Lett Spat Resour Sci (2017) 10:361–374
https://doi.org/10.1007/s12076-017-0190-1

ORIGINAL PAPER

How internally mobile is capital?

Michael Beenstock[1]

Received: 11 January 2017 / Accepted: 22 May 2017 / Published online: 29 May 2017
© Springer-Verlag Berlin Heidelberg 2017

Abstract If capital is perfectly mobile between regions within countries, and regional TFPs share a common stochastic trend, the ratio of regional capital–labor ratios should remain constant over time. Spatial panel data on regional capital–labor ratios in Israel are used to test this hypothesis. Since the data are nonstationary, pairwise panel cointegration tests are applied. These tests are complicated by cross-section dependence between the spatial panel units. Although the null hypothesis of perfect capital mobility is overwhelmingly rejected, rejection of long-term perfect internal capital mobility is not overwhelming.

Keywords Internal capital mobility · Pairwise panel cointegration · Cross-section dependence

JEL Classification R12 · R32 · R53

1 Introduction

A key assumption in the theory of spatial general equilibrium (SGE) is that capital is perfectly mobile within countries, which ensures that rates of return to capital are equated across the economy (Roback 1982; Krugman 1991; Glaeser and Gottlieb 2009). Although internal labor mobility may be inhibited by personal attachments to places (Nocco 2009), such frictions do not extend to capital, which is expected to be perfectly mobile. Empirical investigation of PICM (hypothesis of perfect internal capital mobility) has been impeded by lack of data. There are no data on regional or

✉ Michael Beenstock
Michael.beenstock@mail.huji.ac.il

[1] Department of Economics, Hebrew University of Jerusalem, Mount Scopus, 91905 Jerusalem, Israel

spatial returns to capital, which may be used to test the hypothesis that spatial returns to capital are equated. Nor are there data on regional capital stocks that may be used to investigate PICM indirectly, by testing restrictions on the relationship between regional capital stocks. It is no doubt for these reasons that we are unable to cite previous empirical investigations of internal capital mobility.[1]

By contrast, there are numerous studies of international capital mobility, which are not reviewed here. Capital is expected to be more mobile internally than internationally since the latter involves exchange rate risk as well as country risk. Therefore, absence of international capital mobility does not necessarily presage the same for internal capital mobility.

We use capital stock data for Israel, constructed using the methodological proposal in Beenstock et al. (2011), to carry out indirect tests of PICM. Unfortunately, direct tests of PICM are not feasible since, in Israel as elsewhere, data on spatial returns to capital are not available. Specifically, annual capital stock data are generated for nine regions of Israel during 1987–2010. These data are used to test PICM according to the theory described in the next section, which predicts that the elasticities between pairs of regional capital–labor ratios should be unity. Since there are nine regions there are 36 pairs of regional capital ratios with which to test PICM. We test the joint hypothesis that these 36 pairwise elasticities are not significantly different from unity. Therefore, we do not require that each individual pairwise elasticity is unity.

Since the data for capital–labor ratios are nonstationary, OLS estimates of pairwise elasticities may be spuriously equal to unity. If, however, pairwise log differences between capital–labor ratios are stationary, these log capital–labor ratios are cointegrated with a unit elasticity, suggesting that capital is perfectly mobile between these pairs. To test the joint hypothesis of PICM we use critical values developed for panel cointegration tests, which do not require that all pairwise log capital–labor ratios be cointegrated. If the group augmented Dickey–Fuller statistic (GADF) is less than its critical value, the joint hypothesis that the 36 estimates of pairwise elasticities equal unity cannot be rejected, in which event the null hypothesis of PICM cannot be rejected. A methodological complication is that the 36 pairwise comparisons are unlikely to be independent because each region is involved in eight separate comparisons. The implications of this complication are discussed.

In summary, this is mainly an applied study that uses conventional econometric methods to test canonical theory regarding internal capital mobility.

2 Theory

Each region is assumed to produce homogeneous output using a common Cobb–Douglas production technology:

$$Q_j = A_j K_j^{\alpha} L_j^{1-\alpha} \tag{1}$$

[1] By contrast, there is an empirical literature on internal labor mobility. For example, Bernard et al. (2013) find that wages are not equated within the US. However, evidence of internal labor mobility does not necessarily presage the same for internal capital mobility.

where j labels regions, Q, K and L denote output, capital and labor, and A denotes total factor productivity (TFP). The cost of capital in region j is $r + \delta - s_j$ where r is the national rate of interest, δ is the rate of depreciation assumed to be the same across regions, and s denotes a subsidy to capital investment set by regional policy with respect to region j. It is assumed that all firms can raise finance at the rate r. Firms are assumed to equate the marginal product of capital (MPK) to the cost of capital, hence:

$$MPK_j = r + \delta - s_j = \alpha A_j k_j^{\alpha-1} \tag{2}$$

where $k = K/L$. PICM implies that Eq. (2) applies across all regions, in which case the capital–labor ratios (k) in regions i and j should be related:

$$\begin{aligned} A_j k_j^{\alpha-1} &= A_i k_i^{\alpha-1} + d_{ji} \\ d_{ji} &= (s_i - s_j)/\alpha \end{aligned} \tag{3}$$

where d_{ji} denotes the degree to which capital investment in region j is subsidized relative to region i, scaled by the inverse of the elasticity of output with respect to capital (α). According to Eq. (3) the gap between MPK_j and MPK_i increases with the gap between subsidies to capital in regions i and j. In the absence of regional investment subsidies ($d_{ji} = 0$) Eq. (3) implies that MPK is equated between regions i and j so that:

$$\ln k_j = \mu_{ji} \ln k_i + \frac{1}{1-\alpha} \ln\left(\frac{A_j}{A_i}\right) \tag{4}$$

where $\mu_{ji} = 1$ if capital is perfectly mobile between j and i. PICM predicts that the elasticities between pairs of k_j and k_i are unity.

