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Public capital and productive economy profits: evidence from OECD economies

Ivan D. Trofimov[*]

Abstract

This paper examines the effects of public capital and government final consumption expenditure on the rate of profit in the productive sectors of the OECD economies over the period of 1977-2006. Public capital (expressed as a proportion of private capital) is considered in a multivariate setting, alongside other determinants of profit. The panel cointegration and panel vector autoregressive (PVAR) models are used to remedy the shortcomings of the time series analyses in the short samples and the stationary data panel models. The study demonstrates the absence of cointegration between the variables, but the positive and significant effects of public capital that are particularly manifest in the short-run, as well as the negative and insignificant impact of overall government consumption expenditure. The paper highlights the importance of public capital for macroeconomic outcomes, the relevance of the real channels of fiscal policy, and the non-neutrality of the type of government expenditure for economic outcomes.

Keywords: Public capital, profit, panel data

JEL Codes: E22, H54, C23

1. Introduction

The role of public capital[1] in the economy is considered salient in academic research and policymaking circles for a number of reasons. First, in recent decades, fiscal activism by the state has been an important economic factor, despite the deregulation and liberalisation processes that have been underway since the 1980s. Public investment and fiscal spending have been particularly substantial in the areas of physical and social infrastructure. According to Kamps (2004: 28-29), across the 22 OECD economies, government net capital stock (as a percentage of GDP) ranged from 57.8% in 1960 to 51.4% in 2000, while the average government net capital stock per capita (at 2000 PPP in US dollars) increased from $9,271 to $12,137 between 1980 and 2000 and tracked closely the GDP per capita growth. The shares of government fixed capital formation and final consumption expenditure in GDP likewise remained substantial: according to the World Bank (2019), the former variable fluctuated between 23.6% and 29.1% across the OECD economies during the 1980-2014 period, and the latter ranged from 16.3% to 17.3% during 1980-2017, peaking during the GFC years (19.2%).[2]

Second, the pace of public capital accumulation has been considered against the backdrop of falling productivity in the developing economies. Following Aschauer (1989), the slowdown of public capital formation was seen as an important explanatory factor in the productivity slowdown, with the acceleration of public investment a force intended to reverse these trends. More recent literature on the infrastructure gap (Woetzel et al. 2016) has examined infrastructure underinvestment and associated problems (economic inefficiencies caused by infrastructure deterioration, threats to countries' certain critical systems) across the globe, with research concerning ways of addressing the problem (in particular, whether public capital could be part of the solution). The rise of public-private partnerships and the greater involvement of the private sector in infrastructure provision stimulated the debate around the (new) role of public capital in such partnerships, as well as other related issues (risks, incentives, funding arrangements, transparency, and the level of public-private partnership capital (IMF 2017a)).

Third, the 1970s witnessed the decline of profit rates across the industrialised economies and the profit squeeze (in turn, indicating the decline in capital productivity); and in the 1980s and 1990s, profit rates rebounded. The literature tended to attribute this revival to the decline of the bargaining power of labour (and the reduction in labour shares), changes in the composition of labour force and the rise of unproductive labour, and rising capital productivity, among other factors (Moseley 1997;

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[1] This paper adopts a narrow definition of the public capital, i.e. the tangible capital stock (core infrastructure, such as roads, railways, airports, utilities, as well as hospitals, educational and other public facilities) owned by the public sector, excluding military structures and equipment (Ligthart and Suarez 2011: 8), intangible assets and human capital.

[2] Indicator code - NE.CON.GOVT.ZS, available at https://data.worldbank.org/indicator/NE.CON.GOVT.ZS?locations=OE, accessed on 4 April 2019.

Duménil and Lévy 2004a). It is instructive, however, to examine whether fiscal activism and government investment made any contribution to the revival. As this paper demonstrates, there was ambiguous evidence of the long-term association between profit rates and government capital and investment, but the short-term effects were positive.

Research into the impact of public capital and investment on the economy has been extensive, exploring issues such as estimation of the value of the public capital stock (Kamps 2006); the contribution of public capital to economic and productivity growth (Aschauer 1989, 1990); the complementarity between public and private capitals, and, in contrast, the crowding out phenomenon (Gramlich 1994); the effects of government spending and fiscal policy on economic activity in general (Mishkin 1982); methods of financing a stream of government expenditure (Barro 1974); and the impact of public capital on profitability (Aschauer 1988). The latter aspect has also been examined in heterodox economics literature within the Kalecki equation framework: specifically, it was hypothesized that expenditure by the public sector and the running of a budget deficit by the government would have positive effect on the profitability of the private sector, boosting the revenues of the private sector through government purchases and the purchases of the households through transfer payments, and by reducing the tax burden on the private sector (Toporowski 1999: 363).

