WP2/17 SEARCH WORKING PAPER

The Effect of FDI on Regional Inequality in the ENPs; Evidence from Israel

Michael Beenstock, Daniel Felsenstein and Ziv Rubin

September 2013

SEARCH
SHARING KNOWLEDGE ASSETS: INTERREGIONALLY COHESIVE NEIGHBORHOODS

EUROPEAN COMMISSION
European Research Area

SEVENTH FRAMEWORK PROGRAMME

Funded under Socio-economic Sciences & Humanities

The research leading to these results has received funding from the European Community's Seventh Framework Programme (FP7/2010-2.2-1) under grant agreement n° 266834

The Effect of FDI on Regional Inequality in the ENPs; Evidence from Israel

SEARCH Working Paper, 2/17.

MICHAEL BEENSTOCK[*], DANIEL FELSENSTEIN[+] AND ZIV RUBIN[+]

[*]Department of Economics msbin@mscc.huji.ac.il
[+]Department of Geography msdfels@mscc.huji.ac.il
Hebrew University of Jerusalem, Mount Scopus, Jerusalem 91905, ISRAEL

Abstract

FDI is an important source of capital for the ENPs. This paper investigates whether FDI polarizes regional inequality in host counties. In the absence of regional FDI data we propose a method for estimating the effects of FDI on regional inequality. An empirical application of this method is presented for Israel. We use time series data to show that regional capital stocks vary directly with the stock of national FDI and other variables, and that the sensitivity of regional capital stocks to FDI varies by region. We use regional panel data to show that regional wages vary directly with regional capital-labor ratios. In this way a link is established between FDI and regional wages via regional capital. Finally we decompose the factors driving regional wage inequality, as measured by the variance of regional wages. One of these factors is the polarizing effect of FDI on regional wages. Our results show capital stocks in the central (wealthier) regions of the country are more sensitive to FDI shocks. Also, the polarizing effect of FDI has increased absolutely during 1987-2010. However, it has decreased relatively; the contribution of FDI to regional wage inequality dercreased from 50 percent in 1987 to 30 percent in 2010. Policy implications of these findings are discussed.

KEY-WORDS: FDI, regional inequality, capital-labor ratio, panel data, ENP countries
JEL: C23, R12, R53

Acknowledgement: Thanks to Dai Xieer for assistance with the data analysis

1. Introduction

This paper attempts to answer the question: does FDI polarize inequality in host countries? In our context, this translates into estimating the effect of FDI from the EU15 countries on regional income disparities within ENP countries. Theory suggests that the motivation for FDI can either be 'horizontal' such as market access (Markusen 1984) or 'vertical' such as access to cheap factor inputs (Helpman 1984). However, while this may shed some light on factors promoting economic growth, it says very little about the effects of FDI on inequality. It is also not clear the extent to which this distinction is really relevant to the experience of developing countries in general and the majority of the ENP's in particular. The existence of overlapping domestic and target markets in developing economies and the importance of supra- national rather than national presence often serves to blur this distinction (Alessandrini and Resmini 2000).

Whereas the motivation for FDI has been widely discussed in the literature, the effect of FDI on regional inequality has received limited attention. Moreover, this issue is sometimes confounded with the related topic of MNE's as a vehicle for diffusing FDI, at both the national and regional levels. A major problem that has impeded empirical research on the polarizing effects of FDI is the absence of regional data on FDI. We have drawn attention to the fact regional capital stock data are not available for even the most advanced OECD countries (Beenstock, Ben Zeev and Felsenstein 2011). It therefore comes as no surprise that regional capital stock data by foreign ownership are not available. To circumvent this data problem we propose an empirical methodology for estimating the effects of FDI on regional inequality, which does not require data on regional FDI. This methodology uses data on national FDI which is universally available. We provide an empirical application of this methodology using data for Israel.

First, we model regional capital stocks as a function of national FDI, and other variables including regional incentives, regional population and human capital. We show that regional capital stocks vary in their sensitivity to national FDI shocks. Second, we estimate a regional wage model in which regional wages depend on capital-labor ratios and regional demographics. We show that given everything else, regional wages vary directly with capital-labor ratios. Since regional capital stocks may be more or less sensitive to FDI shocks, and regional wages vary directly with capital-labor ratios, a connection is established between FDI and regional wage inequality. Third, we use the regional wage model to decompose the factors driving regional wage inequality, as measured by the variance of regional wages. One of these factors is the effect of FDI on polarizing regional wages.

We proceed by reviewing the theoretical and empirical literature on the effect of FDI on host countries and regions in both developed and developing countries. We then chart patterns of FDI from the EU-15 in both the new EU eastern countries and the ENP's. The latter are divided into two blocs: Eastern ENP countries and Southern ENP countries, using FDI stocks during 1995-2010 and FDI flows during 1990-2010. The empirical methodology for estimating the polarizing effects of FDI on regional wage inequality is then presented. Using Israel as a prototype, we show how this methodology may be applied to estimate the polarizing effects of FDI on regional inequality. Given suitable regional data, we suggest that this approach can be replicated for other countries including ENP countries.

2. Literature Review

There is much theoretical ambiguity concerning the effects of FDI on human capital and relative wages in host countries. At the outset, it is important to differentiate the effect of FDI on developed and non- developed destination countries. In addition

it is useful to distinguish between national (domestic) and regional impacts. The tradition grounded in general equilibrium trade models with comparative advantage is highly sensitive to the initial equilibrium posited and to the parameter changes specified in such models. As such, these models can show both positive and negative effects associated with FDI (Markusen and Venables 1998). Endogenous growth models generally show more positive long run effects. Labor productivity grows because of imported knowledge and skill and TFP increases because of new technologies that accompany FDI. Much of this occurs through spillover effects (Blomstrom and Kokko 1998).

In terms of income effects, theory posits two different effects of inward FDI. On the one hand FDI can exacerbate income differentials by raising wages in recipient sectors. This is roughly in line with the dependency theory of FDI which views foreign control as an instrument for impoverishing host countries, creating employment opportunities for those with high opportunity costs, increasing capital intensity, raising unemployment in traditional sectors and consequently, exacerbating income differentials (Bornschier and Chase Dunn 1985). In similar vein, endowment-driven theoretical North- South models (eg Feenstra and Hanson 1997) also predict greater income inequality in host countries as FDI raises the skill premium.

Alternatively, FDI can be conceived as stimulating growth and employment that serve to narrow income gaps. This conforms with the modernist theory of FDI highlighting the diffusion of knowledge and technology associated with FDI that in the long run leads to a more equitable distribution of income (Figinia and Gorg 2011). FDI is considered a conduit for transferring new technologies and skills and upgrading local capacity. This is typically the case for FDI in developed host countries. An alternative view sees FDI activity as more skill intensive than local domestic activity thereby generating increased income inequality by increasing the demand for skilled labor (Taylor and Driffield 2005).

For developing countries, Figinia and Gorg (2011) develop a two stage non-linear model whereby FDI initially increases inequality between skilled and unskilled workers through the introduction of advanced technology. In the second stage, domestic capacity begins to imitate the production technologies introduced by FDI and gaps close. Thus FDI has a Williamson type inverted-U effect on equality in developing countries.

The empirical evidence with respect to the effect of FDI on domestic income inequality is as inconclusive as the theoretical models. For individual countries, FDI intensity is shown to be negatively related to income equality. This is true for both developed countries (Taylor and Driffield 2004) and developing countries (Feenstra and Hanson 1997). The latter suggest that in host countries where per capita incomes are lower than in origin countries, FDI is likely to be cost-driven and vertical. In countries where FDI host country per capita income is higher than incomes in the source country (eg Mexican investment in the US), FDI is likely to be horizontal and focused on market access.

Aggregate studies such as Tsai (1995), Choi (2006) and Chintrakan et al (2012) are equally ambivalent. Tsai's (1995) study of 33 developing countries does not find any support for a causal relationship between inward FDI and income inequality. Conversely, Choi (2006) using World Bank data for nearly 120 countries during 1993-2002 finds inward FDI stock related to a deterioration in the income distribution. This effect is more pronounced in larger, poorer and slower growth countries. Using panel data for all US states over a 24 year period, Chintrakan et al (2012) find that FDI reduces inequality over the long run but there is great heterogeneity across the individual states. For 21 out of 48 states there is a direct relationship between FDI and income inequality, suggesting a trade-off between productivity gains and widening social fissures.

In terms of FDI impacts on regional inequality, the geographic concentration of inward FDI has been observed for many developing countries. The most extreme example is probably China where 90% of inward FDI is clustered in coastal areas

accounting for 40% of population and 30% of area (Madariaga and Poncet 2007). In the case of India, Brazil and Indonesia, high levels of spatial concentration leading to a direct relationship between inward FDI and regional disparities, has also been noted (Sjoholm 1999, Daumal 2010). China has predictably been the focus of empirical attention relating to FDI and regional inequality (Zhang and Zhang 2003, Fu 2004). Much of this work shows that Chinese economic growth over the last two decades was fueled by FDI and accompanied by widening regional gaps. However, whether FDI inherently causes these disparities or whether they are a result of the uneven distribution of FDI, is unclear (Wei et al 2009).

Finally, spatial spillovers in the effects of FDI on regional inequality generally receive only indirect attention . Bode et al (2009) construct a Marshallian model of agglomeration with spillovers where firms' productivity is a function of the cluster of workers in proximate firms. The model is estimated using US data on FDI at the state level 1977-2003. The findings show FDI generated positive spillovers in contrast to the negative spillovers associated with domestic activity. Bloningen et al (2007) investigate spillovers associated with outbound US FDI to European countries. They find spatial interdependence between neighboring countries interpreting this as broad support for the export platform motivation for FDI. In the context of FDI in developing countries, Coughlin and Segev (2000) incorporate spatial effects in a study of US FDI and its impact on Chinese provinces. As in the above aggregate studies, they find that FDI in a given province has positive effects on FDI in proximate provinces. Similar findings have been reported for spillovers from Chinese cities where, as expected, these effects are strongest are for coastal cities rather than inland locations (Madariaga and Poncet 2007).

