Does Foreign Direct Investment Polarize Regional Inequality? A Methodological Proposal Applied to Israel

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Abstract

This paper investigates whether FDI polarizes regional inequality in host nations. We hypothesize that the link between FDI and regional inequality is mediated by the regional capital-labor ratio. In the absence of regional FDI data we propose a method for estimating the effects of FDI on regional inequality and present an empirical application 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 that capital 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 decreased from 21 percent in 1987 to 10 percent in 2010. Policy implications of these findings are discussed.

Keywords: *FDI, regional inequality, capital-labor ratio, panel data, wage decomposition*

JEL Classification: *C23, R12, R53*

1. Introduction

Foreign direct investment is an important source of capital for recipient countries.. However, little is known about the potentially polarizing effects of FDI on regional inequality. This is largely due to the difficulty in obtaining data on the regional distribution of FDI. We have drawn attention to the absence of regional capital stock data even for 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. This paper seeks to highlight the channels through which FDI affects income inequality in recipient areas and presents a methodology for estimating the effects of FDI on regional income inequality without recourse to regional FDI data.

The motivation for FDI has been widely discussed in the literature but 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. To estimate the polarizing effect of FDI on regional income inequality a two-stage approach is proposed. In the first stage, regional capital stocks are specified 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. In the second stage, a model is estimated in which regional wages depend on regional 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. Finally, 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.

The paper proceeds as follows. We begin by reviewing the theoretical and empirical literature on the effect of FDI on host countries and regions in both developed and developing countries. 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 empirically applied. Given suitable regional data, we suggest that this approach can be replicated for other countries.

2. Literature Review

There is considerable 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).

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. Coughlin and Segev (2000) incorporate spatial effects in a study of US FDI and its impact on Chinese provinces. 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). More recently, Monastirioitis and Borrell (2013) have used firm level data to investigate productivity and spillover effects of EU FDI in select EN countries, such as Morocco. They find that FDI adversely affects productivity in the sectors in which FDI is concentrated, but there is less evidence of spatial spillovers.

Absence of data on regional FDI stocks has forced investigators into using proxies for these missing data. For example, Haskel, Pereira and Slaughter (2008) proxy these data by the share of employment in foreign-owned plants, and Ascani and Gagliardi (2015) use regional (provincial) FDI flow data constructed by the Bank of Italy to study spillovers in innovation. We are unaware of any studies which use regional data on FDI stocks.

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. Empirical Analysis:

In this section we present a method for investigating the potentially polarizing effects of FDI on regional wage inequality in the absence of regional data for FDI in host countries. The method requires data on regional capital stocks and wages. We present the approach and subsequently, illustrate it with an empirical application for Israel.

3.1 Methodology

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 may also include agglomeration effects on labor productivity, as defined below. Finally, e 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 vary 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 autoregregressive 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 regional capital stocks 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 may 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. While we treat this as an empirical issue below, endogenous growth theory suggests that over the long run, θ will grow beccause of the imported knowledge and skills that accompany FDI (Blomstrom and Kokko 1998). Unlike neoclassical analysis which suggests that θ will only grow in the short run due to diminshing returns on the marginal product of capital, exogenous growth theory posits that factors such as technology and knowledge generate the necessary positive feedbacks to sustain the long terms growth of θ . The stock of FDI creates positive exzternalities through the mechanism of subcontracting, strategic alliances, technology licensing, imports of capital goods and migration (Romer 1990).Technology transfer and local spillovers prevent the decline in marginal productivity of capital suggested in neo-classical analysis. Therefore 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:

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

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, the penultimate term in equation (5) may be written as:

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

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*:

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

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 because internal migration enforces regional wage equality, 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 *r*, the correlation between logK and logL. Equation (7) shows FDI induces sigma divergence, and 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 stock of FDI but at a decreasing rate.

We have measured regional inequality using the variance. The proposed method may be cast in terms of other metrics. Following Yitzhaki and Schechtman (2013), the Gini counterparts to equations (2), (3) and (5) are:

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

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

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

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:

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

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.

