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Abstract:
This paper examines the relationship between foreign direct investment (FDI) and income inequality for a sample of ten European countries over the period 1980 to 2000. Using panel co-integration and causality techniques that are robust to omitted variables, slope heterogeneity, and endogenous regressors, we find that: (1) FDI has a positive short-run effect on income inequality in Europe, (2) the long-run effect of FDI on inequality, however, is negative on average, (3) long-run causality runs in both directions, suggesting that an increase in FDI reduces income inequality and, in turn, higher inequality leads to lower FDI inflows, and (4) there are large differences in the long-run effect of FDI on income inequality, with two countries (Ireland and Spain) exhibiting a positive relationship between FDI and income inequality.

Keywords: FDI, income inequality, panel co-integration, Europe

JEL classification: F21, D31, C23

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1. Introduction

Foreign direct investment (FDI) is often found to have positive effects on growth and productivity in the host country, but "what is generally neglected is the issue of equality" (Figini and Görg 2006: 1). Evidence on the distributional consequences of inward FDI is particularly scarce in the European context. This is surprising as public concerns about globalisation widening the income gap between rich and poor segments of the population figure prominently in current political and economic discourse in various European countries.

Previous studies focused on the labour market repercussions of outward FDI in the home countries of European companies engaged in international outsourcing and offshoring. Summarising this literature, Geishecker et al. (2008: 169) conclude that international outsourcing leads to a shift in the relative demand for labour, which has "implications for the wage bill as well as the employment prospects of high- versus low-skilled labour." However, "undesirable effects upon the labour market" (Taylor and Driffield 2005: 243) may also result from European countries being the target of FDI by companies based within or outside the region. According to the World Investment Report, Europe accounted for 46 per cent of the worldwide stock of inward FDI in 2009—just slightly less than its share of 53 per cent of outward FDI stocks (UNCTAD 2010).

We argue in Section 2 that the effects of inward FDI on income inequality in European host countries are theoretically ambiguous. Against this background, we explore the relationship between FDI stocks and income inequality for a sample of 10 European countries over the period 1980 to 2000. We employ panel co-integration and causality techniques that are robust to omitted variables, slope heterogeneity, and endogenous regressors (Section 3). Our results reported in Section 4 reveal that inward FDI reduces income inequality, though only in the longer run and not necessarily in less advanced host countries such as Ireland and Spain. At the same time, less unequal host countries attract more FDI.

2. Theoretical background and related empirical literature

The European host countries in our sample differ considerably in their overall level of economic development. The sample includes Malta, Spain, and Ireland, with an average per-capita income in the range of 3,400-5,200 US$ in 1980 (i.e., at the beginning of our period of observation) as well as Germany, the Netherlands, Norway, and Sweden, with a per-capita income in the range of 10,800 and 14,000 US$ (see also Section 3). Arguably, the distributional consequences of FDI from relatively rich source countries (within and outside the sample) in the poorer host countries in the sample may be similar to those predicted by North-South models, such as Feenstra and Hanson's (1997). Accordingly, the availability of relatively cheap labour in the poorer host countries may encourage firms based in the richer source countries to undertake cost-oriented, vertical FDI by offshoring labour-intensive parts of the production process. This type of FDI may increase the skill premium not only in the richer source country, but also in the poorer host country. This could happen when offshoring involves activities that are relatively skilled-labour intensive in the host country, even though they are relatively unskilled-labour intensive in the source country.

Several empirical studies support the hypothesis derived from endowment-driven North-South models, according to which FDI is associated with greater inequality by raising the skill premium in poorer host countries. The case of Mexico has received particular attention (Aitken et al. 1996; Feenstra and Hanson 1997; Hanson 2003). Additionally, FDI has benefited skilled workers more than unskilled workers in other emerging economies, including Indonesia (Lipsey and Sjöholm 2004) and Korea (Mah 2002). Cross-country studies covering a varying number of developing countries produced more ambiguous findings (e.g., Tsai 1995; Sylwester 2005; Choi 2006).

North-South FDI models are less useful for analysing the distributional consequences of FDI in the richer host countries in our sample; vertical FDI in these models tends to result from large wage differentials between the source and the host country. However, the argument of Markusen (1995), that advanced countries with similar relative factor endowments host the bulk of FDI, applies in the European context, too. For instance, just five EU countries accounted for two thirds of total inward FDI stocks in the EU in 2000, even though their average per-capita income varied by less than 10 per cent at that time (UNCTAD 2010).[1] Markusen (1995) posits that advanced countries attract mainly market-oriented, horizontal types of FDI. Blonigen and Slaughter argue that "it is quite difficult to assess theoretically whether and how this [type of] inward FDI should affect wages."

The relevant theoretical literature departs from the observation that multinational enterprises (MNEs) possess firm-specific assets, such as technological knowledge and management skills, granting them a productivity advantage over domestic firms. MNEs can then be regarded as vehicles for transferring new technologies, know-how, and ideas to the host country, acting "as 'role models' for indigenous firms" (Figini and Görg 1999: 596). Yet, recent MNE models make different predictions on the distributional effects of FDI in advanced countries. Assuming that the operations of MNEs are more skill intensive than the operations of domestic firms, MNEs could lead to greater income inequality by increasing the demand for skilled labour (see, e.g., Taylor and Driffield 2005).

[1] Belgium, France, Germany, the Netherlands, and the United Kingdom.

The opposite conclusion, that FDI results in less inequality, can be reached, however, when refining the ranking of skill intensities at the firm and plant level. Markusen and Venables (1997) consider headquarter (HQ) services to be more skill intensive than plant operations. It has also to be taken into account that advanced countries are typically host, as well as home, to (foreign and, respectively, domestic) MNEs. Under such conditions, the establishment of foreign plant operations through FDI may reduce the relative demand for skilled labour in the host country. This is most likely to happen where foreign plant operations are considerably less skill intensive than the HQ services supplied from the foreign MNEs' home base, and where the HQ services of various domestic MNEs have traditionally shaped the demand for skilled labour. The United States spring to mind as a case in point (Blonigen and Slaughter 2001), but advanced European countries are also likely to fit into this reasoning.