In summary, both theory and empirics offer mixed insights on the polarizing effects of FDI on developing host countries. It can be argued that FDI can both exacerbate income differentials and close income gaps. When a spatial dimension is added this ambiguity is further compounded. Regional inequalities can be conceived as the result of FDI location choices, and FDI spatial behavior can be interpreted as a result of regional disparities.

3. EU15 FDI in the ENP's

Alessandrini and Resmini (2000) describe the pattern of FDI by EU15 countries in the southern Mediterranean (SM) over the period 1980 -1998 and compare this with EU15 investment in Eastern and Central European countries (EE) over the same time frame. They note the trend of increasing marginalization of the SM in comparison with the growing attractiveness of EE. Data deficiencies notwithstanding, they find three EU core countries responsible for over 50% of EU investment in the SM area (Germany, France and Netherlands). They note a general weakening in EU interest in the region by the end of the 1990's in contrast to buoyant EU interest in Eastern Europe (FDI to EE doubled during the 1990's). They attribute this investment re-focusing to the completion of the single market and increased Eurocentricity over this period. A panel regression highlights those factors motivating FDI in EE and SM countries. For the former they find natural resource endowment, skilled labor and trade to be important in attracting FDI. For the latter, economic (not political) stability, (regional) market size and resource access are found to be significant.

We extend and update this analysis by charting trends in EU15 FDI (stocks and flows) from 1995-2009. We divide ENP destinations into and Eastern and Southern blocs (ENP-E and ENP-S)[1]. In addition, using data from OECD and Eurostat[2] we observe FDI flows to those Eastern Europe (EE) countries that joined the EU in its successive enlargements from 1995 to 2007[3].

[1] ENP-E is comprised of Belarus, Ukraine Moldova Georgia, Armenia Azerbaijan. ENP-S consists of Morocco, Algeria, Tunisia Libya, Egypt, Israel, Jordan, Lebanon Syria

Capital flows from the EU15 to ENP-E, ENP-S and EE are shown in Figures 1-3. The magnitude of total flows to all three areas in 1995 was $11.9b (in constant 2005 prices). The equivalent flow 15 years later was virtually the same at $11.5b. This static picture, however, hides inter-temporal fluctuations and a highly volatile pattern of investment that can be heavily skewed by contraction or expansion of a single foreign firm. In 1995 the flow of FDI to both ENP blocs was miniscule and narrow in origins. Traditionally EU investment was the single most important source of FDI in the Euro-Med (MEDA) countries that includes all the ENP-S bloc[4]. However, over the period 1995-2002 it declined as a share of total EU FDI, from 11 percent to 2 percent (Portelli 2004). The value of EU FDI in ENP-S in 1995 was only $0.36b, virtually all originating in France and the UK. The corresponding flow to ENP-E was $0.65b with 60% of this coming from Germany. At the level of the individual country, EU FDI in Morocco for example, accounted for 60 and 95 percent of all FDI over the period 1995-2002. However this was generally small scale and highly volatile investment in low value added industry such as textiles (Hemal 2004).

By 2009 EU15FDI in ENP-E was S4.04b with 75 percent of this coming from Germany, 20% from Sweden and small scale contraction of FDI originating in Belgium, Denmark, France and Luxembourg. In ENP-S 2009 was characterized by net disinvestment of $1.75b. Much of this could be ascribed to large scale contraction of a French cement plant investment in Egypt of nearly $5b that had injected double that sum in the previous year and despite UK and German FDI over $1b in 2009 (Fig 2). EU FDI to EE was larger and more diversified in terms of origin countries than in the ENPs. In 1995 FDI from the EU15 to EE was $10.9b and in 2009 it was $9.3b. Germany and Netherlands were the main sources of foreign investment in the earlier period. In the latter period, large scale German, French and Swedish FDI was offset by a massive Belgian contraction in EE.

Capital stocks of EU15 FDI are shown for the three destination blocs in Figures 4-6, which have positive time trends. In ENP-E the capital stock grew over this period from $1.2bEU to $62.4b. In ENP-S this growth was lower, rising from $2.5b to $48.7b. EU investment in EE increased from $19.5b in 1995 to $309b in 2009. This reflected the increasing attractiveness of EE as an investment location throughout this period given the enlargement of the EU. This probably came at the expense of investment in ENP-E. In both EU-E and EU-S much of the investment came from a few major origins. For example, over 70 percent of EU 15 investment in ENP-E in 2009 came from just three countries, Germany, UK and Holland. In the case of EU FDI in ENP-S two countries accounted for over 66 percent of all FDI (2009). With respect to the EE countries, in 1995 Germany and Austria accounted for over 65% of all investment while by 2009 France and Spain joined Germany and Austria as important origins. The magnitude of Austrian FDI in both ENP-E and EE most probably reflects the role of proximity and spillover in its foreign investment behavior.

---

[2] Eurostat: Stock (positions) come from: http://appsso.eurostat.ec.europa.eu/nui/show.do?dataset=bop_fdi_pos&lang=en
Data on flows comes from:
http://epp.eurostat.ec.europa.eu/NavTree_prod/AppLinkServices?pid=458_1209540_458_211810_211810&lang=en&appId=nui&appUrl=http%3A%2F%2Fappsso.eurostat.ec.europa.eu%2Fnui%2Fshow.do%3Fdataset%3Dbop_fdi_flows%26lang%3Den
Other FDI data is available from the OECDstat: http://stats.oecd.org/

[3] Bulgaria, Czech Republic, Estonia, Hungary Lithuania, Latvia Poland Romania Slovakia, Slovenia

[4] The Euro-Med Partnership is an EU free trade initiative incorporating twelve Mediterranean countries. In addition to the ENP-S group, it includes Malta, Cyprus and Turkey.

