coefficients of the long-run relationship for both methods show that *FDIGFC* has a positive effect and has therefore the sign implied by some of the theories but it is statistically insignificant. This indicates that FDI growth has no significant effects on inequality in the long-run. This result is to be expected for the Turkish economy because the amount of FDI inflows is not enough to affect inequality in the long-term by much. FDI is

Table 1. Estimated Long-Run Coefficients

Dependent variable is *GINI*					
		Regressor	Coefficient	Standard Error	*t*-statistics (*p*-values)
A.	ARDL	α	1.258	0.66	1.89 (0.043)**
	estimates	*FDIGFC*	0.107E-3	0.18E-3	0.61 (0.278)
		INF	0.217E-3	0.52E-3	0.41 (0.344)
		LR	-0.019	0.009	-2.06 (0.033)**
		POPGR	0.126	0.05	2.43 (0.017)**
		GDPGR	-0.003	0.002	-1.74 (0.056)*
B.	FM-OLS	α	1.400	0.77	1.83 (0.048)**
	estimates	*FDIGFC*	0.144	0.56	0.26 (0.401)
		INF	0.189	0.68	0.28 (0.396)
		LR	-0.025	0.007	-3.73 (0.002)***
		POPGR	0.193	0.03	5.64 (0.001)***
		GDPGR	-0.005	0.003	-1.54 (0.077)*

Notes: 1. ***, **, * denote the 1%, 5% and 10% significance levels.
2. FM-OLS was estimated with Parzen weights and a truncation lag equal to 1.
3. An ARDL (1,1) was selected with the Schwarz Bayesian Criterion (see Table A2 in the Appendix for other diagnostic statistics for the ARDL model supporting the results of Table 1).

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[9] No multicollinearity was found in the model. Multicollinearity would indicate indirect or imprecise relationships.

Generally going into telecommunication and service sectors, like banking and finance. Employment positions in these sectors are being mostly taken by high-skilled labor. This increases aggregate income inequality and also magnifies the differences between rural and urban earnings (Shahbaz and Aamir, 2008). However, our results show that such sectoral effects have no significant influence on the aggregate income distribution, or might be offset by other influence that move in the opposite direction. The inflation rate has a positive coefficient estimate and the literacy rate a negative one. The effect is not statistically significant for inflation in the long run. On the other hand, the literacy rate, *LR,* has a statistically significant influence at the 5% level on inequality in the long run and, as one would expect, decreases inequality. A 1% increase in the literacy rate lowers the Gini coefficient by 1.9 points, using the usual Gini scale from 0 to 100[10]. Therefore, increasing literacy rates is an effective way to decrease income inequality in Turkey in the long run. From the estimates, the population growth rate has a positive and statistically significant effect at the 5% significance level. A 1% increase in the population growth rate increases the Gini coefficient by 12.6 points, which is the largest estimated effect in absolute terms. The GDP growth rate, *GDPGR,* has a negative and statistically significant effect at a 10% level, though it is a borderline case at the 5% level. A 1% increase in GDP growth reduces the Gini coefficient by 0.3 points per year in the long term. This result implies that the poor benefit from economic growth.

Table 2 reports the short-run coefficient estimates obtained from the ECM version of the ARDL model and the tests for normality of the residuals, for serial correlation, for heteroscedasticity, and for misspecification of the functional form of the ECM. The *ECM* coefficient demonstrates how quickly or slowly variables go back to equilibrium and it should have a statistically significant coefficient with a negative sign. The error correction term, ECM(-1), measures the speed of adjustment to restore equilibrium in the dynamic model. It appears with a negative sign and is statistically significant at the 5% level, ensuring that long-run equilibrium can be attained. The coefficient of *ECM*(-1) is equal to -0.38 for the short-run model and implies that deviations from the long-term inequality are corrected by about one third each year.

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[10]We rescaled the Gini coefficient in our regressions to be between 0 and 1 by dividing it by 100.