Theoretical predictions on the distributional effects of FDI in advanced host countries become still more complex when accounting for learning and skill upgrading in the "transition to a new technological paradigm" (Aghion and Howitt 1998: 262). While domestic firms may benefit from FDI-induced spillovers, their absorption of new technologies may increase inequality in the short run and reduce inequality in the longer run. Aghion and Howitt (1998, chapter 8) model such a transition by explicitly referring to the Kuznets inverted-U hypothesis of rising and then falling inequality. Accordingly, the skill premium increases as long as learning efforts result in a high demand for skills that are in short supply. Subsequently, wage inequality declines to the extent that the supply of the required skills improves and firms have managed the transition to the new technological paradigm.

Empirical studies addressing these theoretical predictions are still few. Some country-specific analyses report ambiguous findings. For the United States, Blonigen and Slaughter (2001) do not find any evidence that inward FDI contributed to skill upgrading in manufacturing industries until the mid-1990s. Chintrakarn et al. (2011) perform panel co-integration analysis across the US states; FDI at the state level reduced income inequality during the period 1977 to 2001, on average, but the effects proved to be heterogeneous. By contrast, Taylor and Driffield (2005) show for the United Kingdom that FDI had a positive impact on wage inequality within industries. According to Figini and Görg (1999), the Irish case reveals an inverted U-shaped pattern, with FDI first increasing and then later reducing inequality.

In addition to country-specific studies, some indications exist that the distributional effects of FDI differ between advanced and developing host countries. Gopinath and Chen (2003) analyse time-series data on FDI and wages for 15 advanced and 11 developing countries. Inward FDI is associated with higher labour shares in GNP in both sub-samples, but F tests reveal that FDI effects differ significantly between advanced and developing host countries. These authors also report some evidence that FDI increases the wage gap between skilled and unskilled workers in developing countries. However, they do not address the wage gap in advanced host countries. Similarly, Figini and Görg (2006) find that wage inequality increases with FDI in developing host countries, though this effect diminishes with further increases in FDI. By contrast, inequality decreases with FDI in advanced host countries.

While focusing on relatively advanced European host countries, the varying level of economic development within our sample offers more insights into the heterogeneity of FDI effects on income inequality. Furthermore, the subsequent analysis helps overcome several shortcomings of almost all previous studies. With the exception of Chintrakarn et al. (2011), earlier analyses are typically restricted to wage disparity within the manufacturing sector. Including FDI in services is warranted, and not only because it is quantitatively important; its distributional effects may also differ from those of FDI in manufacturing, as vertical FDI is more likely to exist in manufacturing than in services. Finally, as stressed by Lindert and Williamson (2001) and Basu and Guariglia (2007), wage inequality provides an incomplete picture of income inequality.

3. Empirical model and data

The analysis will examine the relationship between FDI and income inequality in Europe. This section presents the basic empirical model, discusses some empirical issues, and describes the data.

3.1. Empirical specification and econometric issues

Following Chintrakarn et al. (2011), we assume that the correct specification of the long-run relationship between income inequality and FDI is given by a bivariate model of the form

$$Inequality_{it} = a_i + \delta_i t + b(FDI/GDP)_{it} + \varepsilon_{it}, \tag{1}$$

where $i = 1, 2, ..., N$ is the country index, $t = 1, 2, ..., T$ is the time index, $Inequality_{it}$ stands for the estimated household income inequality (EHII) in Gini format (measured on a 0 to 100 scale), and $(FDI/GDP)_{it}$ represents the percentage share of FDI in GDP. As in Figini and Görg (2006) and Chintrakarn et al. (2011), we use FDI stocks, rather than FDI flows. Stocks should capture long-run effects more effectively than annual FDI flows, which fluctuate considerably. The coefficient b represents the long-run effect of FDI on inequality. The a_i and $\delta_i t$ are, respectively, country-specific fixed effects and country-specific deterministic time trends, capturing any country-specific omitted factors that are relatively stable or which evolve smoothly over time.

An important feature of Equation (1) is the assumption that, in the long-run, permanent changes in the FDI-to-GDP ratio are associated with permanent changes in the inequality index. Empirically, this implies that both the individual time series for the EHII Gini coefficient and FDI (relative to GDP) must exhibit unit-root behaviour, and that $Inequality_{it}$ must be co-integrated with $(FDI/GDP)_{it}$.[2] A regression consisting of co-integrated variables has a stationary error term, ε_{it}, implying that no relevant integrated variables are omitted; any omitted non-stationary variable that is part of the co-integrating relationship would enter the error term, thereby producing non-stationary residuals and a failure to detect co-integration. If there is co-integration between a set of variables, this same stationary relationship also exists in extended variable space. Thus, an important implication of finding co-integration is that no relevant integrated variables in the co-integrating vector are omitted. Co-integration estimators are therefore robust (under co-integration) to the omission of variables that do not form part of the co-integrating relationship. This justifies a reduced-form model such as Equation (1) (if co-integrated).

Another assumption inherent in Equation (1) is that inequality is endogenous in the sense that, in the long run, changes in FDI cause changes in income inequality. However, although the existence of co-integration implies long-run Granger causality in at least one direction, causality may also run from inequality to FDI. For example, growing income inequality may reflect a decline in the real wages of less-skilled workers. Under such conditions, multinationals may undertake vertical FDI and locate their low-skilled activities in countries with a higher level of inequality in order to take advantage of lower wages. On the other hand, foreign investors may shy away from countries with greater levels of income inequality, fearing social conflict and political instability. Horizontal FDI may also be discouraged to the extent that greater inequality implies limited effective demand because of a shrinking middle class. It follows that greater inequality could be the cause of increased or reduced FDI activity. The empirical implication is that, in addition to examining the time-series properties of the variables and testing whether the variables are co-integrated, it is also important to deal with this potential endogeneity problem and to investigate the direction of causality.