Studies of the effects of public capital on the profitability of the private sector are scarce, and the existing literature in this area related exclusively to the US economy (Aschauer 1988; Lynde 1992), covers the 1950s-1980s, and uses early generation econometric techniques, specifically ordinary least squares (OLS) regression. Studies of other economies did not examine the contribution of public capital per se, but rather the consequences of fiscal activism and budget balances for profitability (Asimakopulos 1982; Bakir and Campbell 2016). The purpose of this paper is to re-examine the public capital-profitability relationship in a larger sample, using alternative regressors, more recent data, and econometric methods. Given the biased results of the time series models in short samples, we apply panel vector autoregression (PVAR) and panel cointegration models.

The rest of the paper is organised as follows. Section 2 reviews the previous research concerning the effects of public capital on economic variables, and specifically on profitability. Section 3 describes the methodology used and the data sources, while Section 4 presents empirical findings. Section 5 contains concluding remarks.

2. Literature review

The literature review that follows provides a cursory summary of the previous studies pertaining to various aspects of public capital effects. (Given the sheer volume of the literature, this includes only a proportion of the relevant studies.)

The first strand of research concerned data issues. The estimation of the value of the public capital was initially limited to individual economies: Sweden (Berndt and Hansson 1991), the Netherlands (Sturm and de Haan1995), and the local and state governments of the US (Munnell 1990). Given the absence of comparable capital stock data, further studies attempted to construct methodologically consistent capital stock series for a subset of countries (22 economies (Kamps 2006), 48 (Arslanalp et al. 2010), 71 (Gupta et al. 2014) or the majority of countries (IMF 2017)), using some variant of the perpetual inventory method, obtaining initial levels of capital stock from the hypothetical investment series pertaining to the earlier periods, and allowing different depreciation rates for public versus private capital, and for respective country groups. The major challenges in this research strand included too stringent assumptions about initial stocks and depreciation profiles, changes in the composition of the capital stocks (e.g., relative decrease in road and highways investments), and a lack of disaggregated data (such data is currently available in just a handful of economies).[3]

The second group of studies concerned productivity and growth effects of public capital. The empirical research employed a wide range of methods: OLS models applied to production functions, with public capital as one of the regressors (Aschauer 1989; Cohen and Morrison Paul 2004); VAR/VECM to deal explicitly with dynamic effects, without imposing aprioristic theoretical structures (Otto and Voss 1998; Pereira 2001; Kamps 2004b); behavioural models, employing cost or profit functions to examine the impact of public capital on firms' costs or profits (Ishaq Nadiri and Mamuneas, 1991); cross-sectional or panel models, relating per capita output growth and public investment-GDP ratios (Ligthart and Suarez 2011: 9); and dynamic stochastic general equilibrium (DSGE) models (Kamps 2004b).

As to the findings, the early work by Aschauer (1989) identified the positive consequences of government capital expenditure, though later studies raised some qualifications.

[3] An example of detailed data is the US Bureau of Economic Analysis series for the ten categories of capital.

1). The positive aspects were not experienced uniformly, with core infrastructure being the most productive in the developed economies (Aschauer 1990; Sturm and De Haan, 1995) and public investment in education, health, and infrastructure the most productive in the developing economies (World Bank 2007).

2). The size of the effects was found to be much smaller than Aschauer's original estimate for the US economy (Gramlich 1994; Sturm et al. 1998; Demetriades and Mamuneas, 2000). The impacts of public capital at the regional level were smaller than those at the national level, indicating spillover (Evans and Karras, 1994).

3). The countries with higher income levels and initial levels of the capital stock were likely to experience more substantial positive effects of the public capital, thus highlighting the important role of institutional factors (Arslanalp et al. 2010: 5).

4). Unlike that for public capital, the evidence for positive results of public investment was mixed (Romp and De Haan, 2005), as a result of the differential growth pace of public capital stock and investment, the cyclical fluctuations in public investment (public investment as a first expenditure item to be cut in times of tight budgets (Roy et al. 2006)), and the possibility of negative consequences of public investment for output (due to the contractionary effects of the tax rates needed to maintain the current level of public capital (Aschauer 1998)).

5). In many instances, the positive effects of public capital were absent (e.g., Kamps (2004b) found no such effects for employment, while Ferrara and Marcelino (2002) identified none for output in a number of Italian regions) or insignificant (Ligthart (2002) in the case of Portugal, and Otto and Voss (1998) in the case of Australia).