3.2 Data

We create annual regional panel data during 1987-2010 for nine regions of Israel 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 1 plots the panel data for regional wages (deflated by national CPI). Since wages grew over time these data cannot be stationary. Figure 2 uses the data in Figure 1 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 3. The pattern that emerges is one of 'inverted convergence' (Beenstock et al 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 4 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 1).

Figures 5 and 6 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 investment incentives are a potentially important control variable when modeling the effect of FDI on regional capital stocks. These reflect government preferences for influencing industrial location. The extent of government involvement in business location changes over time. Figure 7 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 8, the share of government incentives in regional capital stock is depicted. In certain areas the effect of public policy 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 increase 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.

3.3 Econometric issues

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 1999). 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 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 (Beenstock and Felsenstein 2015). In finite samples, however, OLS estimates may be biased (Banerjee et al 1993), however this bias is mitigated due to diversification across the panel units.

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.

3.4 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 (A), 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).

4. Results

4.1 Regional Wage Model

Figures 1 and 4 clearly show that the panel data for wages and capital-labor ratios have positive time trends and are thus nonstationary. We use the heterogeneous panel unit root test (IPS) proposed by Im Pesaran and Shin (2003) that assumes independence between the panels in the data. This test is chosen as it allows for heterogeneity in the roots of each panel unit. Since some of the variables in equations (1) and (4) are non-stationary but are stationary in first differences, the equations are panel co-integrated if the residuals are stationary The IPS test shows 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 3.3.

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 based on Im, Pesaran and Shin (2003) for first differences of variables. Critical value of the IPS statistic when N=9 and T=23 is -1.99 ($p<0.05$).

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[1] are

[1] The coefficient on the spatial lagged dependent variable is estimated by OLS rather than maximum likelihood since OLS is super-consistent (see section 3.3). We use the following asymmetric spatial weight:

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

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.

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where d_{ni} denotes the distance between regions n and i, and Z is a variable (population) 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	0.875
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.

4.2 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 estimates 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

Note: Regional panel data, 1987-2010. KFDI is regional FDI capital stock, KGI is regional incentives stock, POP is regional population and Educ is regional average years schooling. All variables are defined in the Data Appendix.

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, matters are different because their GADF z-statistic[2] is approximately -4.51, which clearly indicates that the models in Table 4 are jointly cointegrated. Figure 9 plots the residuals for the nine models in Table 4, and clearly indicates mean reverting tendencies. Figure 9 also indicates that fluctuations in regional capital stocks share a common cyclical component.

[2] Calculated as $\sqrt{N}\,\dfrac{\bar{\tau}-E(\tau)}{sd(\tau)}$ 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.

4.3 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 10. 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 eighty five percent of the direct effect. This results from the fact that elasticities of labor supply are relatively large due to internal migration. Column 2 shows that the absolute polarization effect has increased over time, raising the variance of the logarithm of regional wages by about 0.0017 in the beginning of the period and by 0.0023 at the end. The third column of Table 6 shows that the contribution of FDI to regional wage inequality decreased from about 21 percent at the beginning of the period to about 10 percent at the end.

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	0.001780374	21.35
1988	0.012444371	0.001843310	13.80
1989	0.012414319	0.001847685	19.09
1990	0.012379729	0.001865233	11.81
1991	0.012408454	0.001907461	15.92
1992	0.012501847	0.001884009	14.79
1993	0.012592772	0.001824618	14.46
1994	0.012626442	0.001878572	13.24
1995	0.012742209	0.001902896	12.57
1996	0.013013542	0.001878911	19.79
1997	0.013399068	0.001905664	18.91
1998	0.013720289	0.002002021	18.97
1999	0.014277144	0.002107193	16.69
2000	0.014394551	0.002046174	11.96
2001	0.014320613	0.002022834	12.80
2002	0.014325355	0.002033382	13.17
2003	0.014551529	0.002129081	11.45
2004	0.014619286	0.002083965	12.37
2005	0.014874799	0.002090731	9.54
2006	0.015303313	0.002107490	8.89
2007	0.015338443	0.002243345	12.09
2008	0.015189192	0.002248963	10.68
2009	0.01546068	0.002242734	10.69
2010	0.015523439	0.002300284	10.24

5. 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 investigated the linkage between FDI and regional inequality as mediated through the channel of regional capital stocks. We have also tested a methodology which exploits data on FDI at the national level and enables us to estimate the polarizing effects of FDI in the absence of regional FDI data. 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.