3 Bringing theory to data

In our empirical study for Israel there are 9 regions and 23 annual observations (1987–2010). If $d = 0$ Eq. (4) may be written as:

$$\ln k_{jt} = \mu_{ji} \ln k_{it} + \frac{1}{1-\alpha_{ji}} \left(\ln A_{jt} - \ln A_{it}\right) + u_{jit} \tag{5}$$

where u denotes a residual. There are $N - 1$ pairwise comparisons for region j, so the total number of pairwise estimates of Eq. (5) is $1/2N(N - 1)$, which in our case is 36. Suppose that the panel data for lnk happen to be nonstationary (as we show below) but they are difference stationary, i.e. $\ln k_j$ and $\ln k_i$ are integrated time series of order 1, denoted by I(1). Estimation of Eq. (5) requires data on regional TFP, which in turn requires data on gross regional product. In the absence of such data, we assume that regional TFPs share common stochastic trends generated by $\ln A_{jt} = \ln A_t + e_{jt}$, where A_t refers to national TFP assumed to be difference stationary, and e denotes idiosyncratic shocks to regional TFP. The log differences in TFP in Eq. (5) are stationary because $\mathrm{lnA_{jt}} - \mathrm{lnA_{it}} = \mathrm{e_{jt}} - \mathrm{e_{it}}$. Since I(1) time series such as $\mathrm{lnk_i}$ are asymptotically independent of I(0) time series such as $\mathrm{e_j} - \mathrm{e_i}$, OLS estimates of μ and α are independent. Therefore, even if TFP is unobserved, it is possible to test hypotheses regarding μ which should be unity if capital is perfectly mobile. In this case the test simplifies to:

$$\ln k_{jt} = \psi_{ji} + \mu_{ji} \ln k_{it} + u_{jit} \tag{6}$$

which is estimated with T observations, and where ψ_{ji} equals the log difference in average TFP between regions j and i, and $\mathrm{u_{jit}}$ is the residual at time t from the ji'th pairwise time series regression. These residuals are unlikely to be independent for two reasons. First, $\mathrm{u_{ji}}$ and $\mathrm{u_{jk}}$ are directly related because they both involve region j. Second, $\mathrm{u_{ji}}$ and $\mathrm{u_{mn}}$ may be indirectly related through $\mathrm{u_{jm}}$ and $\mathrm{u_{in}}$. The econometric implications of this cross-section dependence are discussed below.

Unfortunately, the assumption that regional TFPs share common stochastic trends cannot be investigated empirically because regional data for capital stocks are required to measure regional TFP.[2] Whereas this assumption may be less plausible internationally since national TFP gaps may increase over time, it might be more justified intra-nationally, and especially in small countries such as Israel, within which productivity shocks are more likely to be common. As a robustness check, we assume that $\mathrm{lnA_{jt}} - \mathrm{lnA_{it}}$ is trend stationary[3] by specifying a deterministic time trend in Eq. (6) with coefficient τ_{ji}.

We distinguish between strong PICM where Eq. (4) holds in every time period, and weak PICM where it only holds in the long run. In either case the residuals (u) must be stationary so that $\mathrm{lnk_j}$ and $\mathrm{lnk_i}$ are cointegrated. In the former case, u must be serially independent for otherwise error correction would imply that PICM does not apply in the short-run. In the latter case error correction occurs between $\mathrm{lnk_i}$ and $\mathrm{lnk_j}$.

At first Eq. (6) is estimated imposing the restriction that $\mu_{ji} = 1$ for all regional pairs. If the residuals ($\mathrm{u_{ji}}$) are stationary $\mathrm{lnk_j}$ and $\mathrm{lnk_i}$ are cointegrated, in which case PICM applies to regions j and i. If all pairwise residuals are stationary PICM applies in all regions. If the residuals are not stationary PICM is rejected. In the latter event Eq. (6) is freely estimated, i.e. without imposing $\mu_{ji} = 1$. If the estimated residuals are stationary, $\mathrm{lnk_i}$ and $\mathrm{lnk_i}$ are cointegrated with $\mu_{ji} < 1$ as estimated, in which case capital is imperfectly mobile between j and i. In this event the average degree of imperfect capital mobility is summarized by μ-bar, which is the average of the estimates of μ_{ji}. If instead these residuals are nonstationary, the hypothesis of imperfect capital mobility is rejected too.

[2] Measures of regional TFP for e.g. US states are based on regional allocations of the national capital stock as in Garofalo and Yamarik (2002), or capital is ignored as in Caliendo et al. (2014). Bernard et al. (2013) used wage bill data to overcome unobserved differences in regional labor productivity. In the absence of "capital bill" data, we rely on Eq. (6) to account for unobserved differences in capital productivity.