Table 2. Estimated Short-Run Coefficients

Dependent Variable is $\Delta GINI$			
Variables	Coefficient	Standard Error	t- statistics (p-value)
α	-0.477	0 .20	-2.42 (0.046)**
$\Delta FDIGFC$	0.731E-4	0 .36E-4	2.05 (0.080)*
ΔINF	0 .823E-4	0 .19E-3	0.44 (0.671)
ΔLR	-0.007	0 .004	-1.91 (0.098)*
$\Delta POPGR$	0.012	0 .02	0 .75 (0.480)
$\Delta GDPGR$	-0.001	0 .45E-3	-2.62 (0.034)**
$ECM(-1)$	-0.379	0.13	2.99 (0.020)**
Diagnostic Tests:			
Test Statistics	LM Version (p-value)		
Serial Correlation	χ^2_{Auto} (1)= 1.46 (0.227)		
Functional Form	χ^2_{RESET} (1)= 1.59 (0.207)		
Normality	χ^2_{Norm} (2)= 0.57 (0.753)		
Heteroscedasticity	χ^2_{White} (1)= 0.06 (0.801)		
$R^2 = 0.980$, adjusted $R^2 = 0.932$, $\sigma = 0.006$			

***, **, * denotes 1%, 5%, 10% significance levels and "*ECM*" is the error-correction term.

Results for short-run dynamics provide evidence that inequality increases with FDI in the short-run, at a 10% level of significance, though not at a 5% level. However, the quantitative effect is quite small. A 1% increase in the FDI to gross fixed capital formation ratio increases the Gini coefficient by only 0.007 points in the short run. Inflation is also positively associated with inequality but the effect is insignificant. An increase in the literacy rate is lowering inequality in the short run, at a significance level of 10%. *POPGR* does not affect

income inequality in the short-run in a statistically significant way. *GDPGR* decreases income inequality in short-run at a significance level of 5%.The effects of LR and GDP growth on the Gini coefficient are quite small, in absolute terms, in the short run.

Diagnostic tests for serial correlation, normality, heteroscedasticity and functional form all support the model as specified and results are shown in Table 2. These tests show that the error-correction model passes all diagnostic tests. The results show that the model passes the test for normality so that the error term is normally distributed. The functional form of the model is well specified and also there is no heteroscedasticity and no autocorrelation in the model. In addition, we used the cumulative sum (CUSUM) and the cumulative sum of squares (CUSUMsq) of the standardized recursive residuals of the ARDL regression for analyzing the stability of the model. The plots of both the CUSUM and the CUSUMsq in Figures 1 and 2 are within the 95% confidence bands and henceforth these statistics verify the stability of the ARDL model coefficients for income inequality for Turkey.
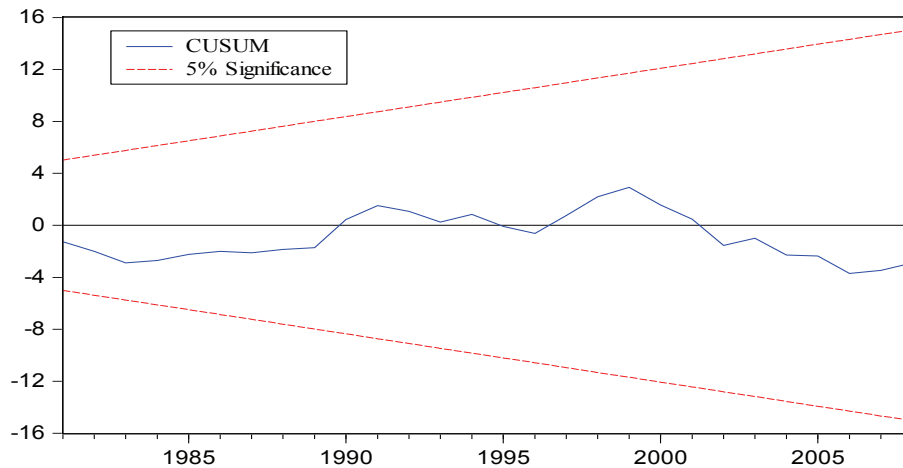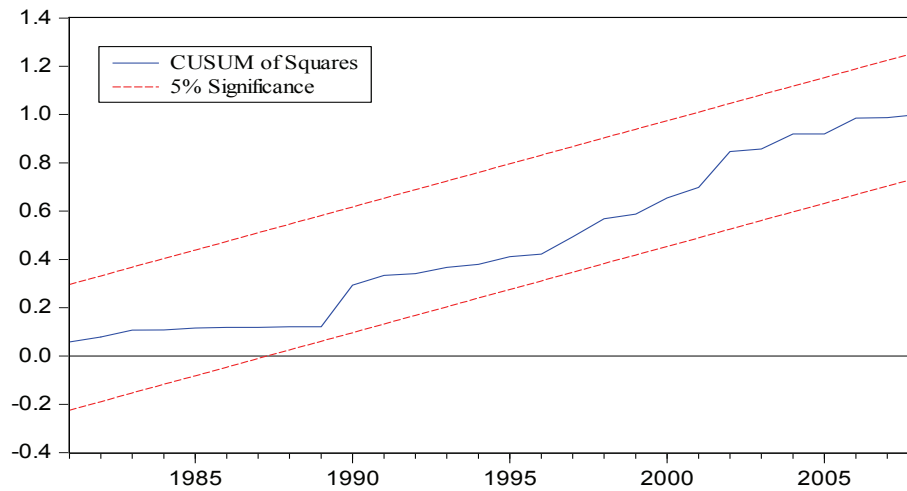
Figure 1. CUSUM