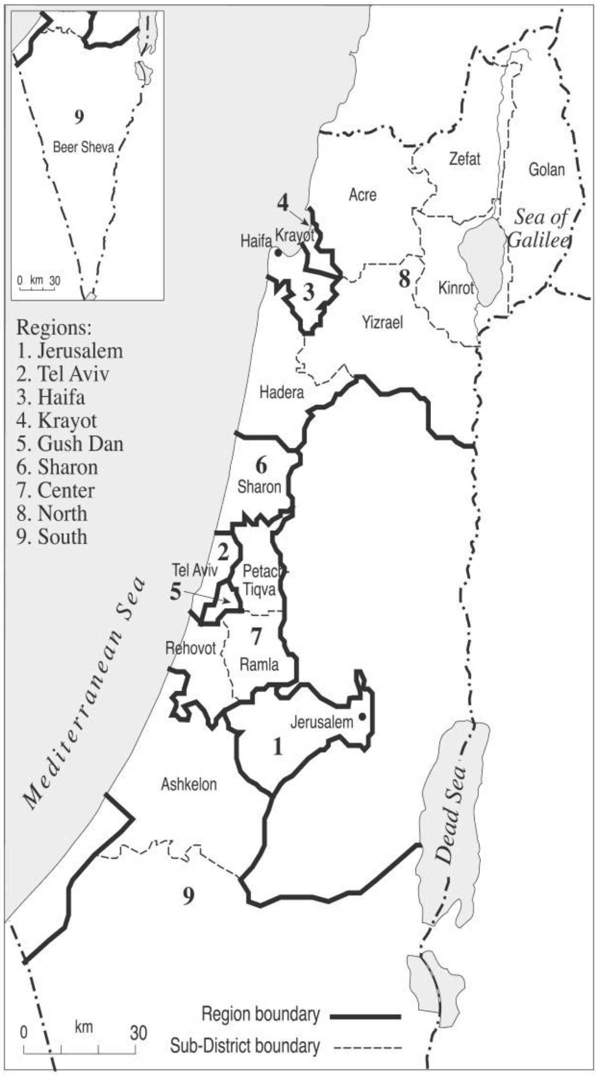
We use regional data for Israel to illustrate the proposed methodology. Empirically, we find substantial evidence of regional heterogeneity in investment to FDI shocks. The elasticities of regional capital stocks with respect to the national stock of FDI range between 0.14 and 0.45. Estimates of the polarizing effect of FDI on regional wage inequality turn out to be quite large. In the late 1980s FDI accounted for more than twenty percent of the variance in regional wages. The polarizing effect of FDI increased by 30 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 due to internal migration, 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 where the physical distances between center and periphery are greater. In larger FDI-destination countries such Egypt or Ukraine, there may be entire regions not reached by FDI, which naturally would exacerbate the polarizing effect of FDI. Therefore, in other countries, 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 consider targeting its regional investment policy on those regions which benefit less from FDI. For example, in Israel the peripheral regions benefit 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. In this way, the judicious use of regional policy may counter-balance some of the polarizing tendencies associated with FDI.

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: average regional earnings in shekels at constant prices. Source: National Insurance Institute data on earnings by localities.

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[3].

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.

FDI stock : available nationally. Source: CBS[4].

Regional Demographics: regional data on population, ultra-orthodox population, age and education levels (years schooling). Source: CBS Labor Force Survey (LFS) , micro data aggregated to 9 regions.

Regional Incentives: 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).