[3] If TFP is trend stationary $\mathrm{lnA_{it}} = \mathrm{a_j} + \mathrm{b_j lnA_{jt-1}} + \mathrm{c_j t} + \mathrm{e_{it}}$ where $0 \leq \mathrm{b_i} < 1$ and $\mathrm{e_i}$ is stationary. Hence, $\mathrm{lnA_{jt}} - \mathrm{lnA_{it}}$ tends to $\mathrm{f_{ji}} + \tau_{ji}\mathrm{t} + \mathrm{h_{jit}}$ where $f_{ji} = \frac{a_j}{1-b_j} - \frac{a_i}{1-b_i}$, $\tau_{ji} = \frac{c_j}{1-b_j} - \frac{c_i}{1-b_i}$, and $h_{jit} = \sum_{n=0}^{\infty} b_j^n e_{jt-n} - \sum_{n=0}^{\infty} b_i^n e_{it-n}$.

Since capital may be perfectly or imperfectly mobile, and mobility may be strong or weak there are four possible types of corroboration:

i) If in Eq. (6) $\ln k_j$ and $\ln k_i$ are cointegrated with $\hat{\mu}_{ji} = 1$ and $\hat{\lambda}_{ji} = -1$ the hypothesis of strong PICM is corroborated between j and i.
ii) If $\hat{\mu}_{ji} = 1$ and $\hat{\lambda}_{ji} > -1$ the hypothesis of weak PICM is corroborated.
iii) If $\hat{\mu}_{ji} < 1$ PICM is rejected, but capital is imperfectly mobile. If $\hat{\lambda}_{ji} = -1$ imperfect capital mobility is strong.
iv) If $\hat{\lambda}_{ji} > -1$ imperfect capital mobility is weak.
v) If in Eq. (6) $\ln k_j$ and $\ln k_i$ are not cointegrated, capital is not even imperfectly mobile.

3.1 Cross-section dependence and critical values for panel cointegration tests

Each pair constitutes a separate test of PICM. A joint test of PICM involves the 1/2N(N − 1) pairwise comparisons. Let GADF denote the group average ADF statistic formed by the 1/2N(N − 1) ADF statistics using the residuals of Eq. (6). In panel data the critical value of GADF is normally distributed due to the central limit theorem:

$$z = \frac{\sqrt{\frac{1}{2}N(N-1)}\,(GADF - E(GADF))}{sd(GADF)} \sim N(0,1) \tag{7}$$

This formula is similar to the one proposed by Pedroni (1999) for panel cointegration tests. E(GADF) denotes the expected value of GADF under the null hypothesis of no cointegration, and its standard deviation is denoted by sd(GADF). We use the critical values of E(GADF) and sd(GADF) calculated by Pedroni (1999, 2004) for heterogeneous panel data, because each region has T observations for N − 1 comparisons. These critical values would be appropriate in the absence of cross-section dependence, i.e. where u_{ji} and u_{jk} are independent for all j, i and k. However, as noted, these residuals are unlikely to be independent within and between pairs. Since Pedroni assumed cross-section independence, his critical values are also likely to be too "liberal". We therefore calculate the residual cross-section correlation matrix (ρ_{ji}), which should be diagonal under the null hypothesis. We use the Breusch–Pagan LM test statistic for cross-section dependence (BP), which is expected to be zero under the null hypothesis of no cross-section dependence. BP is unbiased provided $N < T$ (Sarafides and Wansbeek 2012), which applies here.

If this null hypothesis is rejected, there are two main possibilities; cross-section dependence may be weak or strong (Chudik et al. 2011). In the former case the cross-section dependence is spatial and localized (Anselin 1988) whereas in the latter case the dependence is generic and is induced by common factors. In the former case shocks to u_{ji} dissipate across space, whereas in the latter case they do not. Pesaran (2015) suggested the following test for weak cross-section dependence:

$$CD = \sqrt{\frac{TN(N-1)}{2}}\,\bar{\hat{\rho}} \approx N(0,1) \tag{8}$$

where $\bar{\rho}$ is the average estimate ρ_{ji}. If this average is not significantly different from zero, the null hypothesis of weak cross-section dependence cannot be rejected. Unlike BP, CD depends on the sign of ρ. For example, if N = 3, T = 100 and the three values of ρ_{ji} are 0.4, 0.4 and −0.8, CD = 0 but BP = 96, in which the cross-section dependence is weak (spatial). If instead the values of ρ_{ji} are 0.4, 0.4 and 0.8, BP still equal 96 but CD = 32, in which case the cross-section dependence is strong. In general, both types of cross-section dependence may be present.

Cross-section dependence should lower the critical value for GADF below Pedroni's critical values. This means that if these pairs of lnk are not panel cointegrated using Pedroni's critical values, they cannot be panel cointegrated according to panel cointegration tests with cross-section dependence.