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Figure 2. CUSUM of Squares

5. Conclusion

In the literature, there are only a few empirical study of analyzing the relationship between FDI and income inequality but none exists for Turkey. This study investigates the importance of FDI in respect to inequality within the country. We apply ARDL and FM-OLS methods to investigate the long-run relationships among inequality and FDI in an error-correction version of the ARDL model and showed that the error-correction coefficient, which determines the speed of adjustment, had the expected negative sign and is significantly different from zero, despite the fact that we have available only a relatively small sample of observations. The results indicate that deviations from long-term inequality are corrected by approximately 38 percent in each of the following years. The model passes all of the diagnostic and stability tests. The error term is normally distributed. The CUSUM and CUSUMsq stability tests revealed that the estimated coefficients of the error correction model are stable.

Results show that increasing FDI inflows have caused income inequality in Turkey to increase in the short run but not in the long run. This is in line with the literature that suggests that FDI tends to worsen inequality (and poverty) initially. An increase in the literacy rate and GDP growth rate reduce inequality in the short and long run. The effect of the literacy rate is particularly statistically significant in the long run and the effect of GDP growth is so in the short run. On the other hand, population growth has a strongly adverse

effect on income inequality in the long run, though not in the short run. This study implies that policies that place GDP growth alone at the center of reducing income inequality will be insufficient in the long run. Improving literacy rates (education) is crucial for a sustainable solution to income inequality, in addition to sustained economic growth.

 A future study is planned to assess income inequality for urban and rural incomes in Turkey. In this regard, other factors of income inequality components will be included, such as environmental, political, governmental, and regional factors. For this purpose, we would like to design a questionnaire for measuring changes in rural and urban incomes.

References

Adams, S., 2008. Globalization and income inequality: Implications for intellectual property rights. Journal of Policy Modeling, 30**,** 725–735.

Aghion, P., Howitt, P. 1998. Endogenous growth theory. Cambridge, MA: MIT Press.

Aitken, B., Harrison, A., 1999. Do domestic firms benefit from direct foreign investment? Evidence from Venezuela. American Economic Review, 89, 605-18.

Anderson, E., Tang, P.J.G, Wood, A., 2006. Globalization, co-operation costs, and wage inequalities. Oxford Economic Papers, 58**,** 569-595.

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.

Basu, P., Guariglia, A., 2007. Foreign direct investment, inequality, and growth. Journal of Macroeconomics, 29, 824-839.

Bernstein, J., Michel, L., 1997. Has wage inequality stopped growing? Monthly Labor Review, 120, 3-16.

Bhandari, B., 2007. Effect of inward foreign direct investment on income inequality in transition countries. Journal of Economic Integration, 22, 888-928.

Bircan, C., 2007. The effects of foreign direct investment on wage inequality in developing countries: A case study of Turkey. Honors Thesis in Economics, Retrieved May 3, 2008, from http://www.williams.edu/Economics/Honors/2007/Bircan_thesis.doc

Caselli, F., 1999. Technological revolutions. American Economic Review, 89, 78-102.

Chintrakarn, P., Herzer, D., Nunnenkamp, P., 2012. FDI and income inequality: Evidence from a panel of US States. Economic Inquiry, 50, 788-801.

Choi, C., 2006. Does foreign direct investment affect domestic income inequality? Applied Economics Letters, 13, 811-814.

Clark, D. P., Highfill, J., 2011. FDI, technology spillovers, growth, and income inequality: A selective survey. Global Economy Journal, 11 (2), article 1, 1-42.

Cornia, G. A., 2005. Policy reform and income distribution. DESA Working Paper, 3, Retrieved May 12, 2008, from http://www.un.org/esa/desa/papers/2005/wp3_2005.pdf

Cortez, W. W., 2001. What is behind increasing wage inequality in Mexico? World Development, 29, 1905-1922.