[2] Strictly speaking, of course, the stochastic process for Gini coefficients cannot be a pure unit-root process. Gini inequality measures are bounded, but we know that a unit-root process will cross any finite bound with probability 1. Nevertheless, as argued by Jones (1995), it may be the case that in the relevant range, such variables are well characterised by a unit-root process. Specifically, if the determining factors of Gini coefficients, such as tastes, time preferences, and government policies, change over time, we observe Gini coefficient series with permanent movements that can be well approximated by a unit-root process. For example, the individual country unit-root tests used by Guest and Swift (2008) suggest that the Gini coefficients for all countries in their study (UK, USA, Australia, Japan, and Sweden) are I(1). Thus, it may well be that inequality measures and FDI-to-GDP ratios are integrated and co-integrated, as the results by Chintrakarn et al. (2011) for the United States suggest.

Cross-country heterogeneity represents a further econometric issue to be addressed. As discussed in Section 2, theoretical arguments suggest that the effect of FDI on inequality may be highly heterogeneous across countries. Moreover, individual country studies tend to find different results across regions; the study by Chintrakarn et al. (2011), in particular, suggests that the inequality effects of FDI are not constant across the US states. Given that traditional homogeneous panel estimators produce inconsistent and potentially misleading point estimates of the sample mean of the co-integrating vectors when the true slope coefficients are heterogeneous (see, e.g., Pesaran and Smith 1995), we use heterogeneous panel estimators based on the mean group approach. This also allows us to explicitly analyse the potential heterogeneity in the effects of FDI across countries.

A final empirical issue is the potential non-linearity in the relationship between FDI and inequality. As discussed above, the model of Aghion and Howitt (1998) predicts that wage dispersion first increases after the introduction of new technologies and then decreases (due to learning processes). This implies that the effect of FDI on inequality may differ, not only from country to country, but also over time. More specifically, one might expect that the long-run effect differs significantly from the short-run effect. In our analysis, we therefore attempt to provide insight into the dynamics of FDI over time by examining both the long-run and short-run effects of FDI on income inequality.

3.2. Data and descriptive statistics

We now briefly describe the data. The data on $(FDI/GDP)_{it}$ come from UNCTAD's FDI database (available at http://www.unctad.org/templates/Page.asp?IntItemID=3277&lang=1). The EHII data are taken from the University of Texas Inequality Project (available at http://utip.gov.utexas.edu/data.html). The major advantage of the EHII source is that the data are fully comparable across space and time. The EHII data set combines information from the United Nations Industrial Development Organization (UNIDO) with information from the Deininger and Squire data set; it also uses other relevant information, such as the ratio of manufacturing employment to total population, the degree of urbanisation, and population growth.[3]

[3] See Galbraith (2009) for a detailed discussion of the data set and its construction.

Table 1

Countries and country summary statistics

Country	Average of $Inequality_{it}$	Average of $(FDI/GDP)_{it}$
Finland	31.72	5.37
Germany	33.51	6.34
Ireland	39.58	112.57
Italy	35.30	5.03
Malta	33.78	23.91
Netherlands	33.02	24.47
Norway	33.00	12.63
Spain	37.13	12.22
Sweden	27.71	9.61
United Kingdom	32.40	17.28

The identification and estimation of co-integrating relationships requires the use of data over a sufficiently long period of time. Panel co-integration methods can be implemented with shorter data spans than their time-series counterparts due to exploitation of both the time-series and cross-sectional dimensions of the data. Consequently, a period of about 20 observations should be long enough in the present panel context. We include all countries for which continuous data are available over the period from 1980 to 2000 (21 years), resulting in a balanced panel of 210 observations on 10 countries.

Table 1 lists the countries along with the average values for $Inequality_{it}$ and $(FDI/GDP)_{it}$ over the period of observation. Ireland is the country with the highest inequality, followed by Spain and Italy, while Sweden ranks at the bottom of the inequality scale. Ireland also heads the list of FDI recipients (relative to GDP), followed by the Netherlands and Malta. The countries with the lowest FDI-to-GDP ratios are Germany, Finland, and Italy.

4. Empirical analysis

The pre-tests for unit roots and co-integration are reported in the Appendix. They suggest that the variables are non-stationary and co-integrated, as assumed in Equation (1). In this section, we provide estimates of the co-integrating relationship between FDI and income inequality. We also investigate the direction of causality between the two variables, explore the short-run dynamics of the relationship, and examine the degree of heterogeneity in the effects of FDI on income inequality across countries.

4.1. Long-run relationship

We estimate the long-run effect of FDI on income inequality using the between-dimension group-mean panel DOLS estimator that Pedroni (2001) argues has a number of advantages over the within-dimension approach. First, it allows for greater flexibility in the presence of heterogeneous co-integrating vectors, whereas under the within-dimension approach, the co-integrating vectors are constrained to being the same for each country. This is an important advantage for applications such as the present one, as the effect of FDI on income inequality cannot reasonably be assumed to be the same across countries. Another advantage of the between-dimension estimators is that the point estimates provide a more useful interpretation in the case of heterogeneous co-integrating vectors; they can be interpreted as the mean value of the co-integrating vectors, which is not appropriate for the within estimators. Finally, the between-dimension estimators suffer from much lower small-sample-size distortions than is the case with the within-dimension estimators.