The spatial variant of Eq. (6), which allows for weak cross-section dependence is:

$$\begin{aligned} \ln k_{jt} &= \psi_{ji} + \mu_{ji} \ln k_{it} + \delta_{ji} \ln \widetilde{k}_{jt} + \pi_{ji} \ln \widetilde{k}_{it} + u_{jit} \\ \widetilde{k}_{jt} &= \sum_{n \neq j}^{N} w_{jn} k_{nt} \\ \widetilde{k}_{it} &= \sum_{n \neq i}^{N} w_{in} k_{nt} \end{aligned} \tag{9}$$

where w denote spatial weights row-summed to one, $\widetilde{k}_j$ denotes the spatial lagged dependent variable, and $\widetilde{k}_i$ denotes its spatial Durbin counterpart. If the data are stationary the spatial lag coefficients (δ and π) need to be estimated by maximum likelihood (Anselin 1988). However, if the data happen to be nonstationary, OLS estimates of these parameters are super-consistent (Beenstock and Felsenstein 2015) because $ln\tilde{k}$ is I(1) whereas u is I(0). The elasticity of k in region j with respect to k in region i is:

$$\mu_{ji} + \delta_{ji} \frac{\partial \ln \widetilde{k}_j}{\partial \ln k_i} + \pi_{ji} \frac{\partial \ln \widetilde{k}_i}{\partial \ln k_i} \tag{10}$$

which depends on the direct elasticity μ as well as the two spatial elasticities. Critical values for spatial panel cointegration have been calculated by Beenstock and Felsenstein (2017, chapter 7).

For strong cross-section dependence Pesaran (2006) has suggested the common correlated effects (CCE) estimator in which Eq. (6) is specified as:

$$\ln k_{jt} = \psi_{ji} + \mu_{ji} \ln k_{it} + \kappa_{ji} \ln \bar{k}_t + u_{jit} \tag{11}$$

where $\bar{k}$ denotes the average capital–labor ratio in period t and κ_{ji} denotes the loading of this common factor on the relation between k_j and k_i. Critical values for panel cointegration tests using CCE have been calculated by Banerjee and Carrion-I-Silvestre (2011, 2014).

A final variant of Eq. (6) takes account of regional rates of capital subsidy, which represent s in Eq. (3). If the subsidy to capital investment increases in region j relative to region i, k_j is expected to increase relative to k_i. We use data on cumulative regional investment grants to represent s:

$$\ln k_{jt} = \psi_{ji} + \mu_{ji} \ln k_{it} + \eta_{ji}(s_{jt} - s_{it}) + u_{jit} \tag{12}$$

where η is expected to be positive.

4 Data

Annual capital stock data for 9 regions of Israel (see Fig. 1) were constructed during 1987–2010 using the method proposed by Beenstock et al. (2011). The capital stock comprises plant and machinery. This method calculates capital investment in plant directly from regional data on building completions (square meters) in the business sector published by the Central Bureau of Statistics (CBS). The perpetual inventory method is used to construct regional capital stocks for plant in square meters, which are converted into constant price shekels using the implicit deflator for plant in the base year. Finally, CBS data for machinery capital at the national level (measured in constant price shekels) is allocated to the 9 regions according to the ratio of machinery to plant across the economy, i.e. plant and machinery are strict complements.[4] For example, if the value of plant in a region is \$100, and the ratio of machinery to plant across the economy is 1.3, the value of machinery in the region is imputed to be \$130 so that K = \$230. Whereas Garofalo and Yamarik (2002) allocate the national capital stock using regional shares in value added; we use direct measures of plant. In the absence of direct measures of machinery, we allocate national machinery according to regional shares in plant. This allocation rule is more direct than the one used by Garofalo and Yamarik. Data on employment (L) for these 9 regions were constructed from Labor Force Surveys (CBS). Since geographic disaggregation in these surveys is not continuously available prior to 1987, this determines the starting point for our investigation.

Capital–labor ratios (k) are plotted in Fig. 2 in 1000s of shekels at 2005 prices. Haifa region stands out as the most capital-intensive region of Israel because heavy industry has been concentrated in Haifa since Ottoman times and during the British Mandate (1921–1948). There are persistent and substantial differences between capital–labor ratios in the rest of the country. In 1987 the Dan region was the least capital intensive, but by 1996 it exchanged positions at the bottom of the distribution with Krayot. The Tel Aviv region, which in 1987 was in 4th position, temporarily moved up to 2nd position in 2003. On the whole,positions in the distribution appear to be quite stable. Following the wave of mass migration from the former USSR (1989–1995) capital–labor ratios naturally decreased especially in Haifa, which absorbed many immigrants. Subsequently, capital–labor ratios recovered, eventually surpassing what they were in the late 1980s. Because capital–labor ratios tend to increase over time, they cannot be stationary.

Since 1967 the Ministry of Trade and Industry (now the Ministry of Economics) has operated an Investment Center, which provides investment grants as part of its regional

[4] Hence $K_{it} = [1 + (K_{Mt}/K_{pt})]K_{pit}$ where K_M and K_P denote the national stocks of machinery and plant respectively. The allocation rule rescales regional capital stock data for plant. Consequently, the allocation rule has no effect on tests for $\mu = 1$ since K_i and K_j are rescaled identically.