The third stream of literature concerned the relationship between public and private capital and the relative salience of each type in stimulating economic activity. Erden and Holcombe (2006: 480) noted the likely complementarities between the capitals: the reduction of private sector costs and increase in private sector productivity due to public spending (which is particularly the case in public infrastructure investment); the reduction of uncertainty associated with the irreversibility of private investment (Dixit and Pindyck 1994) or with missing credit markets when complementary public capital (e.g., infrastructure) is provided; and the signalling role of public capital investment, in which it leads by example and signals to the private sector the adequate level of profitability. Erden and Holcombe (2006: 481) also identified the reasons for crowding out: the competition between the capitals for scarce resources, negative macroeconomic effects when public investment is financed through deficit spending or debt creation, and competition in the product markets when similar types of goods are produced using public and private capital. Gutierrez (2005) further noted that the issues of public capital productivity and the complementarity of the capitals are related: when public investment complements private, its impact on productivity and growth is positive; in contrast, the negative consequences of public capital are experienced when there is low complementarity or crowding out, (e.g., when public projects with low or negative social returns are implemented and the scarce resources of the private sector are drained).

The studies yielded mixed findings regarding the effects of public capital on private: crowding out was identified in Germany (Knot and De Haan 1999), India (Serven 1996), in a panel of 116 developing economies (Cavallo and Daude, 2011), and in a set of five European economies (Knot and De Haan, 1995). Complementarities of various magnitude were identified in the US (Erenburg 1993) and Canada (Dadgostar and Mirabelli 1998), and for a panel of 17 developed economies (Afonso and St. Aubyn 2008), six industrialised countries (Mittnik and Neumann 2001), and developing economies (Khan and Kumar 1997), among others. Mixed effects were identified in Greece (Apergis 2000) and in developing countries (Atukeren 2006).

Concerning the specific topic of this paper, the competing paradigms (Keynesian-Kaleckian and neoclassical) postulate different relationships between profits and investment. In a neoclassical paradigm, profits determine investment (given the technological conditions of full employment and savings propensities of the business sector); while in the Keynesian or Kaleckian paradigm, the causality runs from investment to profits (given the absence of full employment, and exogenous investment and money wage rates (Gupta 1988)). Government investment and fiscal activity may thus add to private investment and stimulate private profit rates (by helping the private sector create its own profits) along Keynesian and Kaleckian lines; or, they may positively alter the marginal product of private capital and its rate of return and further stimulate private investment expenditure along neoclassical lines.[4]

[4] According to Meade (1952), the public capital is not purchased by the private businesses in the same way as private capital or labour, but is rather provided by the government in return for a lump-sum tax payment. Assuming little or no control of the private sector over the public capital supply decision, the variable is treated as unpaid fixed input that affects private firms' variable costs and profits (Meade 1952; Deno 1988: 400).

Aschauer (1988), in the early study of the relationship between public capital and profitability in the private sector, hypothesized the relevance of fiscal decisions on economic outcomes and complementarity between public and private capitals, proposing that public investment policy affects the level of private investment by changing the marginal product of capital and private capital returns. He argues that in the 1950-1980s period, the decline in public capital vis-à-vis private capital was responsible for private profitability deterioration. In the study – which considered US series over the 1953-1985 period and used the OLS model and instrumental variables estimation – the overall government expenditure (represented as a ratio of government final consumption expenditure and private capital stock) had a negligible effect on private profitability. In contrast, the elasticities of the private rates of return to public and private capitals were positive and negative, respectively (1% increase in public capital increasing gross and net private rates of return by 19.1 and 21.4 basis points, while a 1% increase in private capital resulted in falls of 38.4 and 38.1 basis points).

A study of the US nonfinancial corporate sector in the 1958-1988 period (Lynde 1992) arrived at similar conclusions. This derived profit rate from a homogeneous production function, with labour, public and private capital inputs, and estimated OLS regression with time trend and a set of regressors (the logs of labour-capital ratio, the ratio of public to private capital, and the private capital stock). The study indicated a trend decline of both profit rate and the ratio of public to private capital, thus demonstrating the positive contribution of public capital to output and profits (albeit smaller than that originally estimated by Aschauer (1988)). The 1% increase in public capital stock (US$16 billion at 1982 prices) increased the profit rate by 1.2% (US$2.6 billion in additional profits at 1982 prices). The study further considered the disaggregated public capital data and identified provision of public capital by state and local governments as a major source of the positive effects on private profitability.