Fig 1: EU15 FDI in ENP-E

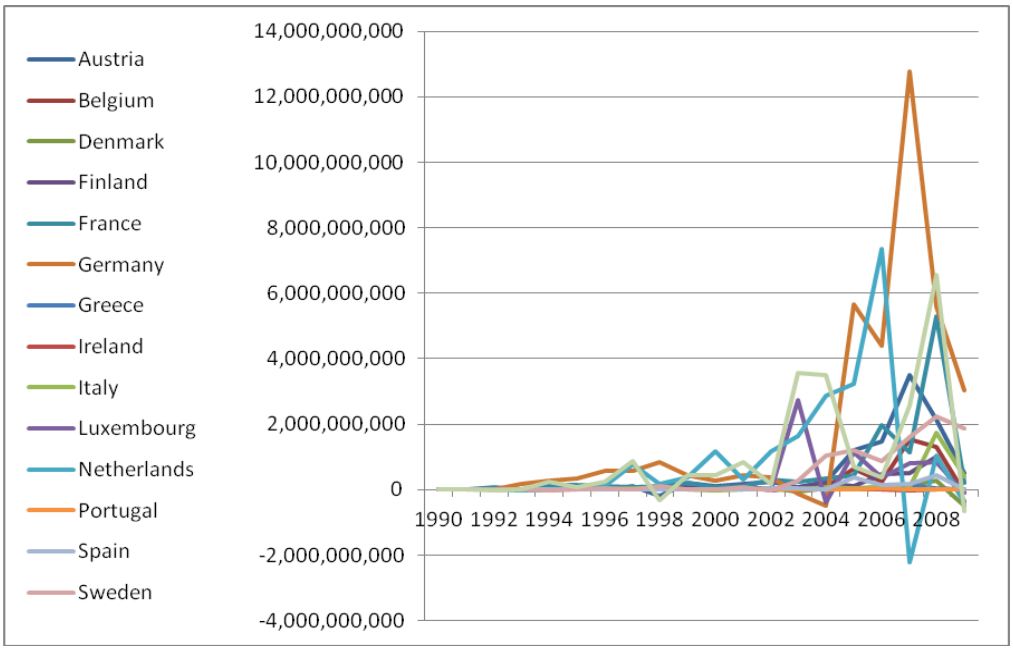


Fig 2: EU15 FDI in ENP-S

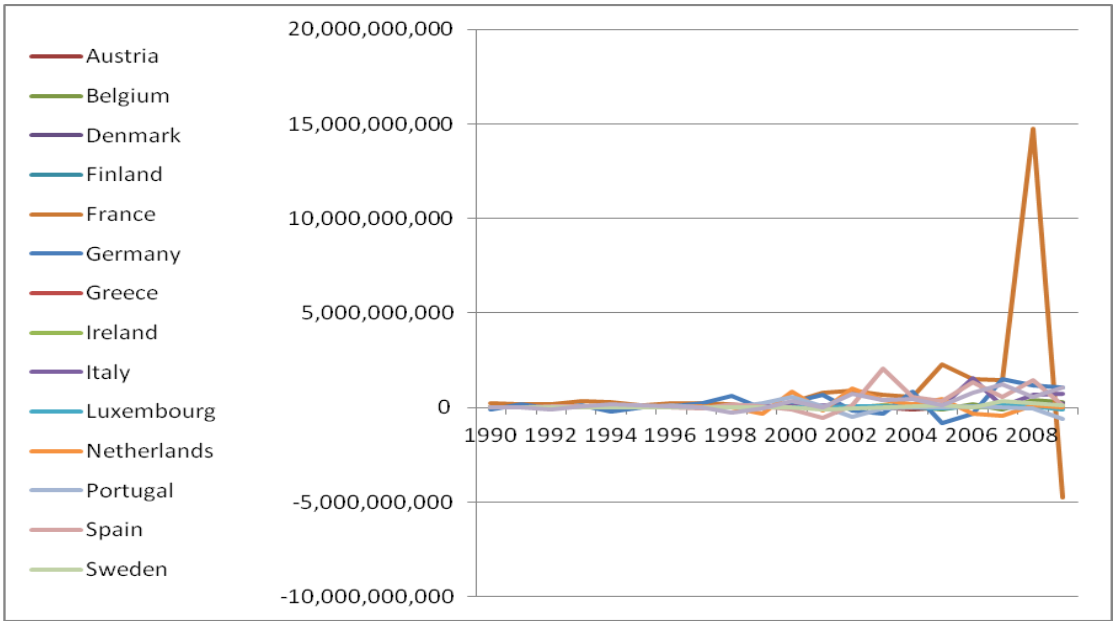


Fig 3: EU15 FDI in EE

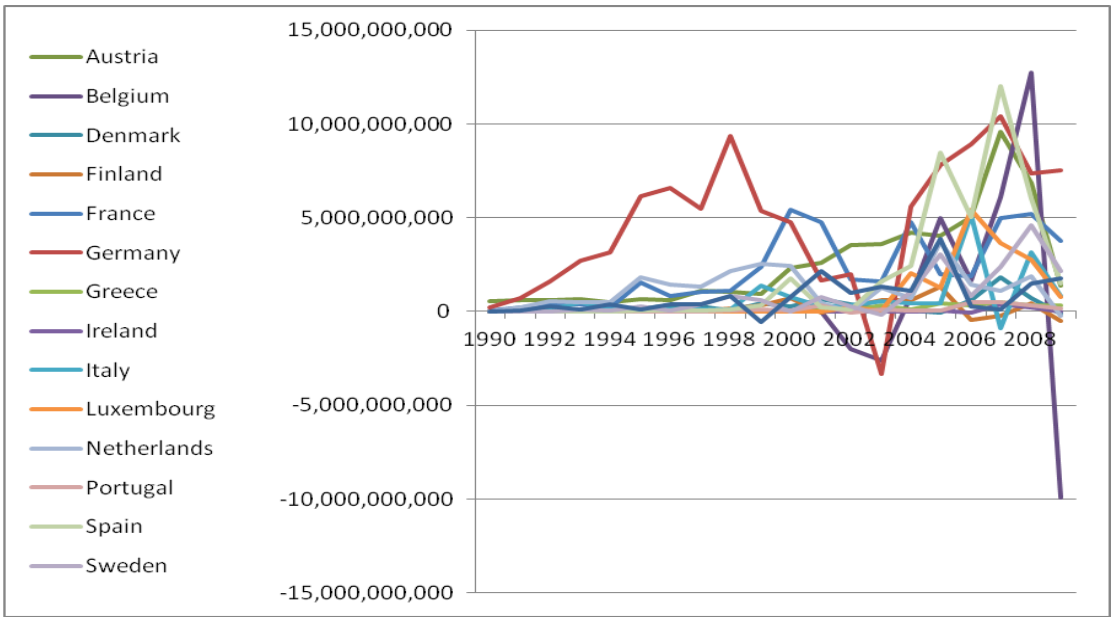


Fig 4: Stock of EU15 FDI in ENP-E

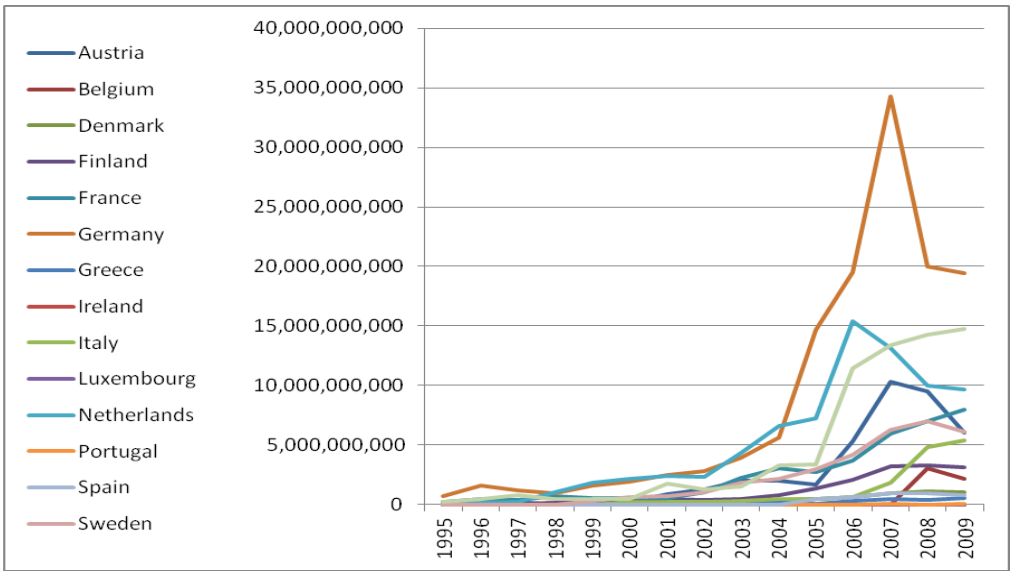


Fig 5: Stock of EU15 FDI in ENP-S

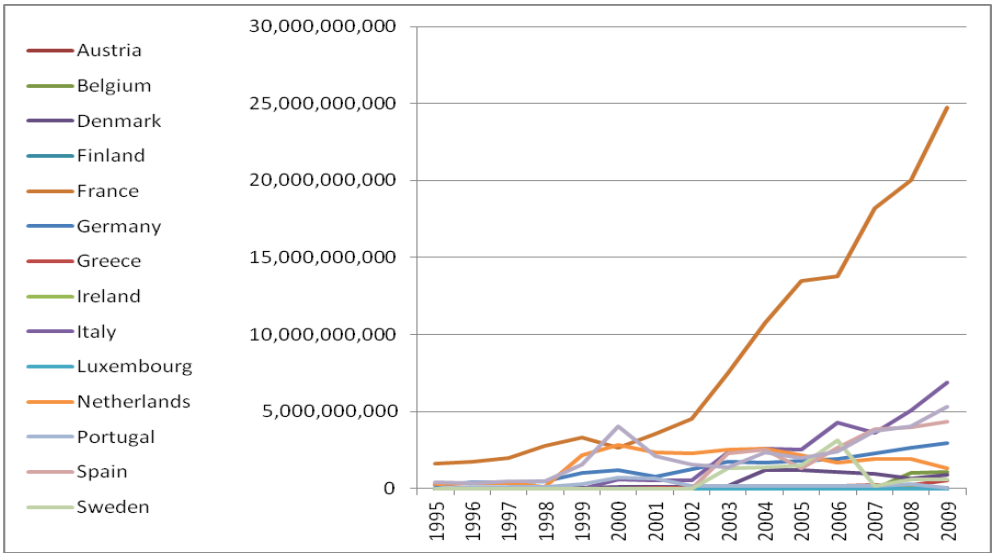
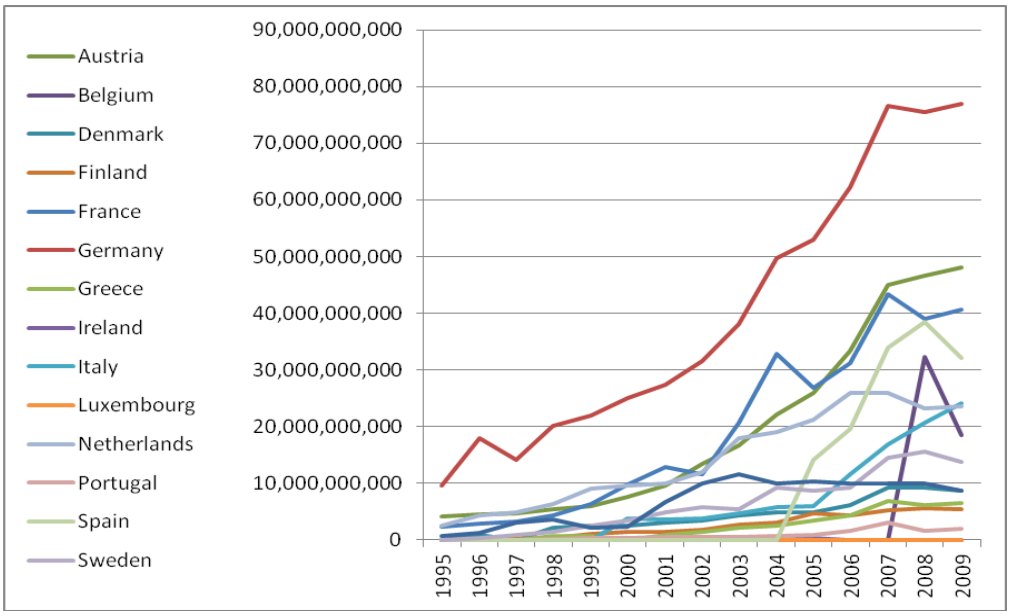


Fig 6: Stock of EU15 FDI in EE



4. Empirical Analysis:

In this section we present a method to investigate the potentially polarizing effects of FDI on regional wage inequality in the absence of regional data for FDI in host countries. However, the method requires data on regional capital stocks and wages. We begin by presenting the method. Subsequently, we illustrate the method with an empirical application for Israel.

4.1 Method

We regionalize a standard "Mincer model" for wages as follows. In the long-run, the real wage (w) in region i is assumed to equal labor productivity, which is hypothesized to vary directly with capital per worker (capital-labor ratio, k = K/L) and a vector of controls (X):

$$\ln w_{it} = \alpha_i + \beta \ln k_{it} + \gamma X_{it} + e_{it} \qquad (1)$$

where α denotes a regional specific effect. X includes "Mincer" variables such as average years of schooling, average age and its square, as well as controls for ethnicity, as defined below. X also includes agglomeration effects on labor productivity, as defined below. Finally, e in equation (1) denotes a "Mincer" residual, which captures unobserved regional heterogeneity in real wages.