Dev, M., 2000. Economic liberalization and employment in South Asia. ZEF-Discussion Papers on Development Policy, 29. Retrieved April 30, 2008, from http://www.zef.de/fileadmin/webfiles/downloads/zef_dp/zef_dp29-00.pdf

Diwan, I., Walton, M., 1997. How international exchange, technology, and institutions affect workers: An introduction. The World Bank Economic Review, 11, 1-15.

Feenstra, R.C., Hanson G.H., 1996. Foreign investment, outsourcing, and relative wages. In R.C. Feenstra, G.M. Grossman, D.A. Irwin (eds.), Political economy of trade policy: Essays in honor of Jagdish Bhagwati. MIT Press, Cambridge (Mass.), 89-127.

Feenstra, R. C., Hanson, G.H., 1997. Foreign direct Investment and relative wages: Evidence from Mexico's maquiladoras. Journal of International Economics, 42, 371-393.

Figini, P., Görg, H., 2011. Does foreign direct investment affect wage inequality? An empirical investigation. The World Economy, 34, 1455-1475.

Giovannetti, G., Ricchiuti, G., 2005. The effects of new patterns of FDI on growth and inequality: The case of southern Mediterranean countries. Economic Research Forum (ERF), Retrieved June 1, 2008, from http://www.erf.org.eg/cms.php?id=conferences_details& conference id=3

Guest, R., Swift, R., 2008. Fertility, income inequality, and labor productivity. Oxford Economic Papers, 60, 97–618.

Gourdon, J., 2007. Trade and wage inequality in developing countries: South-south trade matters. MPRA, Retrieved June 3, 2008, from http://mpra.ub.uni-muenchen.de/4177/1/ MPRA_paper_4177.pdf

Glyyn, J., Nelson, P., Reetu, V., 2007. Unit-root tests and structural breaks: A survey with applications. Journal of Quantitative Methods for Economics and Business Administration, 3, 63-79.

Grossman, G. M., Helpman, E., 1991. Innovation and growth in the global economy. MIT Press, Cambridge (Mass.).

Hanousek, J., Kočenda, E., Maurel, M. 2011. Direct and indirect effects of FDI in emerging European markets: A survey and meta-analysis. Economic Systems, 35, 301-322.

Haug, A. A., 2002. Temporal aggregation and the power of cointegration tests: A Monte Carlo study. Oxford Bulletin of Economics and Statistics, 64, 399-412.

Herzer, D., Klasen, S., Nowak-Lehman, F., 2008. In search of FDI-led growth in developing countries: The way forward. Economic Modelling, 25, 793-810.

Herzer, D., Nunnenkamp, P., 2013. Inward and outward FDI and income inequality: Evidence from Europe. Review of World Economics, 149, 395-422.

Herzer, D., Hühne, P., Nunnenkamp, P., 2014. FDI and income inequality - evidence from Latin American economies. Review of Development Economics, forthcoming.

Jensen, N. M., Rosas, G., 2007.Foreign direct investment and income inequality in Mexico, 1990-2000. International Organization, 61, 467-487.

Jones, C., 1995. Time series tests of endogenous growth models. Quarterly Journal of Economics, 110, 495–525.

Kijima, Y., 2006. Why did wage inequality increase? Evidence from urban India 1983-99. Journal of Development Economics, 81, 97-117.

Kinoshita, Y., 2000. R&D and technology spillovers via FDI: Innovation and absorptive capacity. William Davidson Institute Working Paper Series, 349, William Davidson Institute at the University of Michigan Business School, Ann Arbor.

Kuştepeli, Y., 2006. Income inequality, growth, and the enlargement of the European Union. Emerging Markets Finance and Trade, 42, 77-88.

Lee, J., Strazicich, M.C, 2003. Minimum Lagrange multiplier unit-root test with two structural breaks. Review of Economics and Statistics, 85, 1082-1089.

Lee, E., Vivarelli, M. 2006. The social impact of globalization in the developing countries. IZA DP No. 1925. Retrieved January 30, 2014, from http://ftp.iza.org/dp1925.pdf

Leybourne, S.J., Mills, T.C. and Newbold, P., 1998. Spurious rejections by Dickey-Fuller tests in the presence of a break under the null. Journal of Econometrics, 87, 191-203.

McDonald, C., Yao, J., 2003. Mauritius: unemployment and the role of institutions. IMF Working Paper, WP/03/211.