The regional economics literature (Eberts and Fogarty 1987; Deno,1988) did not focus explicitly on the profitability of the private sector, but instead considered aggregate profit functions for the manufacturing sector in particular regions to estimate the elasticities of variable input demand and output supply with respect to public capital. The underlying hypothesis was of a substantial and positive role of public capital in stimulating regional profits. According to Deno (1988: 400-401), public investment had a positive impact on the variable costs and profits of regional manufacturing businesses and induced existing businesses to expand operations and new firms to enter the respective regions. The virtuous cycle created by rising private profits included the expansion of business investment and labour demand, a rise in real wages, a further increase in private capital stock, and in the long-term, structural change towards more profitable and higher value added industries. The study by Deno, which used normalised translog profit function and data for 36 standard metropolitan statistical areas (SMSA) in the US for the 1970-1978 period, ascertained positive effects on manufacturing performance of disaggregated (roads and highways, sewage facilities, and water supply and treatment facilities) and total public capital – which additionally included health and welfare, police and fire, and recreation facilities. This was most pronounced for aggregate public capital stock, while the role of particular types of the public capital varied across the growing and declining regions.

The studies of the determinants of profitability along Kalecki lines focused on government saving and dissaving (as represented by budget deficit or surplus) as opposed to public capital or the level of government investment or consumption per se. Asimakopulos (1982) demonstrated that budget deficit was not a major determinant of US profits (gross retained earnings in the US national income and product accounts (NIPA)) during any of the post-WWII business cycles (1950-1982) – in contrast to private investment and personal savings, which affected profits positively and negatively, respectively. The budget deficit was a positive contributor to profits, particularly in the earlier part of the sample. Asimakopulos (1982: 17-18) attributed variation in influence to the adverse impact of interest rates (as a consequence of deficit financing) that deter investment and thus offset the positive contributions of the budget deficit, as well as to conditions in financial markets and inflation uncertainty. Bakir and Campbell (2016) expressed Kaleckian variables (profits, net private investment, personal dissaving, government deficit, and international surplus) as a percentage of net private fixed assets and considered after-tax profit in the US domestic sector over the 1947 Q1 to 2014 Q3 period. The finding differed from that of Asimakopulos, as the fiscal activity of the government and its deficit continued to be an important determinant of profitability, second to net private domestic investment, and its contribution persisted despite the deregulation and government downsizing that has taken place since the 1980s.

The analysis that follows builds upon the above studies, while considering a broader set of economies and applying more up-to-date econometric techniques. In the absence of comparable and consistent profit rate series for the non-financial corporate profit rates of the OECD economies, we estimate the approximate productive economy profit rates. (This may be an appropriate approach, given that government economic activities are confined to sectors excluded from the productive profit rate calculation, and that privatisation process of the past decades reduced government production activities in sectors such as utilities and transportation.)

3. Methodology and data

3.1 Model

Following Lynde (1992: 129-30) the relationships between profit rate on one hand and public and private capital on the other are derived from homogeneous production function:

$$Y = f(N, G, K) \tag{1}$$

,where N represents the quantity of labour services, and G and K represent public and private capital stocks respectively. Assuming that function is homogeneous of degree one $(\lambda = 1)$, Equation (1) transforms as:

$$f\left(N/K, G/K, 1\right) = \left(1/K\right)^{\lambda} f\left(N, G, K\right) \tag{2}$$

The output Y is then expressed as:

$$Y = f(N, G, K) = K^{\lambda} \phi\left(N/K, G/K\right) \tag{3}$$

,where

$$\phi\left(N/K, G/K\right) = f\left(N/K, G/K, 1\right) \tag{4}$$

Equation (3) is then substituted into the profit rate equation:

$$\pi = (1 - \lambda + S_g)\left(Y/K\right) \tag{5}$$

,where Y/K is capital productivity and S_g is government capital share of output.
Profit rate is then represented as:

$$\pi = \left(1 - \lambda + \frac{\partial \phi}{\partial g} \frac{g}{\phi}\right) K^{\lambda - 1} \phi(n, g) \tag{6}$$

,where $n = N/K$ and $g = G/K$ are ratios of labour to private capital and public to private capital and λ are returns to scale.