The DOLS regression is given by

$$Inequality_{it} = a_i + \delta_i t + b_i (FDI/GDP)_{it} + \sum_{j=-p_i}^{p_i} \Phi_{ij} \Delta (FDI/GDP)_{it-j} + \varepsilon_{it}, \qquad (2)$$

where Φ_{ij} are coefficients of lead and lag differences, which account for possible serial correlation and endogeneity of the regressor(s), thus yielding unbiased estimates. Thus, an important feature of the DOLS procedure is that it generates unbiased estimates for variables that co-integrate, even with endogenous regressors. In addition, the DOLS estimator is superconsistent under co-integration, and it is also robust to the omission of variables that do not form part of the co-integrating relationship.

From regression (2), the group-mean DOLS estimator for b is constructed as

$$\hat{b} = \left[N^{-1} \sum_{i=1}^{N} \left(\sum_{t=1}^{T} z_{it} z'_{it} \right)^{-1} \left(\sum_{t=1}^{T} z_{it} \tilde{s}'_{it} \right) \right]_1, \qquad (3)$$

where z_{it} is the $2(K+1)\times 1$ vector of regressors $z_{it} = [(FDI/GDP)_{it} - \overline{(FDI/GDP)_i}$, $\Delta(FDI/GDP)_{it-K}$, …, $\Delta(FDI/GDP)_{it+K}]$, $\tilde{s}_{it} = s_{it} - \bar{s}_i$, and the subscript 1 outside the brackets indicates that only the first element of the vector is taken to obtain the pooled slope coefficient. Because the expression following the summation over the i is identical to the conventional time-series DOLS estimator, the between-dimension estimator for b is calculated as

$$\hat{b} = N^{-1} \sum_{i=1}^{N} \hat{b}_i, \qquad (4)$$

where $\hat{b}_i$ is the conventional DOLS estimator applied to the ith country of the panel and $t_{\hat{b}} = N^{-1/2} \sum_{i=1}^{N} t_{\hat{b}_i}$ is the associated t statistic. As found by Stock and Watson (1993), this estimator performs well in short time series compared to other co-integration estimators such as the maximum likelihood estimator of Johansen (1988), or the fully modified ordinary least squares (FMOLS) estimator of Phillips and Hansen (1990).

The DOLS group-mean estimate of the long-run relationship is as follows (t statistics in parentheses; ** and (*) denote significance at the 1% (5%) level):[4]

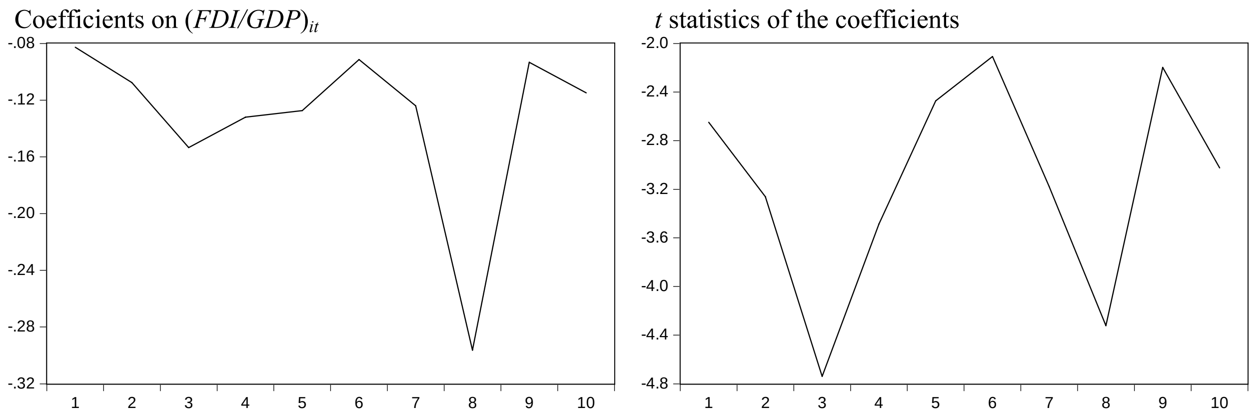
$$\begin{aligned} Inequality_{it} &= \underset{(103.59)}{31.604^{**}} + \underset{(12.79)}{0.261^{**}}t - \underset{(-3.31)}{0.132^{**}}(FDI/GDP)_{it} \end{aligned} \quad . \tag{5}$$

The coefficient on FDI is highly significant and negative, suggesting that FDI reduces income inequality in European host countries (on average) in the long run. This result is consistent with the findings of Chintrakarn et al. (2011) for the US states and supports the notion that the effects of FDI on inequality in advanced countries tend to differ from those in developing countries.

Given the small number of countries in our sample, we need to ensure that the negative average effect on inequality is not due to outliers. To this end, we re-estimate the DOLS regression, excluding one country at a time from the sample. The sequentially estimated coefficients and their t statistics are presented in Figure 1. They fluctuate between -0.29 (due to the exclusion of Spain)[5] and -0.08 and are always significant at least at the 5-per-cent level, suggesting that the negative effect of FDI on income inequality is not the result of possible outliers.

Figure 1

DOLS estimation with single country excluded from the sample



4.2. Short-run and long-run causality

We are interested in detecting the direction of long-run causality, even though estimation by DOLS does not require the regressor(s) to be exogenous (and co-integration implies long-run Granger causality in at least one direction). Specifically, the arguments in Section 3.1 suggest that causality may run in both directions, that is, not only from FDI to inequality but also from

[4] The number of leads and lags in the individual DOLS regressions was determined by the Schwarz criterion with a maximum of two lags.