Mehic, E., Silajdzic, S., Lambert, P., Aranson, J.R., 1993. Inequality decomposition analysis and the Gini coefficient revisited. Economic Journal, 103, 1221-1227.

Miles, D., Rossi, M., 2001. Wage inequality in developing countries: Market forces or government intervention. Retrieved May 13, 2008, from http://decon.edu.uy/publica/2001/Doc1001.pdf

Munshi, F., 2008. Does openness reduce wage inequality in developing countries? A panel data analysis. Retrieved May 1, 2008, from http://gupea.ub.gu.se/dspace/bitstream /2077/2665/3/gunwpe0241corr.pdf

Narayan, P.K., 2005. The saving and investment nexus for China: evidence from cointegration tests. Applied Economics, 37, 1979-1990.

Ostry, J.D., Berg, A., Tsangarides, C.G., 2014. Redistribution, inequality, and growth. IMF Staff Discussion Note, SDN/14/02.

Pesaran, M.H., 1997. The role of economic theory in modeling the long run. Economic Journal, 107, 178-191.

Pesaran, M.H., Pesaran, B., 1997. Working with Microfit 4: Interactive econometric analysis. Oxford University Press, Oxford.

Pesaran, M.H., Shin, Y., 1998. An autoregressive distributed-lag modeling approach to cointegration analysis. Econometrics and economic theory in 20th century. The Ragnar Frisch Centennial Symposium, edited by Steinar Strom. Cambridge University Press, Cambridge, 371-413.

Pesaran, M.H., Shin, Y., Smith, R.J., 2001. Bounds testing approaches to the analysis of level relationships. Journal of Applied Econometrics, 16, 289-326.

Piva, M., 2003. The impact of technology transfer on employment and income distribution in developing countries: A survey of theoretical models and empirical studies. International Labour Office, Policy Integration Department, International Policy Group, Working Paper n.15, Geneva.

Rapanos, V.T., 2004. Minimum wage and income distribution in developing countries. EconWPA, 0404005, Retrieved May 20, 2008, from http://129.3.20.41/eps/it/papers/0404/0404005.pdf

Saglam, B. B., Sayek, S., 2011. MNEs and wages: The role of productivity spillovers and imperfect labor markets. Economic Modelling, 28, 2736-2741.

Saunders, P., 2005. Reviewing recent trends in wage income inequality in Australia. Social Policy Research Centre, University of South Wales. Retrieved May 18, 2008, fromhttp://www.econ.usyd.edu.au/download.php?id=4304

Stringer, J.M., 2006. Foreign direct investment and income inequality in developing countries: An exploration of the causal relationship using industry level FDI data. Paper presented at the annual meeting of the Midwest Political Science Association, Retrieved May 20, 2008, from http://www.allacademic.com/meta/p140292_index.html

Sun, F., 2007. Foreign direct investment, economic growth, and income inequality. Paper presented at the annual meeting of the Midwest Political Science Association, Retrieved May 22, 2008, from http://www.allacademic.com/meta/p198955_index.html

Turkish Statistical Institute [TURKSTAT], 2007. http://www. turkstat.gov.tr

United Nations Conference on Trade and Development [UNCTAD], 2007. http://www.unctad.org

Urban, I., Lambert P.J., 2008. Redistribution, horizontal inequity, and re-ranking: How to measure them properly. Public Finance Review, 36, 563-587.

Tang, S., Selvanathan, S., 2005. Foreign direct investment and regional income inequality in China. Retrieved May 16, 2008, from http://www.uib.es/congres/ecopub/ecineq/papers/245tang.doc

teVelde, D.W., 2003. Foreign direct investment and income inequality in Latin America experiences and policy implications. Overseas Development Institute, Retrieved May 16, 2008, from http://www.odi.org.uk/IEDG/Meetings/FDI_feb2003/fdi_la_dwtv.pdf

Tsai, P.L., 1995. Foreign direct investment and income inequality: Further evidence. World Development, 23, 469-483.

Vijaya, R.M., Kaltani, L., 2007. Foreign direct investment and wages: A bargaining power approach. Journal of World Systems Research, 13, 83-95.

Wagstaff, A., 2005. Decomposing changes in income inequality into vertical and horizontal redistribution and re-ranking, with applications to China and Vietnam. World Bank Policy Research Working Paper WPS3559. Washington, DC: World Bank. Retrieved May 16, 2008, from http://www-wds.worldbank.org/servlet/WDSContentServer/WDSP/IB/2005/04/13/ 000012009_20050413124029/Rendered/PDF/wps3559.pdf

Wood, A., 1997. Openness and wage inequality in developing countries: The Latin American challenge to East Asian conventional wisdom. World Bank Economic Review, 11, 33-57.