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[3] CBS: *Local Authorities in Israel 2010*
http://www1.cbs.gov.il/webpub/pub/text_page.html?publ=58&CYear=2010&CMonth=1

[4] CBS (Time Series-DataBank): *Balance Of Payments-International Investment Position-Assets and Liabilities*

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Fig 1: Regional Real Wages (ln), 1987-2010

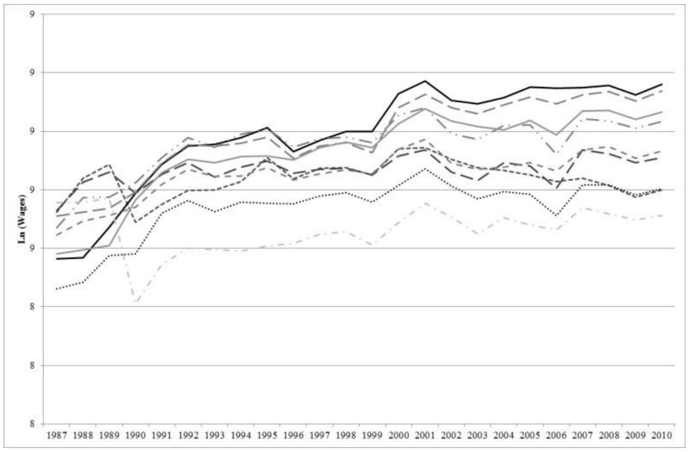


Fig 2: Regional Wage Inequality

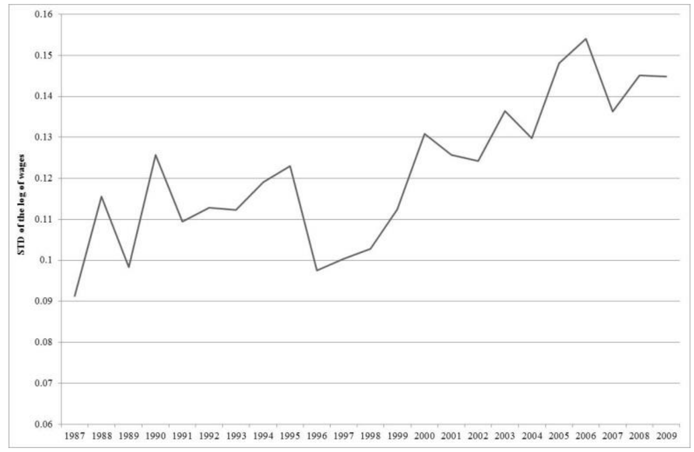


Fig 3: Shares of Regional Capital Stocks in National Capital Stock

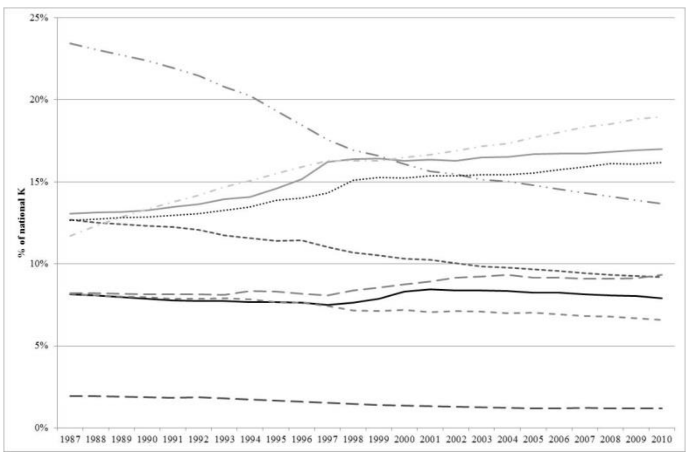
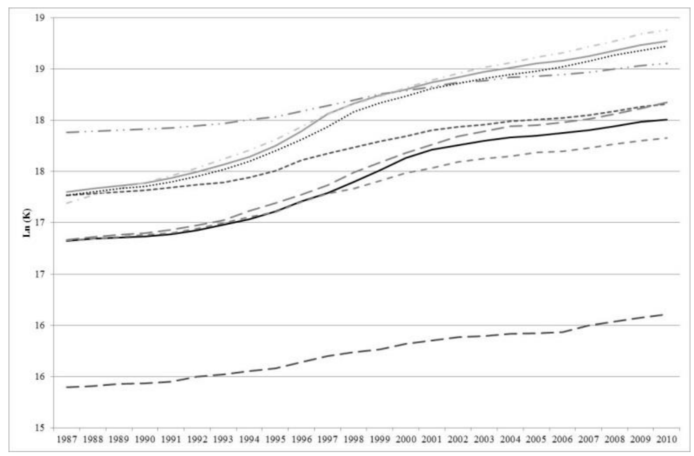