[5] The individual country DOLS estimates are presented in Table 3.

inequality to FDI. To test the direction of long-run causality and to examine the short-run dynamics between the variables, we enter the residuals from the individual DOLS long-run relations,

$$ec_{it} = Inequality_{it} - [\hat{a}_i + \hat{\delta}_i t + \hat{b}_i\,(FDI/GDP)_{it}], \tag{6}$$

as error-correction terms into a simple panel vector-error-correction model (VECM) of the form

$$\begin{bmatrix} \Delta Inequality_{it} \\ \Delta(FDI/GDP)_{it} \end{bmatrix} = \begin{bmatrix} c_{1i} \\ c_{2i} \end{bmatrix} + \sum_{j=1}^{k}\Gamma_j \begin{bmatrix} \Delta Inequality_{it-j} \\ \Delta(FDI/GDP)_{it-j} \end{bmatrix} + \begin{bmatrix} a_1 \\ a_2 \end{bmatrix} ec_{it-1} + \begin{bmatrix} \varepsilon_{1it} \\ \varepsilon_{2it} \end{bmatrix}, \tag{7}$$

where the c_is are fixed effects; the error-correction term, ec_{it-1}, represents the error in, or deviation from, the equilibrium; and the adjustment coefficients a_1 and a_2 capture how $Inequality_{it}$ and $(FDI/GDP)_{it}$ respond to deviations from the equilibrium relationship. From the Granger representation theorem, we know that at least one of the adjustment coefficients must be non-zero if a long-run relationship between the variables is to hold. A significant error-correction term also suggests long-run Granger causality and thus, long-run endogeneity (see, e.g., Hall and Milne 1994), whereas a non-significant adjustment coefficient implies long-run Granger non-causality from the independent to the dependent variable(s), as well as weak exogeneity. Following Herzer (2008), we test for weak exogeneity by first eliminating the insignificant short-run dynamics in the model successively, according to the lowest t values, and then deciding on the significance of the error-correction term. In doing so, we reduce the number of parameters and thereby increase the precision of the weak exogeneity tests on the α coefficients. Since all variables in Equation 7, including ec_{it-1}, are stationary (because the level variables are integrated of order 1 and co-integrated), conventional t tests can be used for this purpose.

Table 2

Vector-error-correction model, long-run causality, and short-run dynamics

Independent variables	Dependent variable $\Delta Inequality_{it}$	Dependent variable $\Delta(FDI/GDP)_{it}$
ec_{it-1}	-0.419** (-6.45)	0.792* (2.37)
$\Delta Inequality_{it-1}$	-0.131* (-2.00)	—
$\Delta Inequality_{it-2}$	—	—
$\Delta(FDI/GDP)_{it-1}$	0.040* (2.19)	0.399** (3.77)
$\Delta(FDI/GDP)_{it-2}$	—	0.223* (2.07)

Notes: ** and (*) indicate significance at the 1% (5%) level. t statistics are in parentheses. The number of lags was determined by the general-to-specific procedure with a maximum of two lags. Given that the insignificant short-run dynamics were eliminated successively according to the lowest t values, they are not reported here.

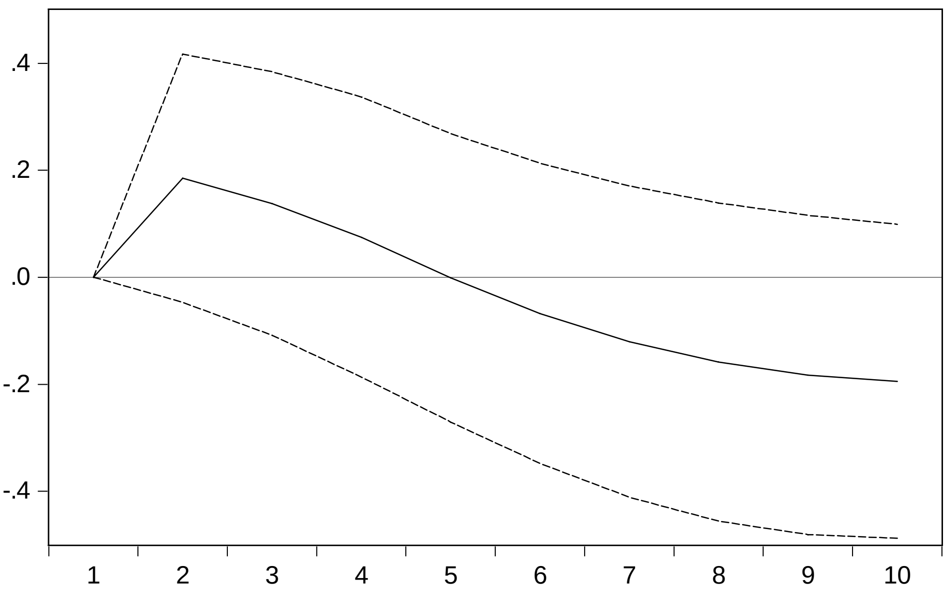
Table 2 presents the results. As the t statistics for the error-correction terms indicate, the null hypothesis of weak exogeneity can be rejected for both $Inequality_{it}$ and $(FDI/GDP)_{it}$, at least at the

5-per-cent level. Both income inequality and FDI are thus endogenous. From this it can be concluded that the statistical long-run causality is bidirectional, in turn suggesting that decreased inequality is both a consequence and a cause of increased FDI.

Another important result is that the coefficient on $\Delta(FDI/GDP)_{it\text{-}1}$ is statistically significant and positive in the equation with $\Delta Inequality_{it}$ as the dependent variable (Table 2, second column). This indicates that FDI has a *positive* causal effect on income inequality in the short run, while its long-run effect on income inequality is negative.

To test the robustness of this conclusion, we compute the impulse-response functions from the VECM residuals, applying a standard Choleski decomposition (with $\Delta Inequality_{it}$ ordered first and $\Delta(FDI/GDP)_{it}$ ordered last). The responses of income inequality to a one-standard-deviation innovation in FDI over a 10-year horizon are shown in Figure 2, where the dashed lines mark plus and minus two standard errors obtained through Monte Carlo simulations using 1,000 replications. The figure reveals that a one-standard-deviation innovation in FDI first produces an increase in income inequality, reaching a peak in two years. After four years, however, income inequality gradually and permanently decreases in response to a one-standard-deviation innovation in the FDI-to-GDP ratio. Consequently, the impulse responses are consistent with the above finding that FDI has a positive effect on inequality in the short run, whereas the long-run effect of FDI on inequality is negative.