This regionalized Mincer model may be used to decompose regional inequality in terms of its variance at time t:

$$\mathrm{var}(\ln w)_t = \mathrm{var}(\alpha) + \beta^2\,\mathrm{var}(\ln k)_t + \gamma^2\,\mathrm{var}(X)_t + 2\beta\gamma\,\mathrm{cov}(\ln k\, X)_t + \mathrm{var}(e) \qquad (2)$$

$$\mathrm{var}(\ln k)_t = \mathrm{var}(\ln K)_t + \mathrm{var}(\ln L)_t - 2\,\mathrm{cov}(\ln K \ln L)_t \qquad (3)$$

Equation (2) decomposes regional wage inequality into the contribution of the regional specific effects var(α), which do not vary over time, the contribution of inequality in regional capital-labor ratios var(lnk), which varies over time, and the contribution of inequality in the Mincer controls var(X), which also vary over time. Finally, regional wage inequality depends on

var(e), or unobserved heterogeneity, which will not vary over time unless it happens to be autoregregrssive conditionally heteroscedastic (ARCH). Covariance terms between e and α, and k and X are assumed to be zero, as they are in the method of estimation (see below). Equation (3) decomposes the variance of the capital-labor ratio into its capital and labor components. Equations (2) and (3) may be used to investigate the determinants of regional sigma divergence and convergence over time.

The auxiliary model for the regional capital stock is assumed to be:

$$\ln K_{it} = \phi_i + \theta_i \ln KFDI_t + \pi_i Z_{it} + v_{it} \qquad (4)$$

where KFDI denotes the stock of FDI and Z is a vector of controls hypothesized to determine the regional capital stock. If, for example, physical and human capital are complements or substitutes, Z will include average school years. It will also include regional investment incentives provided by the government. Notice that KFDI is defined nationally but not regionally. Also, the parameters in equation (4) vary across regions. A key parameter of interest is θ. If θ_i is larger, region i is more sensitive to FDI.

We may use equation (4) to decompose regional inequality in capital stocks since:

$$\mathrm{var}(\ln K)_t = (\ln KFDI_t)^2 \,\mathrm{var}(\theta) + \mathrm{var}(\pi_i Z_{it}) + \mathrm{var}(v) \qquad (5)$$

Since the variance of a product of two random variables is the product of their variances minus the square of the product of their expected values:

$$\mathrm{var}(\pi Z_t) = \mathrm{var}(\pi)\,\mathrm{var}(Z_t) - [E(\pi)E(Z_t) + \mathrm{cov}(\pi Z_t)]^2 \qquad (6)$$

This variance depends on time because the means and variances of Z vary over time. It varies directly with the variance of Z (regional inequality in Z) and inversely with the mean of Z.

The effect of FDI on regional wage inequality may now be calculated by differentiating the variance of log regional wages at time t with respect to the log of KFDI at time t:

$$\frac{\partial\,\mathrm{var}(\ln w_t)}{\partial \ln KFDI_t} = 2\beta^2 \ln KFDI_t \,\mathrm{var}(\theta)\left[1 - r_{\ln L \ln K}\, sd(\ln L)/sd(\ln K)\right] \qquad (7)$$

The first term in equation (7) refers to the direct polarizing effect of FDI under the assumption that the supply of labor is fixed, or perfectly inelastic. If the supply of labor is elastic, capital and labor will be positively correlated, as a result of which the increase in wages resulting from FDI will be smaller. This mitigating or indirect effect is captured by the second term in equation (7). If the supply of labor is perfectly elastic FDI has no effect on regional wages, there is complete mitigation, in which case equation (7) has a lower bound of zero. In general, however, equation (7) is positive; FDI polarizes regional wages.

Polarization varies directly, as expected, with β and the variance of θ, and it varies inversely with the elasticity of labor supply as reflected in the correlation between logK and logL. Equation (7) shows not only that FDI induces sigma divergence, the

elasticity of the variance of log wages with respect to the stock of FDI varies directly with KFDI. Therefore, the polarizing effect of FDI also increases with the level of the stock of FDI.

We have measured regional inequality using the variance. The proposed method may be cast in terms of other metrics. For example, the Gini counterparts to equations (2), (3) and (5) are:

$$G_{\ln w} = [G_\alpha^2 + \beta^2 G_{\ln k}^2 + \gamma^2 G_X^2 + G_u^2 + \beta\gamma G_{\ln k} G_X (\Gamma_{\ln k, X} + \Gamma_{X, \ln k})]^{\frac{1}{2}} \quad (8)$$

$$G_{\ln k} = [G_{\ln K}^2 + G_{\ln L}^2 - G_{\ln K} G_{\ln L} (\Gamma_{\ln K, \ln L} + \Gamma_{\ln L, \ln K})]^{\frac{1}{2}} \quad (9)$$

$$G_{\ln K_t} = [G_\phi^2 + (\ln KFDI_t)^2 G_\theta^2 + G_{\pi Z}^2 + G_v^2]^{\frac{1}{2}} \quad (10)$$

Where G_j denotes the regional Gini coefficient for variable j and Γ_{ji} is the regional Gini correlation coefficient between variable j and variable i. Equation (8) assumes that the Gini correlations between k and X and α and u are zero. Equation (10) assumes that the variables in equation (4) are independent, in which case their Gini correlations are zero.

The Gini counterpart to equation (7) is:

$$\frac{\partial G_{\ln w}}{\partial \ln KFDI_t} = \frac{\beta^2 \ln KFDI_t G_\theta^2 [1 - \frac{1}{2}(\Gamma_{\ln K, \ln L} + \Gamma_{\ln L, \ln K}) G_{\ln L} / G_{\ln K}]}{G_{\ln w} G_{\ln k}} \quad (11)$$

As in equation (7) Gini varies directly with KFDI. There is a direct effect and a mitigating effect. The polarizing effect of FDI on regional wage inequality varies directly with β, KFDI and inequality in θ, and it varies inversely with the elasticity of supply of labor as expressed by the Gini correlations between capital and labor. Equation (7) and (11) differ insofar as the polarizing effect of FDI does not depend on the variances of wages in the former but it varies inversely with the Gini coeffcient for wages in the latter. Below we use equation (7) rather than equation (11) since it is simpler.

4.2 Data Description

We create annual regional panel data during 1987-2010 for nine regions used by the Central Bureau of Statisics (CBS) for publishing house price data. These regions vary greatly in size but less so in population and roughly coincide with spatial housing markets (Map 1). The construction of the variables is described in the Data Appendix.

Figure 7 plots the panel data for regional wages (deflated by national CPI). Since wages grew over time these data cannot be stationary. Figure 8 uses the data in Figure 7 to chart regional wage inequality, as measured by the standard deviation of the logarithm of earnings. Inequality has increased over time, and especially since 2000. Sigma divergence clearly applies to wages. The shares of regional capital stocks in the national capital stock are plotted in Figure 9. The pattern that emerges is one of 'inverted convergence' (Beenstock, Ben Zeev and Felsenstein 2011); wealthier regions such as Tel Aviv and the Central region have increased their share and have closed the capital stock gap with respect to the those regions traditionally the recipients of public support (such as the North and Haifa regions). Figure 10 plots the panel data for capital per worker, which are also nonstationary. Note, however, that with the arrival of almost a million immigrants from the former USSR during the 1990s the capital-labor ratio stalled temporarily and real wage growth moderated (Figure 7).

Figures 11 and 12 chart the stock of FDI and the share of FDI in national GDP. The former shows clearly that FDI stock has a positive time trend, while the latter shows that FDI is volatile, peaking in 1999-2000 and 2006-8. These peaks were generated by a few flagship foreign direct investments such as multinational branch plant contruction, large scale mergers of Israeli firms with international conglomerates or celebrated high tech exit sell-outs.

Regional incentives are an important control variable when modeling the effect of FDI on regional capital stocks. These reflect government preferences for influencing industrial location behavior. The extent of government involvement in business location changes over time. Figure 13 shows the clear and consistent policy preference for investment in the North and South (and to a lesser extent in Jerusalem) over the other six regions. In Figure 14, the share of government incentives in regional capital stock is depicted. In certain areas the effect of public polcy is quite pronounced. Over the period 1995-2010, government incentives rose in the South from 5.9 percent of total capital stock to 9.0 percent. In the North the rise was from 3.8 to 5.8 percent over the corresponding period. Government share rose in all areas until the early 2000's and continued to rise in the Southern region until 2006. Subsequently, government rolled back regional incentives in all regions.