Fig 4: Capital per worker by region (ln), 1987-2010



Jerusalem Tel Aviv Haifa Krayot Dan Sharon Center North South

Fig 5 : Stock of Real FDI

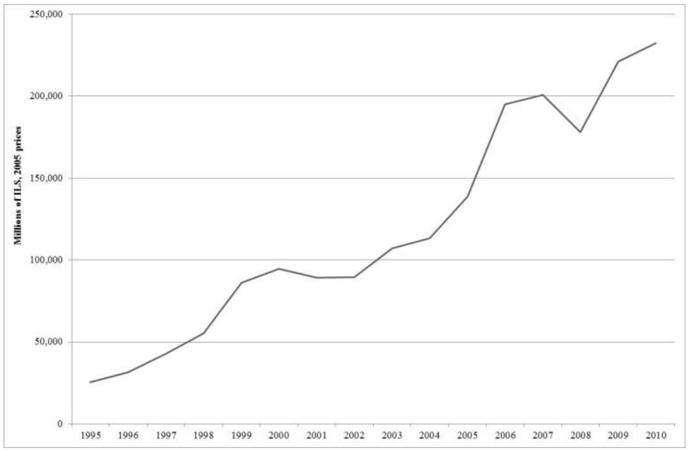


Fig 6: FDI as % of GDP

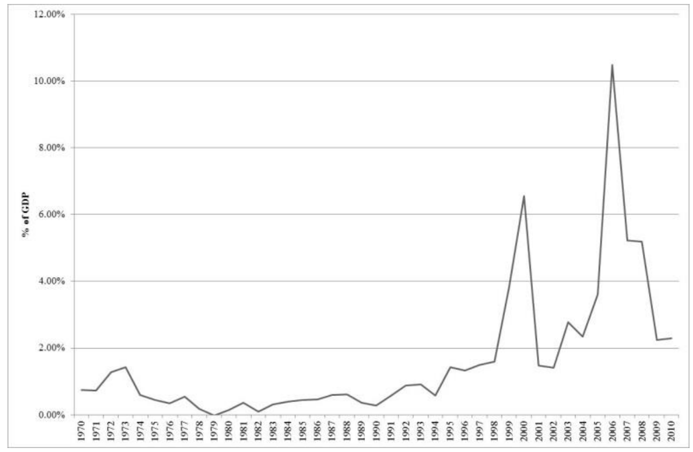


Fig 7: Stock of Government Incentives by Region: 1987-2012

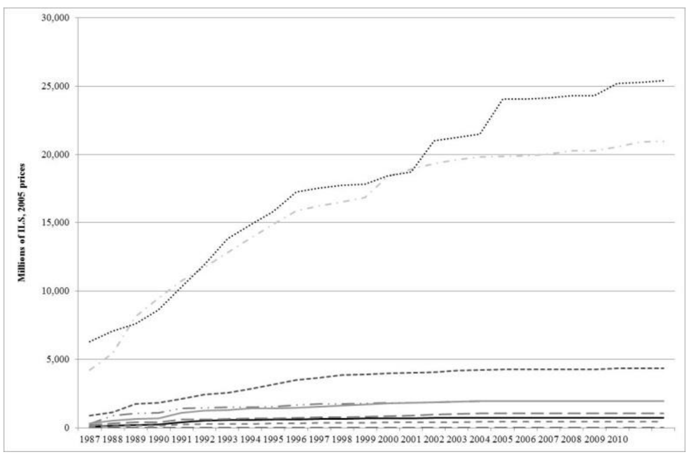
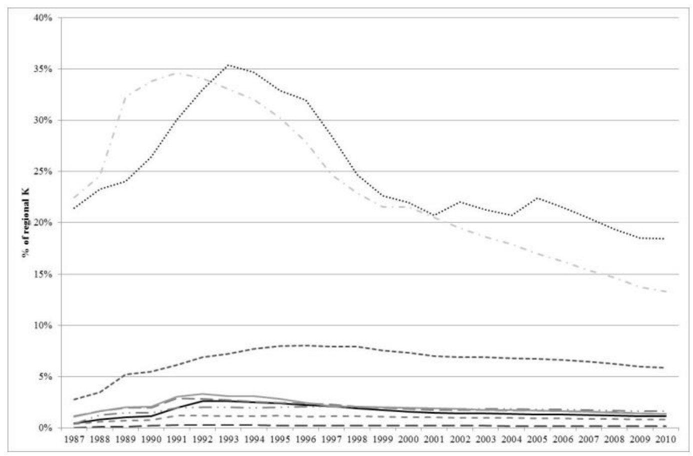