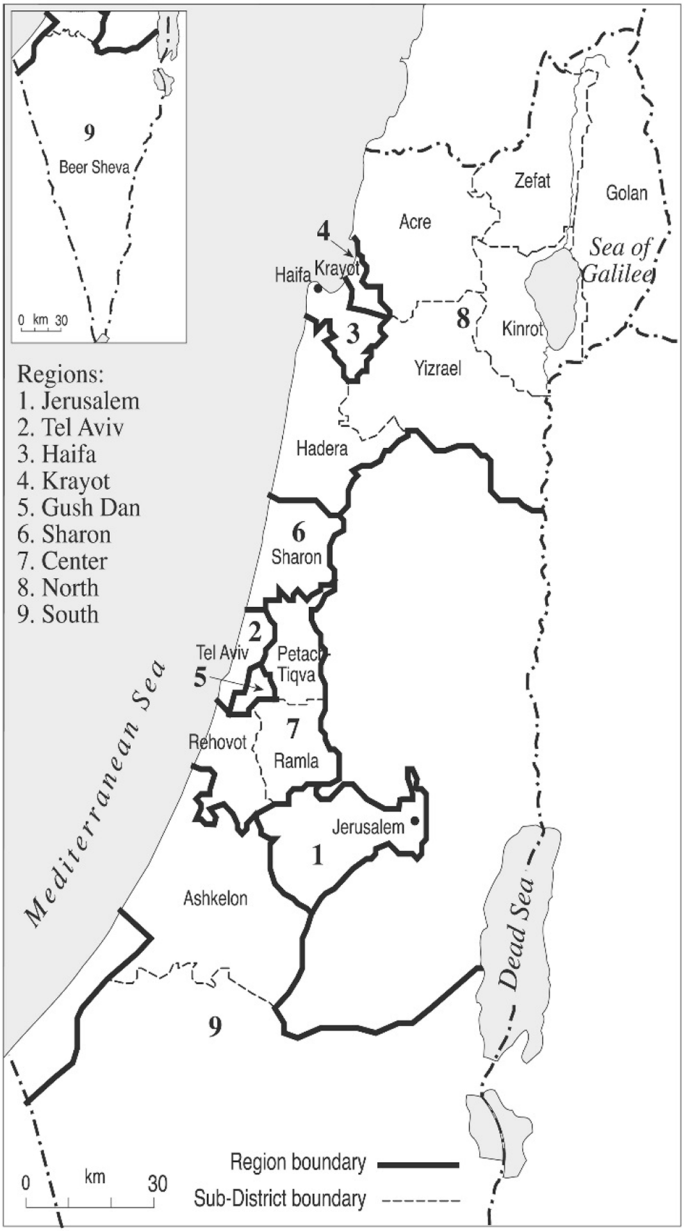

Fig. 1 Map

development policy. Businesses in designated regional development zones (A, B and C) are eligible to apply to the Investment Center for investment grants, which are awarded as percentages of the total investment, and which are highest in zone A and lowest in zone C. Priority is given to export businesses, and to industry rather than to services. The criteria have varied over time as have the zones eligible for regional

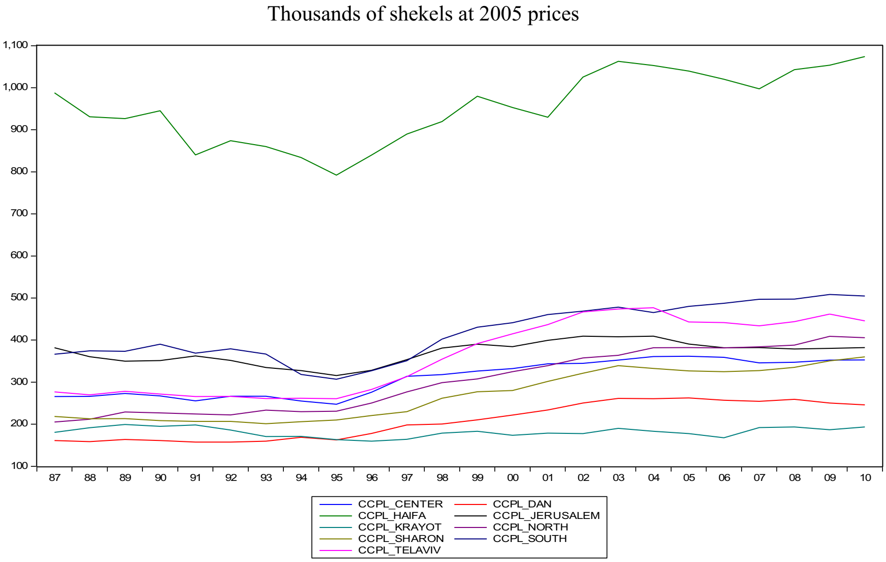

Fig. 2 Capital–labor ratios. Thousands of shekels at 2005 prices

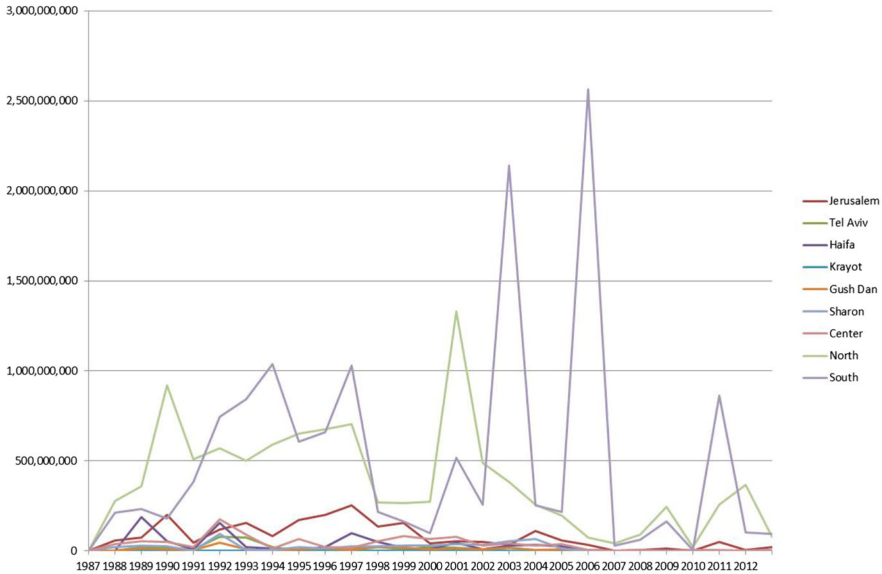

Fig. 3 Investment grants (shekels at 2005 prices)