Figure 2. Response of inequality to FDI



4.4. Individual country effects

The results reported so far indicate that FDI has, on average, a negative long-run effect on income inequality in Europe (and vice versa). This finding for the sample as a whole does not necessarily imply, however, that FDI affects inequality negatively in each individual country. Table 3 (second and third columns) presents the individual DOLS country estimates of the coefficients on $(FDI/GDP)_{it}$. These estimates should be interpreted with caution, given the relatively short time dimension of our data. However, several previous co-integration analyses for individual countries are based on smaller time samples (see, e.g., de Crombrugghe et al. 1997; Irvin and Izurieta 2000). More importantly, we also report the results of the Johansen (1988) maximum likelihood estimator (fourth and fifth columns) to ensure that our conclusions are robust to different estimation techniques. Note that the Johansen procedure corroborates that FDI has, on average, a negative long-run effect on income inequality.

At the same time, there appears to be considerable heterogeneity in the long-run effects of FDI on income inequality across countries. Specifically, we find that in eight countries—Finland, Germany, Italy, Malta, the Netherlands, Norway, Sweden, and the United Kingdom—an increase in FDI is associated with a decrease in income inequality, while in two cases—Ireland and Spain—an increase in FDI is associated with an increase in inequality. This heterogeneity appears to be largely in line with the reasoning in Section 2. Accordingly, countries with relatively low per-capita income, such as Ireland and Spain, are most likely to attract vertical FDI.[6] The prediction of North-South models, that inequality widens with FDI, should thus be relevant for these cases, in contrast to the larger part of our sample, which is characterised by higher per-capita incomes and similar factor endowments.

Table 3

Individual country estimates

	DOLS estimator		Johansen maximum-likelihood estimator	
	$(FDI/GDP)_{it}$	*t* stat.	$(FDI/GDP)_{it}$	*t* stat.
Finland	-0.576*	-2.511	-0.427**	-5.107
Germany	-0.351	-0.677	-1.616**	-4.683
Ireland	0.061**	3.760	0.051**	3.345
Italy	-0.156	-0.178	-1.398+	-1.900
Malta	-0.174**	-3.041	-0.011	-0.292
Netherlands	-0.498**	-4.137	-0.362**	-4.267
Norway	-0.204	-0.928	-3.421**	-5.724
Spain	1.346*	2.507	3.990**	5.248
Sweden	-0.481**	-3.867	-0.298**	-8.382
United Kingdom	-0.287	-1.387	-0.111**	-7.939
Average	-0.132		-0.360	

Notes: **, (*), and [+] indicate significance at the 1% (5%) [10%] level

[6] Recall that Malta represents the third country in our sample with relatively low per-capita income. Unlike Ireland and Spain, however, Malta is not known for having attracted considerable FDI in manufacturing, where vertical FDI figures most prominently. FDI in Malta is probably concentrated in tourism.

5. Conclusion

Trade-offs associated with FDI have contributed to waning public support of economic globalisation. Concerns are primarily about outward FDI helping firm-level competitiveness at the expense of wage pressure or unemployment of low-skilled workers. Inward FDI may involve similar distributional conflicts. This would be the case if productivity-enhancing spillovers went along with widening inequality due to shifts in the relative demand for labour towards higher skills. In Europe, possible trade-offs between average income gains and greater inequality typically rank high on the political agenda. However, strikingly different predictions can be derived from recent theoretical models, and reliable empirical evidence on the distributional effects of FDI is largely lacking. This applies to inward FDI in particular.

The present study contributes to filling this gap. We examined the macroeconomic relationship between FDI and income inequality for a sample of 10 European host countries over the period 1980 to 2000 by employing panel co-integration and causality techniques that are robust to omitted variables, slope heterogeneity, and endogenous regressors. Our results suggest that: (1) FDI has a positive short-run effect on income inequality in Europe, (2) the long-run effect of FDI on inequality, however, is negative on average, (3) long-run causality runs in both directions, suggesting that an increase in FDI reduces income inequality and that, in turn, higher inequality leads to lower FDI inflows, and (4) there are large differences in the long-run effect of FDI on income inequality, with Ireland and Spain exhibiting a positive relationship between FDI in income inequality.

On average, the European experience leads to the conclusion that policymakers competing for inward FDI need not be concerned that productivity and growth promoting access to superior technology and know-how generally deepens the divide between rich and poor segments of the workforce. Possible trade-offs typically appear to be short-lived. In the longer term, relatively few European host countries run the risk of giving rise to social conflict by attracting FDI. Yet this heterogeneity deserves to be analysed in more detail. In particular, country-specific studies may offer deeper insight into the varying effects different types of FDI may have on income inequality. Distinguishing between vertical and horizontal FDI provides one possible avenue for further research, but the distributional effects may also vary between FDI from different sources and depend on the local environment in the host country, including labour market conditions.

Appendix: Panel unit root and cointegration tests

A.1. Panel unit-root tests

To examine the time-series properties of the data, we use the between-dimension panel unit root test of Im, Pesaran, and Shin (2003) (IPS), which is based on the augmented Dickey-Fuller (ADF) regression for the *i*th cross-section unit:

$$\Delta x_{it} = z_{it}'\gamma + \rho_i x_{it-1} + \sum_{j=1}^{p_i} \varphi_{ij} \Delta x_{it-j} + \varepsilon_{it}, \tag{A.1}$$

where p_i is the lag order and z_{it} represents deterministic terms, such as fixed effects or fixed effects combined with individual time trends. The IPS test tests the null hypothesis of a unit root for all i, $H_0: \rho_i = 0$, against the alternative of (trend) stationarity, $H_1: \rho_i < 0$, i = 1, 2, …, N_1; $\rho_i = 0$, $i = N_1 + 1$, $N_1 + 2$, …, N, using the standardised t bar or IPS statistic

$$\Gamma_{\bar{t}} = \frac{\sqrt{N}[\bar{t}_{NT} - \mu]}{\sqrt{v}}, \tag{A.2}$$

where $\bar{t}_{NT}$ is the average of the N = 10 cross-sectional ADF t statistics, and μ and v are, respectively, the mean and variance of the average of the individual t statistics, tabulated by Im et al. (2003).