Fig 8: Government Incentives as a share of Regional Capital Stock



Jerusalem   Tel Aviv   Haifa   Krayot   Dan   Sharon   Center   North   South

Fig 9: Residuals from Regional Capital Stock Models

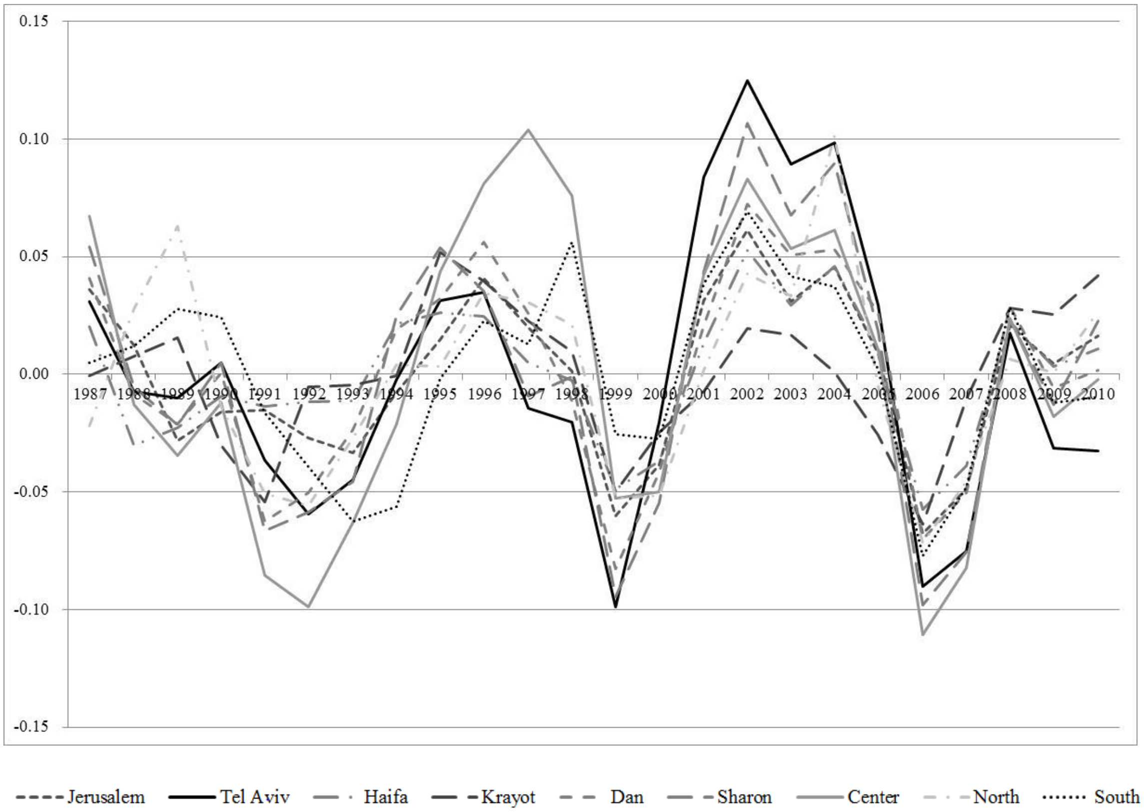


Fig 10: Effect of FDI on Regional Wage Inequality , 1987-2010