Fig 7: Regional Real Wages (ln), 1987-2010

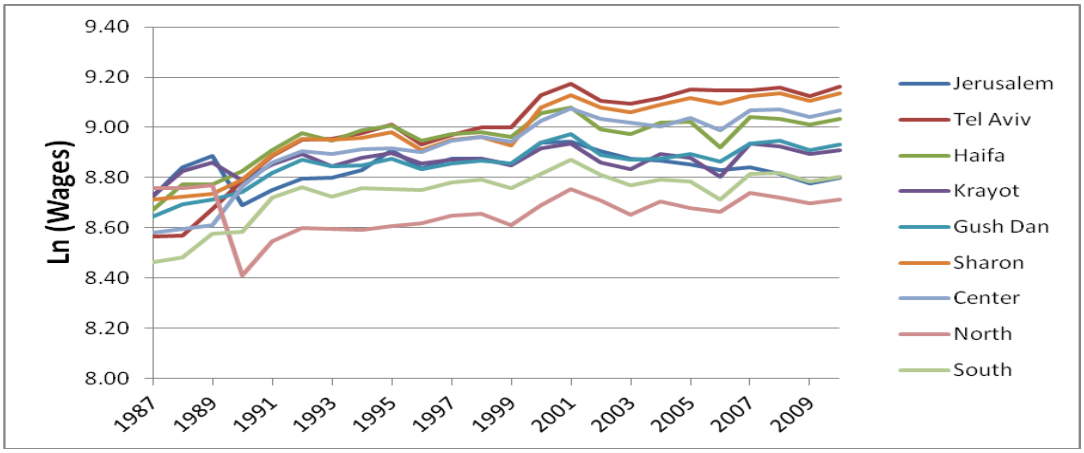


Fig 8: Regional Wage Inequality

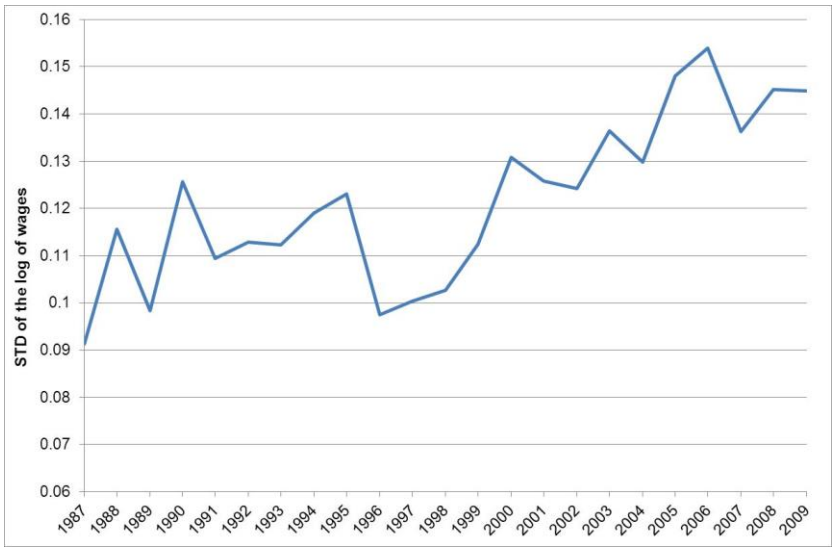


Fig 9: Shares of Regional Capital Stocks in National Capital Stock

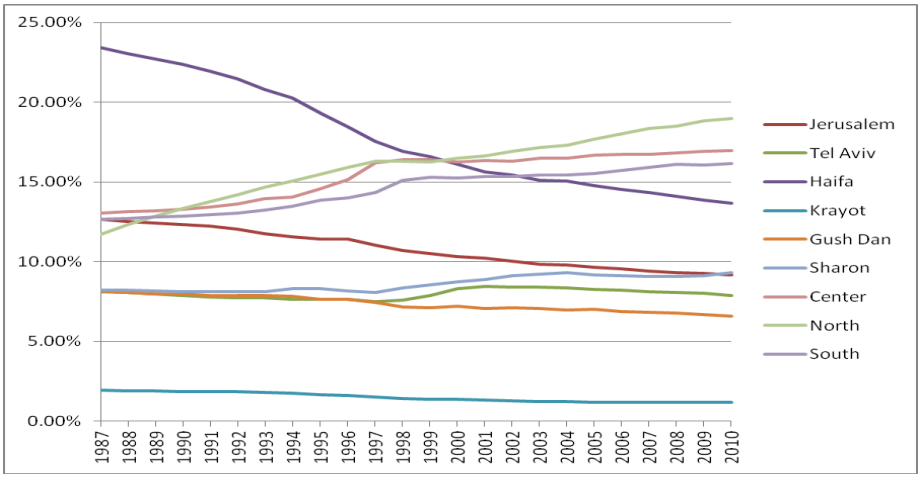


Fig 10: Capital per worker by region (ln), 1987-2010

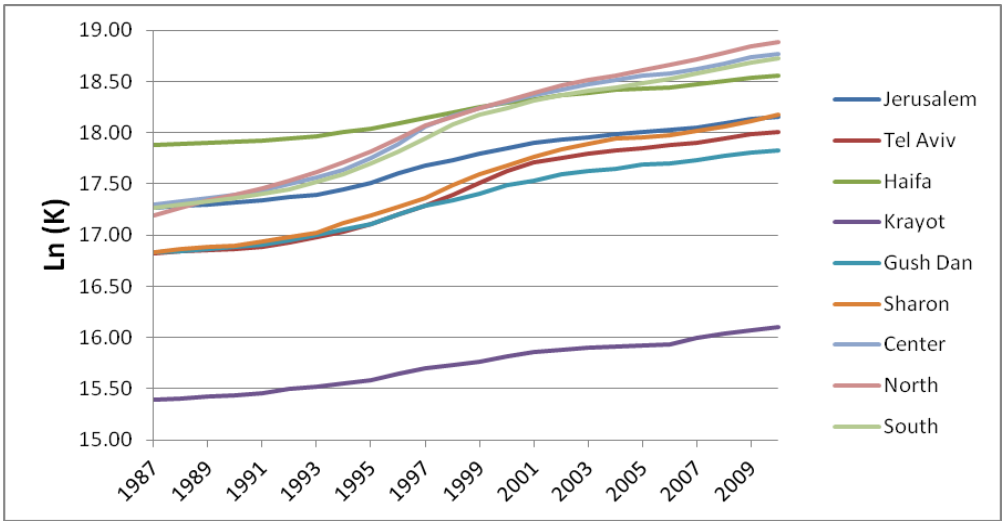


Fig 11 : Stock of Real FDI

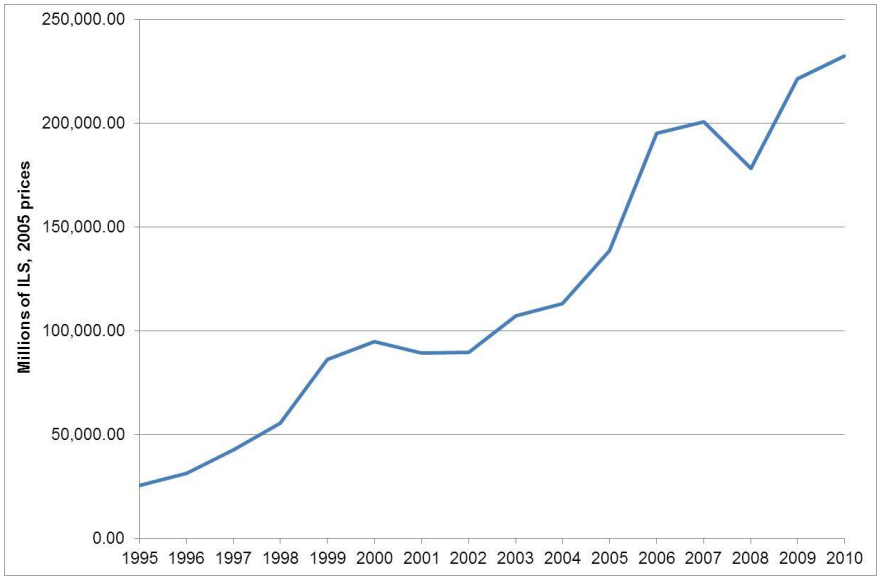

250,000.00
200,000.00
150,000.00
100,000.00
50,000.00
0.00
Millions of ILS, 2005 prices
1995 1996 1997 1998 1999 2000 2001 2002 2003 2004 2005 2006 2007 2008 2009 2010