development support. Figure 3 plots the allocation of investment grants at constant 2005 prices by the Investment Center in each of the 9 regions. The main beneficiaries have been the North and South, and since 2007 the budget of the Investment Center has been cut-back considerably. By contrast, the central regions (excluding Jerusalem) have received almost nothing. Figure 4 plots cumulative (since 1967) development grants received by the 9 regions.

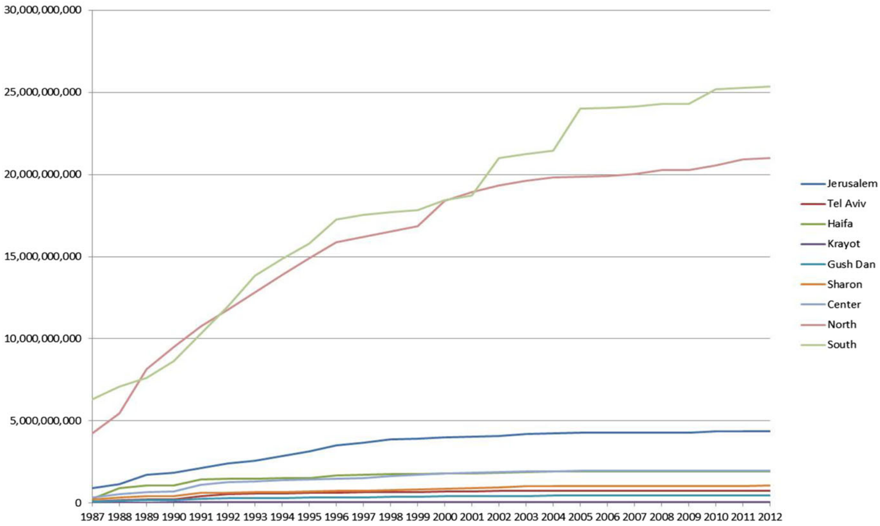

Fig. 4 Cumulative capital investment by the investment center (shekels at 2005 prices)

Table 1 Panel unit root tests

d	IPS		CIPS	
	0	1	0	1
lnk	1.07	−3.92	−1.51	−3.75
lns	−8.89	−8.15	−2.28	−3.64

Order of differencing denoted by d. IPS: Unit root test due to Im et al. (2003). CIPS: Unit root test due to Pesaran (2007) allowing for strong cross-section correlation

Panel unit roots tests are reported in Table 1 for the data in Figs. 2 and 3. The IPS statistics assumes that there is no cross-section dependence between the panel units, whereas the CIPS statistics assume that there is strong-cross section dependence. The IPS statistic confirms that lnk is difference stationary, as does the CIPS statistic. Matters are more complicated in the case of the data in Fig. 4 where both IPS and CIPS suggest that lns may be stationary. The problem is that for most regions lns maybe stationary, but in North, South and Jerusalem it is clearly nonstationary. These stationary components of lns cannot be cointegrated with lnk. Matters may obviously be different for North, South and Jerusalem.

5 Results

Results based on Eq. (6) are reported in Table 2. We begin by assuming that capital is perfectly mobile by imposing $\mu_{ji} = 1$, hence the elasticity between pairs of capital–labor ratios is unity in columns 2 and 3. Results are reported for two auxiliary hypotheses concerning unobserved TFP. If TFP is hypothesized to be difference stationary GADF is −0.69, which easily exceeds its critical value of −2.00. Hence, the

Table 2 Estimates of Equation (6)

Capital mobility	Perfect		Imperfect	
TFP	Difference stationary	Trend stationary	Difference stationary	Trend stationary
Elasticity	1	1	0.701	0.896
GADF	−0.69	−2.25	−2.31	−2.59
GADF*	−2.00	−2.64	−2.25	−2.73
BP	303.7	141.2	135.0	116.1
CD	7.6	13.0	8.7	23.3
$\bar{\rho}$	0.062	0.106	0.071	0.189

Elasticity: average estimate of μ (imposed at 1 in columns 2 and 3). GADF: group ADF statistic generated by residuals of Eq. (6). GADF*: critical value of GADF ($p = 0.05$) from Pedroni (1999, 2004). BP: Breusch-Pagan statistic for cross-section independence (critical value for chi-square at $p = 0.05$ c50). CD: test statistic for weak cross-section dependence (critical value 1.64 at $p = 0.05$)

Table 3 Estimates of Eq. (11): common correlated effects

Capital mobility	Perfect		Imperfect	
TFP	Difference stationary	Trend stationary	Difference stationary	Trend stationary
Elasticity	1	1	−0.077	−0.063
GADF	−2.44	−2.51	−2.84	−3.17
GADF*	−3.59	−4.28	−4.28	−4.30
BP	189.3	159.6	121.3	116.3
CD	3.63	7.01	2.48	4.16
$\bar{\rho}$	0.030	0.057	0.02	0.034

See notes to Table 2. GADF* is taken from Banerjee and Carrion-I-Silvestre (2011, 2014) Table 1

null hypothesis of PICM is clearly rejected. The BP statistic indicates the presence of cross-section dependence, implying that GADF* (which assumes cross-section independence) is too "liberal". Therefore, the rejection of PICM is even stronger than its nominal pvalue.