Appendix

Table A1. Unit-Root Test Results

Variables	ADF Stat	5% Critical Value	PP Statistics	5% Critical Value	KPSS Statistics	5% Critical Value	Lee-Strazicich(2003) break test [a]		
							t-statistic	First beak date	Second break date
GINI	-2.44	-3.54	-1.83	-3.52	0.141	0.146	-4.13 (-1.94)	1980* (1979*)	1997* (1996*)
ΔGINI	-4.60*	-2.94	-4.60*	-2.94	0.09	0.463	-6.40* (-5.81*)	1983* (1983)	1996* (1997)
FDIGFC	1.04	-3.57	-2.20	-3.53	0.15*	0.146	-7.43* (-4.57*)	1999* (1975)	2004 (2005*)
ΔFDIGFC	0.37	-2.96	-6.87*	-2.94	0.14	0.463	-10.94* (-7.20*)	2000* (1987)	2003* (2005*)
INF	-2.31	-3.53	-2.31	-3.53	0.137	0.146	-4.94* (-3.16*)	1996 (1998*)	1999 (2001)
ΔINF	-5.69*	-2.95	-6.85*	-2.94	0.14	0.463	-7.99* (-7.42*)	1997* (1979)	2000* (1982)
LR	-1.79	-3.53	-1.79	-3.53	0.13	0.146	-4.96* (-2.19)	1988* (1979*)	1992* (2001)
ΔLR	-6.29*	-2.94	-6.28*	-2.94	0.12	0.463	-13.40* -6.70*	1988* (1992)	1991* (1997)
POPGR	-2.12	-3.57	-2.24	-3.53	0.154*	0.146	-2.71 (-1.05)	1985* (1983)	1995* (1986)
ΔPOPGR	-1.98	-2.96	-2.96*	-2.94	0.18	0.463	-2.06 (-1.65)	1977* (1984*)	1983* (1993)
GDPGR	-6.18*	-2.94	-6.18*	-2.94	0.04	0.463	-6.95* (-6.59*)	1998* (1979*)	2005* (2004)
ΔGDPGR	-3.80*	-1.95	-11.7*	-1.95	0.08	0.463	-11.36* (-10.44*)	1979* (1998)	1988* (2004)

Notes: The lag augmentations for the ADF test are selected with Akaike's criterion. The PP test is based on a quadratic kernel and Andrews' automatic bandwidth selection. The critical values for these two tests are from EViews 8 and are based on response surface estimates. The KPSS test also uses a Bartlett kernel and Andrews' automatic bandwidth. The KPSS test has the null hypothesis of stationarity (no unit root), while the ADF and PP tests have the null hypothesis of non-stationarity (a unit root). The level tests include a deterministic time trend, except for the GDP growth rate. This time trend cancels out in the first differences. A "*"indicates rejection of the relevant null hypothesis at the 5% significance level.

[a]The first entry in the cell is the break model with two break in each, in the intercept and in the trend; the second entry in the cell, in parentheses, is the crash model with two breaks in the intercept only.

Table A2. ARDL estimates for ARDL (1, 1) selected with SBC

Dependent variable is *GINI*

Variables	Coefficient	Standard Error	t-statistics (*p*-value)
α	4.098	1.88	2.18 (0.030)
GINI(-1)	0.654	0.07	9.20 (0.000)
FDIGFC	0.556E-7	0.89E-7	0.63 (0.558)
FDIGFC(-1)	0.216E-6	0.85E-7	2.53 (0.053)
INF	0.504	0.62	0.85 (0.211)
LR	-0.071	0.03	-2.15 (0.084)
LR(-1)	-0.089	0.03	- 3.15 (0.025)
POPGR	1.484E-6	0.44E-6	3.36 (0.005)
POPGR(-1)	0.139E-5	0.46E-6	3.02 (0.029)
GDPGR	-0.004	0.02	-1.86 (0.050)
R^2	0.987	F-statistics	42.22 (0.000)
Res. Sum Sq.	2.82	Schwarz Bayesian Criterion (SBC)-22.30	
Adjusted R^2	0.964	Akaike Criterion 18.76	
		Durbin's h-statistic 1.96 (0.049)	