Fig 12: FDI as % of GDP

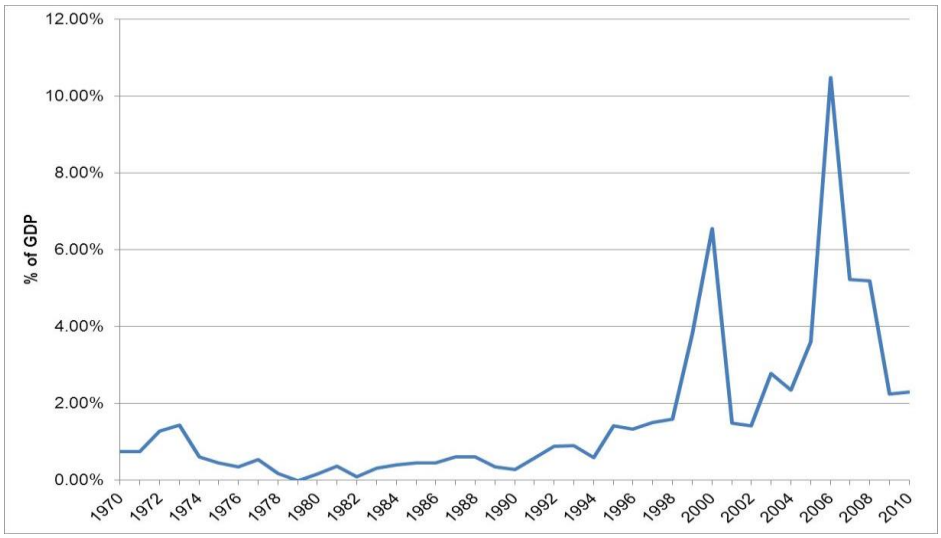


Fig 13: Stock of Government Incentives by Region: 1987-2012

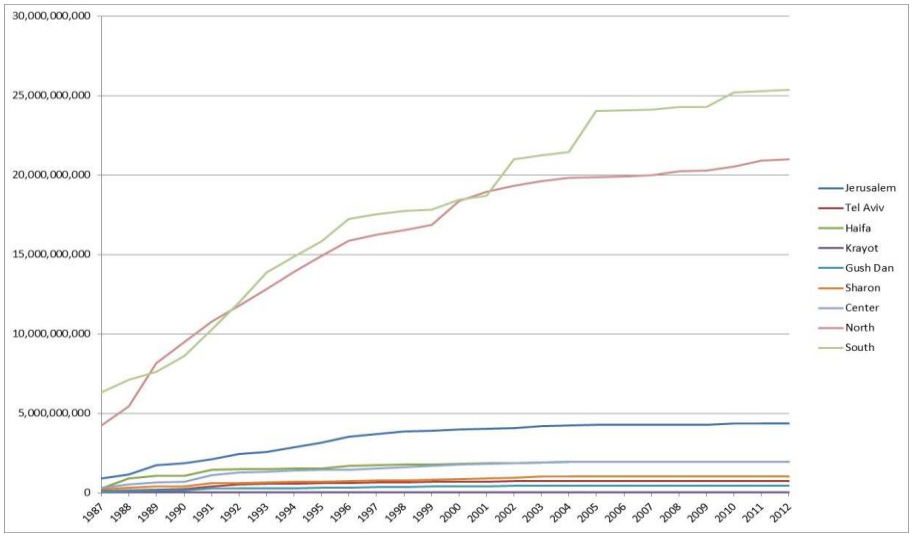
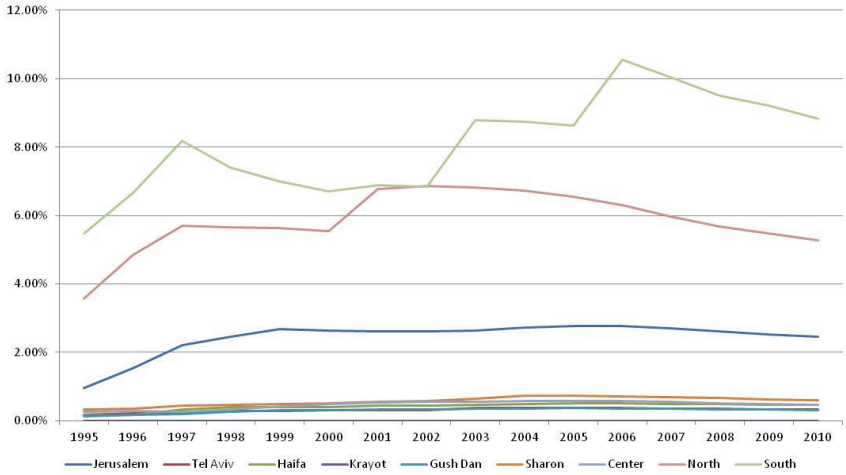


Fig 14: Government Incentives as a share of Regional Capital Stock



4.3 Methodology

Equation (1) is estimated using annual panel data for nine regions in Israel during 1987 – 2010. Since, as shown below, panel unit root tests indicate that the data are nonstationary, but are stationary in first differences, OLS or ML estimates of equation (1) might be spurious (Phillips and Moon). Such estimates are not spurious when estimates of e are stationary, in which case equation (1) is panel cointegrated. Specifically, we use the group ADF statistic (GADF) due to Pedroni (1999, 2004). Since the parameter estimates of cointegrating vectors have non-standard distributions, hypothesis tests concerning estimates of β and γ cannot be carried out using t- statistics, chi-square statistics and F statistics, which are all derived from the normal distribution. Instead such tests are carried out using GADF. For example, if GADF for an unrestricted model is cointegrated, but the restricted model is no longer cointegrated, the restrictions are rejected. On the other hand, if the restricted model is cointegrated, the restrictions cannot be rejected.

Parameter estimates of panel cointegrated models are superconsistent. If N is fixed (as it is in the present paper) these estimates are T-consistent if the variables in the model are driftless, and they are $T^{3/2}$ – consistent when there is drift (as in the present paper). This means that even though w and k might be jointly dependent, estimates of β are consistent. It also means that estimates of the residuals and fixed effects are asymptotically independent of k and X. Matters would have been quite different had the data been stationary. It also means that the spatial lag coefficient λ may be estimated without recourse to instrumental variables or maximum likelihood. In finite samples, however, OLS estimates may be biased (Banerjee et al 1986).

Equation (4) is estimated individually for each region. Since these time series data are difference stationary, cointegration tests are carried out for each region. Since there are only 24 time series observations for each region the power of these cointegration tests is not high. However, the joint power in nine independent cointegration tests is greater than in individual tests. Here too the parameter estimates are $T^{3/2}$ - consistent, and we draw comfort from the fact that the observation period covers almost a quarter of a century. We most probably learn more from 24 observations of annual data than from 48 observations of quarterly data. Here too estimates of the spatial lag coefficients μ_i do not require instrumental variables or ML for consistency.

4.4 Agglomeration in Labor Productivity

The region specific effects (α) capture unobserved differences in labor productivity. Productivity might be higher due to agglomeration or it might be higher for numerous other reasons. To investigate the effects of agglomeration we define A as:

$$A_{it} = (1-d)A_{it-1} + bk_{it-1} + a_{it} \qquad (8)$$

where d is the rate of depreciation on agglommerated knowledge, bk_{it-1} is new knowledge acquired from using capital, and a is an iid agglomeration shock. Since k is I(1) so must A be I(1). Given everything else, A is larger in regions where k was larger in the past. Therefore, even if $k_{it} = k_{jt}$ wages in j might exceed wages in i because $A_{jt} > A_{it}$. We therefore include $\ln A_{it}$ as one of the covariates in X in equation (1).

5. Results

5.1 Regional Wage Model

Figures 7 and 10 clearly show that the panel data for wages and capital-labor ratios are nonstationary. Panel unit root tests (Im ,Pesaran and Shin 2003) show that these key variables are stationary in first differences (Table 1), i.e. they are difference stationary. We therefore carry out tests of equation (1) using panel cointegration methods as described in section 4.3.

Table 1: Panel Unit Roots Tests for Difference Stationarity

Variable	IPS statistic
Lnw	-11.141
ln k	-6.185
Avg. school years	-11.737

Notes: IPS statistics for first differences of variables.

Table 2 presents four variants of the regional Mincer model in equation (1). Since the data are expressed in logarithms their coefficients can be interpreted as elasticities. Models 1 and 2 specify a full set of regressors and demographic controls as well as a spatial effect (in Model 2). Models 3 and 4 are much more parsimonious with model 3 specifying traditional demographic controls. The main difference between models 1 and 2 and models 3 and 4 arises from local capital agglomeration which is present in the former but not in the latter. Agglomeration is path dependent since it depends on evolution of the capital-labor ratio. Spatial spillover effects[5] are positive in model 2, implying that labor productivity in neighboring regions affect wages in the region under consideration. The group ADF statistics are very similar across all models, suggesting that there is not much to choose between them in terms of cointegration. Model 4 is most parsimonious and fits the data as well as model 2 which is the least parsimonious.

Table 2 shows that capital agglomeration lowers β, the coefficient of the capital-labor ratio, but fails to improve the cointegrative properties of the model. It also shows that the specification of demographic controls lowers estimates of β. We use the capital labor ratio from model 3 for estimation the FDI-regional inequality relationship.

[5] The coefficient on the spatial lagged dependent variable is estimated by OLS rather than maximum likelihood since OLS is super-consistent (see section 4.3). We use the following asymmetric spatial weight:

$$w_{nit} = \frac{1}{d_{ni}} \frac{Z_{it}}{Z_{nt} + Z_{it}}$$

where d_{ni} denotes the distance between regions n and i, and Z is a variable that captures scale effects.

Table 2: Real Wages: Estimates of Equation (1)

	1	2	3	4
Log capital-labor ratio	0.026	0.028	0.265	0.316
Log capital agglomeration	0.112	0.053		
Average school years	0.049	0.062		
Jews (percent)	0.184	0.216		
Average Age	0.028	0.032	0.302	
Average Age squared	-0.00062	-0.00069	-0.0035	
Ultra orthodox (percent)	-2.50	-3.12	-0.656	
Log Spatial wage		0.482		
Adjusted R^2	0.943	0.940	0.850	
GADF	-2.35	-2.20	-2.18	-2.25

Notes: Dependent variable – log real wages. Estimation by EGLS with SUR. GADF – z value for group ADF statistic of estimated residuals.

5.2 Regional Capital Stock Model

Unit root tests for key variables by each region, are presented in Table 3. With very few exceptions values fall short of the critical value (~3.0) meaning that virtually all variables in all regions have grown over time and therefore cannot be stationary. The ADF unit root statistic for FDI stock (tested on national data) is 0.298 and is also nonstationary.

Table 4 presents the estimated results of equation (4). A separate model is estimated for each region using annual data for 1987 - 2010. We use the minimized ADF statistic of the residuals to make the model selection in Table 4. We note that the estimates of θ turned out to be insensitive to alternative specifications. All regional capital stocks are sensitive to FDI, however some or more sensitive than others. The regions most sensitive are Tel Aviv, Sharon and Central and those least sensitive are the Krayot, Haifa and Northern regions. Other controls such as education, population and government incentives are specified in some models and not in others, depending on the results of the cointegration test.