Table A.1

Panel unit-root tests

Variables	Deterministic terms	IPS statistics	CIPS statistics
Levels			
$Inequality_{it}$	constant, trend	0.02	-2.02
$(FDI/GDP)_{it}$	constant, trend	6.99	-1.94
First differences			
$\Delta Inequality_{it}$	constant	-4.82**	-2.65**
$\Delta(FDI/GDP)_{it}$	constant	-2.24*	-2.41*

Notes: The number of lags was determined by the Schwarz criterion with a maximum of two lags. The relevant 1% (5%) critical value for the CIPS statistics is -3.15 (-2.88) with an intercept and a linear trend, and -2.60 (-2.34) with an intercept. ** and (*) denote significance at the 1% (5%) level.

However, the standard IPS test can lead to spurious inferences if the errors, ε_{it}, are not independent across i, for example, due to common shocks or spillovers between countries. Therefore, we also employ the cross-sectionally augmented IPS test proposed by Pesaran (2007). This test is designed to filter out the cross-section dependency by augmenting the ADF regression with the cross-section averages of lagged levels and first differences of the individual series. The attractive feature of this test is thus that it permits the individual countries to respond differently to the common time effects as reflected by the country-specific coefficients on the cross-section averages of the variables. Accordingly, the cross-sectionally augmented ADF (CADF) regression is given by

$$\Delta x_{it} = z_{it}'\gamma + \rho_i x_{it-1} + \sum_{j=1}^{p_i} \varphi_{ij} \Delta x_{it-j} + \alpha_i \bar{x}_{t-1} + \sum_{j=0}^{p_i} \eta_{ij} \Delta \bar{x}_{t-j} + v_{it}, \tag{A.3}$$

where $\bar{x}_t$ is the cross-section mean of x_{it}, $\bar{x}_t = N^{-1}\sum_{i=1}^{N} x_{it}$. The cross-sectionally augmented IPS statistic is the simple average of the individual CADF statistics and is defined as

$$CIPS = t\text{-bar} = N^{-1}\sum_{i=1}^{N_i} t_i, \tag{A.4}$$

where t_i is the OLS t ratio of ρ_i in Equation A.4. The corresponding critical values are given by Pesaran (2007).

The results for the variables in levels and in first differences are presented in Table A1. While neither test rejects the null hypothesis of a unit root in levels, the unit-root hypothesis is rejected for the first differences. We therefore conclude that $Inequality_{it}$ and $(FDI/GDP)_{it}$ are integrated of order 1 (the necessary condition for co-integration in a bivariate context).

A.2. Panel co-integration tests

We use several panel co-integration tests to determine whether there is a long-run relationship between FDI and income inequality. The first is the two-step residual-based procedure suggested by Pedroni (1999, 2004), which can be intuitively described as follows. In the first step, the hypothesised co-integrating regression

$$Inequality_{it} = a_i + \delta_i t + b_i (FDI/GDP)_{it} + \varepsilon_{it} \tag{A.5}$$

is estimated separately for each country, thus allowing for heterogeneous co-integrating vectors. In the second step, the residuals from these regressions, $\hat{\varepsilon}_{it}$, are tested for stationarity. To test the null hypothesis of non-stationarity (or no co-integration), Pedroni proposes seven statistics. Four of these test statistics pool the autoregressive coefficients across different countries during the unit-root test and thus constrain the autoregressive parameters to be homogeneous across countries. Pedroni refers to these within-dimension-based statistics as panel co-integration statistics. The other three test statistics are based upon estimators that average the individually estimated autoregressive coefficients for each country, thus allowing the autoregressive coefficient to be heterogeneous across countries. Pedroni refers to these between-dimension statistics as group-mean panel co-integration statistics. The first of the panel co-integration statistics is a non-parametric variance ratio test. The second and the third are panel versions of the Phillips and Perron (PP) *rho* statistic and *t* statistic, respectively. The fourth statistic is a panel ADF *t* test analogous to the Levin et al. (2002) panel unit-root test. Similarly, the first of the group-mean panel co-integration statistics is analogous to the PP *rho* statistic, the second is a panel version of the PP *t* statistic, and the third is a group mean ADF *t* test analogous to the IPS (2003) panel unit-root test (described above). The standardised distributions for the panel and group statistics are given by

$$\kappa = \frac{\varphi - \mu\sqrt{N}}{\sqrt{v}} \Rightarrow N(0, 1)\,, \tag{A.6}$$

where φ is the respective panel or group statistic, and μ and v are the expected mean and variance of the corresponding statistic, tabulated by Pedroni (1999).

A potential problem with the Pedroni approach is that it does not take into account potential error cross-sectional dependence, which could bias the results. To test for co-integration in the presence of possible cross-sectional dependence, we use the two-step procedure suggested by Holly et al. (2010). In the first step, we apply the common correlated effects (CCE) estimator of Pesaran (2006) to the static co-integrating regression. Like the cross-sectionally augmented IPS test, the CCE estimator allows for cross-sectional dependencies that potentially arise from multiple unobserved common factors and permits the individual responses to these factors to differ across countries. The cross-sectionally augmented co-integrating regression for the *i*th cross-section is given by

$$Inequality_{it} = a_i + \delta_i t + b_i (FDI/GDP)_{it} + g_{1i}\overline{Inequality_t} + g_{2i}\overline{(FDI/GDP)_t} + e_{it}\,, \tag{A.7}$$

where the cross-sectional averages $\overline{Inequality_t}$ = $N^{-1}\sum_i^N Inequality_{it}$ and $\overline{(FDI/GDP)_t}$ = $N^{-1}\sum_i^N (FDI/GDP)_{it}$ serve as proxies for the unobserved factors. In the second step, we compute the cross-sectionally augmented IPS statistic for the residuals from the individual CCE long-run relations, $\hat{\mu} = Inequality_{it} - \hat{\delta}_i t - \hat{b}_i\,(FDI/GDP)_{it}$, including an intercept. In doing so, we account for unobserved common factors that could be correlated with the observed regressors in both steps.