If, however, TFP is hypothesized to be trend stationary, matters are different. GADF decreases sharply from −0.69 to −2.25 but is still larger than its critical value. The BP statistic decreases sharply too, hence although cross-section dependence remains, it has weakened. The pvalue of GADF is approximately 0.12, hence the rejection of PICM is not overwhelming. On the other hand, cross-section dependence implies the effective p value is larger than 0.12.

In columns 4 and 5 of Table 2 μ_{ji} is estimated by OLS and the average elasticity is reported at 0.701 when TFP is hypothesized to be difference stationary, and 0.986 when TFP is hypothesized to be trend stationary. GADF is statistically significant in the former case and is almost statistically significant in the latter case. In both cases cross-section dependence is smaller than in columns 2 and 3. These results suggest that if capital is imperfectly mobile, the degree of mobility is quite high. However, the continued evidence of cross-section dependence undermines the reliability of these results.

Table 4 Estimates of Eq. (12)

Capital mobility	Perfect		Imperfect	
TFP	Difference stationary	Trend stationary	Difference stationary	Trend stationary
Elasticity	1	1	0.700	0.874
GADF	−1.37	−2.33	−2.49	−2.68
GADF*	−2.25	−2.73	−2.67	−3.08
BP	257.4	130.6	97.5	103.3
CD	8.22	11.31	8.87	14.06
$\bar{\rho}$	0.067	0.092	0.072	0.114

See notes to Table 2

5.1 Robustness checks

Since the CD statistics in Table 2 reject the null hypothesis of weak cross-section (spatial) dependence, there is little purpose in testing PICM with respect to the spatial model in Eq. (9). On the other hand, Table 3 reports estimates of Eq. (11) since CCE is motivated by the presence of strong cross-section dependence. The CCE estimate of GADF in column 2 is much smaller than its OLS counterpart in Table 2. Nevertheless, it is not statistically significant because the critical value for GADF for CCE is considerably smaller than its counterpart in Table 2. Therefore, the CCE test of PICM rejects the null hypothesis. Moreover, when μ is freely estimated as in columns 4 and 5, the estimate of μ turns out to be zero. Notice that although CD tends to be smaller, as expected, in Table 3 than in Table 2, it continues to be statistically significant. Therefore, CCE does an imperfect job in capturing strong cross-section dependence.

The results in Table 2 ignore the potential effect of regional subsidies to investment, which are taken into consideration in Eq. (12), presented in Table 4. The estimate of η (not reported) in Eq. (12) is positive as predicted by theory. However, the GADF statistics, which tend to be smaller than their counterparts in Table 2, fall short of their critical values. However, the pvalue of GADF in column 4 is approximately 0.1.

6 Conclusions

In the absence of data on regional returns to capital, data for regional capital–labor ratios in Israel are used to carry out indirect tests of the hypotheses of perfect and imperfect internal capital mobility. This issue is important because spatial general equilibrium theory assumes that capital is perfectly mobile within countries. Indeed, this is the first time that tests of internal capital mobility have been undertaken. If capital is perfectly mobile the elasticity between pairs of capital–labor ratios is expected to be 1 under Cobb–Douglas technologies.

Since the panel data in the study are nonstationary, these tests are carried out using panel cointegration methods where each panel consists of pairs of regions. Annual data for almost a quarter of a century are used for 9 regions in Israel, so there are 36 pairwise panels. Matters are complicated by the presence of strong cross-section dependence between these pairwise panels, which weakens standard panel cointegration tests. We therefore carry out robustness checks for strong cross-section dependence induced by common factors. Pairwise differences in unobserved regional total factor productivity are captured by pairwise fixed effects. Robustness checks are also carried out regarding unobserved regional differences in total factor productivity over time, which are hypothesized to be stationary or trend stationary. Finally, robustness checks are carried out with respect to the effect of regional investment grants on the elasticity between regional capital stocks.

The hypothesis of perfect instantaneous internal capital mobility is categorically rejected. Matters are more nuanced in the case of perfect long-term internal capital mobility depending on auxiliary hypotheses regarding total factor productivity. For example, if TFP is hypothesized to be trend-stationary, the null hypothesis cannot be rejected at pvalues of about 0.12, although it is rejected at conventional levels of significance. The same applies to the hypothesis of long-run imperfect capital mobility, regardless of how cross-section dependence is specified. However, if regional TFP differentials are assumed to be difference stationary, the hypothesis of imperfect capital mobility cannot be rejected at conventional levels of statistical significance when cross-section independence is ignored. In this case the average elasticity between pairwise capital–labor ratios is approximately 0.7 instead of 1. In summary, although the hypothesis of instantaneous perfect capital mobility is overwhelmingly rejected, the same does not apply to the hypothesis of long-term perfect capital mobility.