Table 3: ADF Statistics

	J'lem	TA	Haifa	Krayot	Dan	Sharon	Center	North	South
lnK	-0.180	-0.204	0.462	0.596	-0.358	0.062	-0.756	-2.271	-0.490
lnPop	-0.497	-0.226	-1.894	-2.578	-0.283	-1.610	-2.005	-2.649	-3.100
lnEduc	-1.786	-0.419	-1.074	-0.403	-0.727	-0.780	-0.744	0.447	-1.312

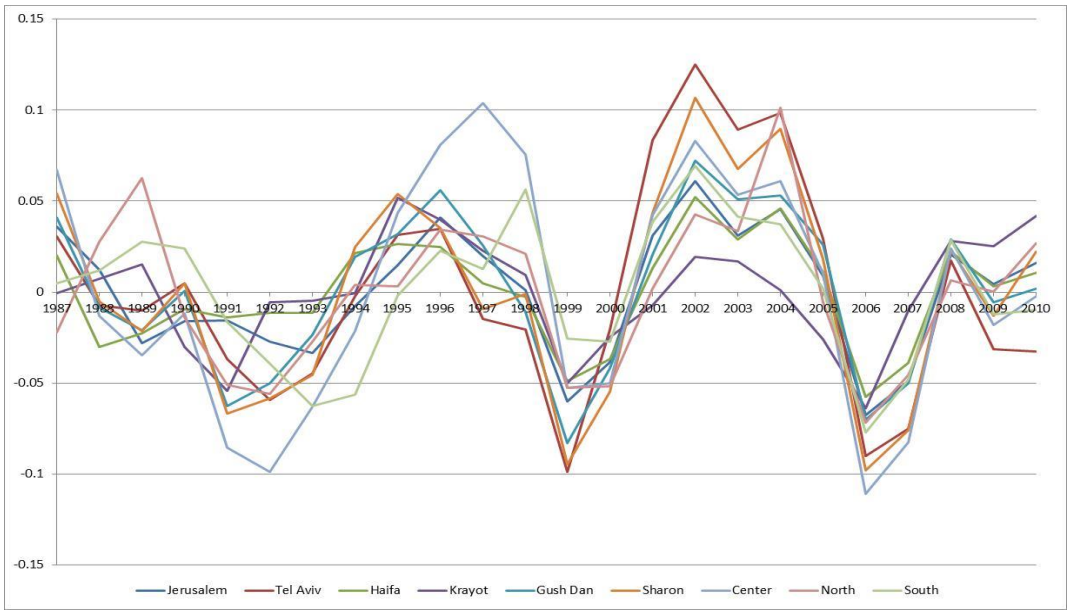
Table 4 : Determinants of Regional Capital Stock

	Jerusalem	Tel Aviv	Haifa	Krayot	Dan	Sharon	Center	North	South
Constant	11.365	10.871	14.563	9.316	10.814	8.337	7.877	4.482	8.252
lnKFDI	0.271	0.426	0.236	0.144	0.315	0.416	0.445	0.255	0.306
lnPOP				0.670				1.156	0.959
Educ				0.121					
lnKGI	0.154	0.092	0.049		0.156	0.225	0.251		
ADF	-4.269	-3.267	-3.887	-3.261	-3.673	-3.456	-3.378	-4.154	-3.364

The MacKinnon (1991) critical values for the cointegration test statistics are -4.11 at p = 0.05. Since most of the individual ADF statistics for the residuals exceed these critical values, not all the models reported in Table 3 are cointegrated. Jointly, however, they are cointegrated because their GADF z-statistic[6] is approximately -4.51, which clearly indicates that the models in Table 4 are jointly cointegrated. Figure 15 plots the residuals for the nine models in Table 4, and clearly indicates mean reverting tendencies. Figure 15 also indicates that fluctuations in regional capital stocks share a common cyclical component.

[6] Calculated as $\sqrt{N}\,\frac{\bar{\tau}-E(\tau)}{sd(\tau)}$ where N = 9 is the number of regions, τ-bar = 3.82 is the average of the ADF statistics, E(τ) and sd(τ) are the expected value and standard deviation of τ from Pedroni (1999) Table 2.

Fig 15: Residuals from Regional Capital Stock Models



5.3 The Effect of FDI on RegionalWage Inequality

Table 5 reports different regional wage sensitivities with respect to FDI. There is a clear center-periphery pattern; wages in the wealthier central regions (Center, Tel Aviv and Sharon) are more sensitive to FDI shocks than peripheral regions (North and South) and regions with an older industrial base (Haifa and Krayot). In Center the elasticity of wages with respect to national KFDI is 0.12, whereas this elasticity is only 0.03 in Krayot.

Table 5: Elasticities of Regional Wages with Respect to the Stock of FDI

	$\frac{\partial \ln w_{it}}{\partial \ln KFDI_t} = \beta\theta_i$
Jerusalem	0.072
Tel Aviv	0.113
Haifa	0.062
Krayot	0.030
Dan	0.083
Sharon	0.081
Center	0.118
North	0.068
South	0.081

The contribution of FDI to regional polarization over time is reported in Table 6 and plotted in Fig 16. The first column of Table 6 reports the direct polarizing effect of FDI on wage inequality. This is the first term in equation (7). Had θ been the same in all regions, regional polarization would have been zero. Since this is not the case, FDI induces regional inequality . The

second column is the total effect which includes the offset or mitigating effect. This is the second term in the RHS of equation (7). Table 6 shows that this offset is typically large; amounting to about two thirds of the direct effect. This results from the fact that elasticities of labor supply are relatively large. Column 2 shows that the absolute polarization effect has increased over time, raising the variance of the logarithm of regional wages by about 0.005 in the beginning of the period and by 0.006 at the end. The third column of Table 6 shows that the contribution of FDI to regional wage inequality decreased from about 50 percent at the beginning of the period to less than a third at the end

Fig 16: Effect of FDI on Regional Wage Inequality , 1987-2010

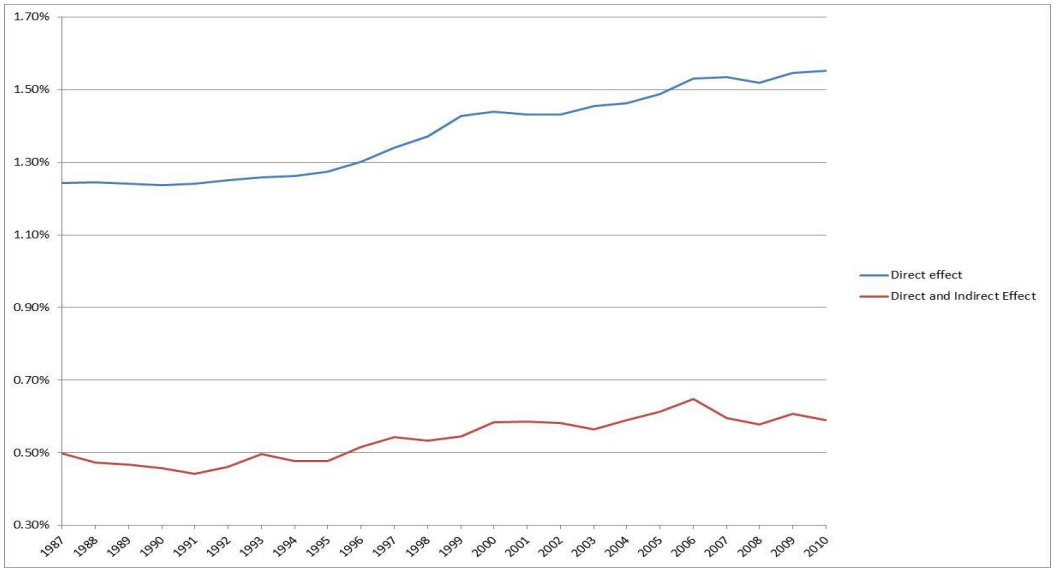


Table 6: Effect of FDI on Regional Wage Inequality

	$2\beta^2[\ln KFDI_t \operatorname{var}(\theta)]$ Direct Effect	$2\beta^2 \ln KFDI_t \operatorname{var}(\theta)\left[1-\frac{r_{\ln L \ln K} sd(\ln L)}{sd(\ln K)}\right]$ Direct and Indirect Effect	Contribution of FDI to wage inequality (%)
1987	0.012438526	0.004987151	60
1988	0.012444371	0.004729593	35
1989	0.012414319	0.00468123	48
1990	0.012379729	0.004573197	29
1991	0.012408454	0.004425186	37
1992	0.012501847	0.004616732	36
1993	0.012592772	0.004956223	39
1994	0.012626442	0.004764081	33
1995	0.012742209	0.004778043	31
1996	0.013013542	0.005149764	54
1997	0.013399068	0.00542332	54
1998	0.013720289	0.005341259	51
1999	0.014277144	0.00545794	43
2000	0.014394551	0.005830728	34
2001	0.014320613	0.005854475	37
2002	0.014325355	0.005815071	38
2003	0.014551529	0.005640713	30
2004	0.014619286	0.005897296	35
2005	0.014874799	0.00612449	28
2006	0.015303313	0.006482866	27
2007	0.015338443	0.005949402	32
2008	0.015189192	0.005776637	27
2009	0.01546068	0.006074199	30
2010	0.015523439	0.005896092	26

6. Conclusions

Research into the polarizing effects of FDI on regional wage inequality has been impeded by lack of data on regional FDI. In this paper we have proposed a methodology which exploits data on FDI at the national level, and which enables us to estimate the polarizing effects of FDI in the absence of data in FDI at the regional level. In a theoretical model we show that polarization varies directly with heterogeneity in the sensitivity of regional investment to national FDI and it varies inversely with the elasticity of regional labor supply. Polarization tends to zero as the labor supply elasticity tends to infinity.