Residual-based (panel) co-integration tests, such as the ones just described, restrict the long-run elasticities to being equal to the short-run elasticites. If this restriction is invalid, residual-based (panel) co-integration tests may suffer from low power (see, e.g., Westerlund, 2007). Another drawback of residual-based (panel) co-integration tests is that they are generally not invariant to the normalisation of the co-integrating regression. Therefore, we also use the Larsson et al. (2001) procedure, which is based on Johansen's (1988) maximum likelihood estimation procedure. Like the Johansen time-series co-integration test, the Larsson et al. panel test treats all variables as potentially endogenous, thus avoiding the normalisation problems inherent in residual-based co-integration tests. In addition, the Larsson et al. procedure allows the long-run elasticities to differ from the short-run elasticities and hence does not impose a possibly invalid common factor restriction. It involves estimating the Johansen vector-error-correction model for each country separately:

$$\Delta y_{it} = \prod_i y_{it-1} + \sum_{i=1}^{k_i} \Gamma_{ik} \Delta y_{it-k} + z_{it}\gamma_i + \varepsilon_{it}, \tag{A.8}$$

where y_{it} is a $p \times 1$ vector of endogenous variables (y_{it} = [$Inequality_{it}$, $(FDI/GDP)_{it}$]'; p is the number of variables), and Π_i is the long-run matrix of order $p \times p$. If Π_i is of reduced rank, $r_i < p$**,** it is possible to let $\Pi_i = \alpha_i \beta_i$, where β_i is a $p \times r_i$ matrix, the r_i columns of which represent the co-integrating vectors, and α_i is a $p \times r_i$ matrix having p rows which represent the error-correction coefficients. The null hypothesis is that all of the N countries in the panel have a common co-integrating rank, i.e., at most r (possibly heterogeneous) co-integrating relationships among the p variables: $H_0 : rank(\Pi_i) = r_i \le r$ for all $i = 1, \dots, N$. The alternative hypothesis is that all the cross-sections have a higher rank: $H_1 : rank(\Pi_i) = p$ for all $i = 1, \dots, N$. To test H_0 against H_1, a panel co-integration rank trace-test statistic is computed by calculating the average of the individual trace statistics, $LR_{iT}\{H(r)|H(p)\}$:

$$\overline{LR}_{NT}\{H(r)|H(p)\} = \frac{1}{N}\sum_{i=1}^{N} LR_{iT}\{H(r)|H(p)\}, \tag{A.9}$$

and then standardising it as follows:

$$\Psi_{\overline{LR}}\{H(r)|H(p)\} = \frac{\sqrt{N}\left(\overline{LR}_{NT}\{H(r)|H(p)\} - E(Z_k)\right)}{\sqrt{Var(Z_k)}} \Rightarrow N(0,1). \tag{A.10}$$

The mean $E(Z_k)$ and variance $Var(Z_k)$ of the asymptotic trace statistic are tabulated by Breitung (2005) for the model (with an intercept and a trend) we use.

Table A.2

Panel co-integration tests

Pedroni (1999, 2004) cointegration test statistics		
Panel *v* statistic	2.12*	
Panel PP *rho* statistic	-3.21**	
Panel PP *t* statistic	-5.49**	
Panel ADF *t* statistic	-5.45**	
Group PP *rho* statistic	-1.26	
Group PP *t* statistic	-6.07**	
Group ADF *t* statistic	-6.21**	
CIPS statistic for the residuals of the CCE long-run relations	-2.75**	
	Co-integration rank	
	$r = 0$	$r = 1$
Panel trace statistics	3.58**	0.06
Fisher statistics	37.40*	19.75

Notes: ** and (*) indicate a rejection of the null of no co-integration at the 1% (5%) level. Under the alternative hypothesis, the panel *v* statistic diverges to positive infinity so that the right tail of the normal distribution is used to reject the null hypothesis. The relevant 1% critical value for the CIPS statistic is -2.60. The number of lags was determined by the Schwarz criterion with a maximum of two lags. The Fisher statistic is distributed as χ^2 with $2 \times N$ degrees of freedom. It has a critical value of 37.57 (31.41) at the 1% (5%) level.

In addition, we compute the Fisher statistic proposed by Madalla and Wu (1999), which is defined as

$$\lambda = -2\sum_{i}^{N} \log(p_i) , \tag{A.11}$$

where p_i is the p value of the trace statistic for country i, calculated from the response surface estimates in MacKinnon et al. (1999). The Fisher statistic is distributed as χ^2 with $2 \times N$ degrees of freedom.

The results of these tests are presented in Table A2. Six of the seven Pedroni statistics reject the null of no co-integration at conventional levels. Specifically, the ADF-type tests reject the null hypothesis at the 1-per-cent level. Given that these tests have been shown to have the highest power for smaller sample sizes, such as $T = 21$ (see, e.g., Pedroni 2004), the ADF test results, in particular, provide strong evidence of co-integration. This conclusion is supported by the CIPS, the panel trace, and the Fisher statistics, which show that $Inequality_{it}$ and $(FDI/GDP)_{it}$ are co-integrated (or exhibit a single co-integrating vector).

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