The usual disclaimers apply. First, if capital is miss-measured the results could be induced by measurement error. Since plant is measured directly the main source of measurement error would be in the regional allocation of machinery. Beenstock et al. (2011) have shown that similar measures of capital provide a good empirical account of regional wages over time, which increases confidence in the quality of the data. Moreover, the results should be robust with respect to stationary measurement error since the methodology is based on panel cointegration.

Second, the results are for Cobb–Douglas technologies, which are conveniently loglinear. In the CES case, for example, where:

$$Q_j = A_j[aK_j^{\rho} + (1-a)L_j^{\rho}]^{1/\rho} \tag{13}$$

it may be shown that PICM implies that the pairwise elasticities are:

$$\frac{\partial \ln k_j}{\partial \ln k_i} = \frac{1 - \alpha \frac{k_i^{\rho}}{\alpha k_i^{\rho} + 1 - \alpha}}{1 - \alpha \frac{k_j^{\rho}}{\alpha k_j^{\rho} + 1 - \alpha}} \tag{14}$$

The elasticity is greater than 1 when k_j exceeds k_i, and it tends to 1 as ρ tends to zero, as in the Cobb–Douglas case. The elasticity varies inversely with k_i and varies directly with k_j. Consequently, testing PICM for CES technologies involves nonlinear spatial panel cointegration theory, which has not yet attracted attention in the literature on spatial panel data, and therefore lies beyond the present terms of reference. Nevertheless, Eq. (14) serves to remind us that indirect empirical tests of PICM depend on auxiliary hypotheses regarding the technology of production.

This unsatisfactory situation would not have arisen had there been data on regional returns to capital, in which event PICM could have been tested directly. Even incomplete data on regional returns to capital would have provided an opportunity for external validation of the results obtained from indirect tests. In the complete absence of data on regional returns to capital external validation remains elusive.

If in a small country such as Israel it is difficult to find evidence in favor of perfect or even imperfect capital mobility, in large countries it might be even more difficult. At the very least, the results challenge the consensus that instantaneous perfect internal capital mobility is an innocuous empirical assumption in spatial general equilibrium theory.

References

Anselin, J.-L.: Spatial Econometrics: Methods and Models. Kluwer, Boston (1988)

Banerjee, A., Carrion-I-Silvestre, J.L.: Testing for panel cointegration using common correlated effects estimators. Dept of Economics, University of Birmingham. (2011, 2014)

Beenstock, M., Ben Zeev, N., Felsenstein, D.: Capital deepening and regional inequality: an empirical analysis. Ann. Reg. Sci. **47**, 599–617 (2011)

Beenstock, M., Felsenstein, D.: Estimating spatial spillover in housing construction using nonstationary panel data. J. Hous. Econ. **28**, 42–58 (2015)

Beenstock, M., Felsenstein, D.: Econometric Analysis of Nonstationary Spatial Panel Data. Springer, Berlin (2017). (forthcoming)

Bernard, A.B., Redding, S.J., Schott, P.K.: Testing for factor price equality with unobserved differences in factor quality and productivity. Am. Econ. J. Microecon. **5**, 135–163 (2013)

Caliendo, L., Parro, D., Rossi-Hansberg, E., Sarte, P.-D.: The impact of regional and sectoral productivity changes on the US economy. NBER Working Paper 20168 (2014)

Chudik, A., Pesaran, M.H., Tosetti, E.: Weak and strong cross-section dependence and estimation of large panels. Econom. J. **14**, C45–C90 (2011)

Garofalo, G.A., Yamarik, S.: Regional convergence: evidence from a new state-by-state capital stock series. Rev. Econ. Stat. **84**, 316–323 (2002)

Glaeser, E.L., Gottlieb, J.D.: The wealth of cities: agglomeration economies and spatial equilibrium in the United States. J. Econ. Lit. **47**, 983–1028 (2009)

Im, K.S., Pesaran, M.H., Shin, Y.: Testing for unit roots in heterogeneous panels. J. Econom. **115**, 53–74 (2003)

Krugman, P.: Geography and Trade. MIT Press, Cambridge (1991)

Nocco, A.: Preference, heterogeneity and economic geography. J. Reg. Sci. **49**, 33–56 (2009)

Pedroni, P.: Critical values for cointegration tests in heterogeneous panels with multiple regressors. Oxford Bull. Econ. Stat. **61**, 653–670 (1999)

Pedroni, P.: Panel cointegration: asymptotic and finite sample properties of pooled time series with an application to the PPP hypothesis. Econom. Theory **20**, 597–625 (2004)

Pesaran, M.H.: A simple panel unit root test in the presence of cross section dependence. J. Appl. Econom. **22**, 265–312 (2007)

Pesaran, M.H.: Estimation and inference in large heterogeneous panels with a multi-factor error structure. Econometrica **74**, 967–1012 (2006)

Pesaran, M.H.: Testing weak cross-sectional dependence in large panels. Econom. Rev. **34**, 1087–1117 (2015)

Roback, J.: Wages, rents and the quality of life. J. Polit. Econ. **90**, 1257–1278 (1982)

Sarafides, V., Wansbeek, T.: Cross-sectional dependence in panel data analysis. Econom. Rev. **31**, 483–531 (2012)