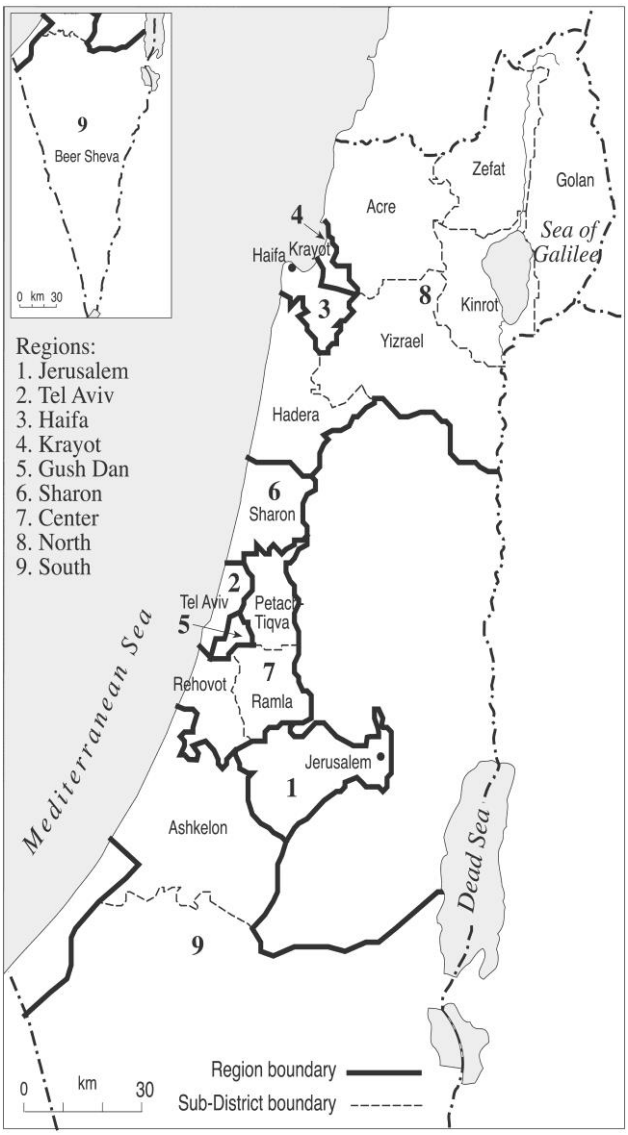
We use regional data for Israel to illustrate the proposed methodology. Empirically, we find substantial evidence of regional heterogeneity in investment to FDI shocks. The elasticities of regional capital stocks with respect to the national stock of FDI range between 0.14 and 0.45. Estimates of the polarizing effect of FDI on regional wage inequality turn out to be quite large. In the late 1980s FDI accounted for more than half the variance in regional wages. The polarizing effect of FDI increased by 20 percent over the subsequent 20 years. However, because regional wage inequality increased in Israel for other reasons, by 2010 the contribution of FDI to regional wage inequality had decreased to less than a third.

In terms of policy implications, we have shown that FDI increases regional capital stocks unequally, thereby exacerbating regional differences in labor productivity. Since regional wages vary directly with labor productivity a mechanism is established between FDI and regional wages. However, if regional labor supplies are elastic, the increase in wages induces employment, which mitigates the increase in wages, thereby offsetting the polarizing effect of FDI, partially and even totally. Since the elasticity of regional labor supply varies directly with internal migration, the polarizing effects of FDI on regional wage inequality may be mitigated by public policy which encourages internal migration.

Our results show that the polarizing effect of FDI on regional inequality may be large. The regional sensitivities to FDI shocks in Israel, reflect distinct core-periphery differences. In a small country such as Israel, this effect is likely to be smaller than in larger countries (like many ENP countries) where the physical distances between center and periphery are greater. In larger countries such as Morocco, Egypt or Ukraine, there may be entire regions not reached by FDI, which naturally would exacerbate the polarizing effect of FDI. Therefore, in other ENPs, which are much larger than Israel, the polarizing effect of FDI is likely to be even greater.

To offset the polarizing effects of FDI on regional wage inequality, public policy might reasonably consider targeting its regional investment policy on those regions which benefit less from FDI. For example, in Israel the periphery benefits less from FDI than the central regions. While the overall budget for regional investment incentives has been cut back, the share of the periphery in the regional development budget has increased.

Map 1: Israeli Regions



Data Appendix:

Regional aggregates are constructed using micro data for 1987-2010. Each variable is created from a different source and all data is aggregated to nine regions which form the basic spatial units of analysis (Map 1) . Variables and their sources are as follows:

Earnings: this refers to average regional earnings in shekels at constant prices. We use National Insurance Institute data on earnings by localities. This data is aggregated to the nine basic regions of the study.

Regional Capital Stock: this is constructed using a 'hybrid' methodology of perpetual inventory and proportional regional allocation described elsewhere (Beenstock, Ben Zeev and Felsenstein 2011). Source: residential and commercial property tax data published by the CBS for each locality[7].

Regional Capital Agglomeration: this is constructed as the cumulative depreciated effect of capital in the region (equation 8).We assume d=.05 and b=1.0. Unit root tests for this variable are not reported in Table 1. As the varibale is derived from regional capital stock it is assumed to have the same time series properties as the regional capital-labor.

FDI stock : this data is only available nationally. Source: CBS[8].

Regional Demographics: this refers to regional data on population, ultra-orthodox population, age and education levels (years schooling). The source for these demographic and human capital controls is the CBS Labor Force Survey (LFS) , micro data aggregated to 9 regions. The ultra-orthodox, age and years of schooling variables are regional averages.

Regional Incentives: this refers to the value (in constant 2005 shekels) of capital incentives (loans and grants) disbursed under Law for Encouragement of Capital Incentives to firms located in preferential areas. We use data on all loans and grants allocated to individual investment projects 1993-2012 and augment this data for the period 1967-1992 with data published in the annual reports of the Investment Center (the government agency charged with administering the policy). Data is then aggregated to 9 regions.

[7] CBS: *Local Authorities in Israel 2010* http://www1.cbs.gov.il/webpub/pub/text_page.html?publ=58&CYear=2010&CMonth=1
[8] CBS (Time Series-DataBank): *Balance Of Payments-International Investment Position-Assets and Liabilities*

References

Alessandrini S and Resmini L (2000) *European FDI in the Mediterranean region: A Comparison with CEE Experience,* MPRA Paper No. 26103, http://mpra.ub.uni-muenchen.de/26103/

Banerjee A, Hendry D.F. and Smith G.W. (1986) Exploring Equilibrium Relationships in Econometrics through Static Models: Some Monte Carlo Evidence, *Oxford Bulletin of Economics and Statistics*, 48, 253-277.

Barba Navaretti, G and Venables, A. (2004), *Multinational Firms in the World Economy*, Princeton University Press.

Beeenstock M, Ben Zeev, N. and Felsenstein, D. (2011), Capital Deepening and Regional Inequality: An Empirical Analysis, *Annals of Regional Science,* 47, 599-617

Blonigen, B.A., Davies, R.B., Waddell, G.R., Naughton, H.T. (2007), FDI in space: Spatial autoregressive relationships in foreign direct investment, *European Economic Review*, 51(5), 1303-1325.

Blomstrom M. and Kokko A (1998),Multinational Corporations and Spillovers, *Journal of Economic Surveys*, 12: 247-277.

Bode E, Nunnenkamp P and Waldkirch A (2009) *Spatial Effects of Foreign Direct Investment in US States,* Kiel Working Paper No. 1535

Bornschier, V. and Chase-Dunn, C. (1985). *Transnational Corporations and Underdevelopment,* New York: Praeger.

Choi C (2006) Does foreign direct investment affect domestic income inequality? *Applied Economics Letters*, 2006, 13, 811–814

Coughlin, C., Segev, E., 2000. Foreign direct investment in China: a spatial econometric study. *The World Economy,* 23 (1), 1–23.

Feenstra, R. C. and Hanson, G. H. (1997) Foreign direct investment and relative wages: evidence from Mexico's Maquiladoras, *Journal of International Economics*, 42, 371–93.

Figinia P and Gorg H (2011) Does Foreign Direct Investment Affect Wage Inequality? An Empirical Investigation ,*World Economy* 34(9), 1455-1475

Fu X. (2004) Limited linkages from growth engines and regional disparities in China, *Journal of Comparative Economics*, 32, 148-164.

Gopinath, M. and W. Chen (2003), Foreign Direct Investment and Wages: A Cross-Country
Analysis', *Journal of International Trade and Economic Development*, 12, 3, 285–309.

Herzer D and Nunnenkamp P (2011) *FDI and Income Inequality: Evidence from Europe,* Kiel Working Paper No. 1675 . Kiel Institute for the World Economy

Helpman, E., 1984. A simple theory of international trade with multinational corporations. *Journal of Political Economy* 92 (3), 451–471.

Hemal A (2004) Enhancing Neighborhood Policy through FDI, pp 68-72 in F. Attina and R.Rosa, (eds.*), European Neighborhood Policy: Political, Economic and Social Issue*, The Jean Monnet Centre "Euro-Med" Department of Political Studies, University of Catania.

Im K,, Pesaran M.H. and Shin Y. (2003) Testing for unit roots in heterogeneous panels, *Journal of Econometrics*, 115(1), 53-74.

Madariaga, N and Poncet S (2007), FDI in Chinese Cities: Spillovers and Impact on Growth, *World Economy* 30(5), 837-862.

Markusen, J.R., 1984. Multinationals, multi-plant economies, and the gains from trade, *Journal of* International Economics 16 (3–4), 205–226.

Mah, J. S. (2002) The impact of globalization on income distribution: the Korean experience, *Applied Economics Letters*, 9, 1007–9.

Pandejchintrakarn K, Herzer D and Nunnenkamp P (2012) FDI and Income Inequality : Evidence for a Panel of US States, *Economic Inquiry,* 50(3), 788–801

Phillips P.C.B. and Moon H. (1999) Linear regression limit theory for nonstationary panel data. *Econometrica*, 67(5),1011–1057

Portelli B. (2004) Foreign Direct Investment in the European Union's Mediterranean Neighbors: Past Trends and Future Potential in the MEDA Region, pp 73-85 in F. Attina and R.Rosa, (eds.*), European Neighborhood Policy: Political, Economic and Social Issue*, The Jean Monnet Centre "Euro-Med" Department of Political Studies, University of Catania.

Sjöholm, F (1999) Productivity in Indonesia: the Role of Regional Characteristics and Direct Foreign Investment, *Economic Development and Cultural Change* 47, 559-584.

Taylor, K., and Driffield N (2005) Wage Inequality and the Role of Multinationals: Evidence from UK Panel Data, *Labor Economics*, 12, 223–49.

Tsai, P.L. (1995) Foreign direct investment and income inequality: further evidence, *World Development*, 23 469–83

Wei, K., S. Yao and A. Liu (2009), Foreign Direct Investment and Regional Inequality in
China, *Review of Development Economics*, 13, 4, 778–91.

Zhang, X. and Zhang, K. H. (2003) How does globalization affect regional inequality within a developing country? Evidence from China, *Journal of Development Studies*, 39, 47–67.