Simplifying Equation (6) we get:

$$\pi = H\left(K, n, g\right) \tag{7}$$

The empirical analysis of the effects of public capital and investment on productive economy profitability was then conducted by means of two specifications as follows:

$$PR_{it} = \beta_0 + \beta_1 GKPK_{it} + \beta_2 LK_{it} + \beta_3 PK_{it} + \varepsilon_{it} \tag{8}$$

$$PR_{it} = \beta_0 + \beta_1 GFCEPK_{it} + \beta_2 LK_{it} + \beta_3 PK_{it} + \varepsilon_{it} \tag{9}$$

,where PR_{it} is the productive economy profit rate of country i in period t, $GKPK_{it}$ is the ratio of the value of the public capital stock to the private capital stock, $GFCEPK_{it}$ is the ratio of the value of the government final consumption expenditure to the value of private capital, LK_{it} is the measure of the capital intensity (the inverse of the ratio of private capital to the number of employees), PK_{it} is the product of public capital and GDP gap (to account for the cyclical fluctuations in capacity utilization, in line with approach taken by Lynde 1992). The effects of public capital and government final consumption expenditure on profit rate are contrasted, following Aschauer (1988). All variables were transformed to logarithms to reduce the effects of heteroscedasticity and skewness on the empirical estimates.

3.2 Data

The paper considered a panel of 9 industrialized economies belonging to the OECD group: Austria, Denmark, Finland, Germany, Italy, Netherlands, Spain, the UK and the USA. These economies formed a long panel consisting of 270 observations (30 observations per panel stretching 1977-2006 period). The choice of the sample and the study period was influenced by the lack of the consistent data on sectoral capital stock, depreciation allowances and gross operating surplus (that are used to calculate productive economy profit rate) for a greater number of OECD economies in the available databases.[5] The use of panel data and models is warranted, given the sufficient amount of trade and flows of investment across these economies (and the economic integration initiatives and processes in Europe), hence certain commonalities in the dynamics of variables.

The productive economy profit rate is estimated as the ratio of the net operating surplus in the productive economy to the net capital stock in the productive economy as follows:

$$PR = \frac{(C / P_k) - D}{K_r - D} \times 100 \tag{10}$$

,where K_r is real fixed capital stock at 1995 prices (K_GFCF code in EUKLEMS database), D is consumption of fixed capital at 1995 prices (D_GFCF), C is capital compensation at current prices (CAP_GFCF), and P_k is gross fixed capital formation price index with the base in 1995 (lp_GFCF). The variables were obtained from the capital input data files of the EU-KLEMS database (November 2009 release).

The data for the corporate (private) sector is not available for all economies in question and estimation of the corporate profit rate is precluded. As an alternative, following Duménil and Lévy (2004b), a productive economy profit rate is used, i.e. the rate for the aggregate of the sectors that are the most relevant for capitalist reproduction (agriculture, mining, manufacturing, utilities, construction, transport, storage and communication, wholesale and retail trade, and hotels and restaurants), excluding those sectors (or economic activities) that are deemed unproductive or that have indirect relationship to the wealth creation and private profitability (real estate, government sector, finance and insurance, as well as social reproduction sectors of health, education, social, and community work).

The government capital-private capital ratio was constructed using net public and private capital series (kpriv_ppp and kgov_rppp indicator codes) obtained from the IMF investment and capital stock data set (IMF 2017).[6] The ratio of government final consumption expenditure to net private capital was estimated using 'final consumption expenditure of general government at 2010 prices' series from

[5] For instance, the EU-KLEMS database (used in this paper) does not include the capital data for Belgium. The capital series for Ireland, Sweden and Greece start in the mid-1990s, and for Portugal and Eastern European members of OECD in the 2000s.

[6] Available at
https://www.imf.org/external/np/fad/publicinvestment/data/data122216.xlsx accessed on 1 February 2019.

AMECO database (indicator code UCTG0), converted to 2011 prices to ensure consistency with other variables in question.

The net public capital values were derived from the net public investment by the general government (national and sub-national), excluding other means and methods of government investment: investment grants, loan guarantees, tax concessions, government-backed saving schemes, and funding by the government financial institutions, such as development banks and public bodies outside the general government (IMF, 2017b: 1). The values of capital stock were separated from the the value of the public-private partnerships (PPP) capital. The private and public capital stocks were estimated by the perpetual inventory method as:

$$K_{it+1} = (1-\delta_{it})K_{it} + (1-\delta_{it}/2)I_{it} \tag{11}$$

,where for country i, K_{it+1} is the stock of public (private capital) at the beginning of period, δ_{it} is time-varying depreciation rate, and I_{it} is the gross fixed capital formation (IMF 2017b: 2).

3.3 Econometric approach

The length of the sample period in this paper is 30 years, which may be relatively short to justify the use of time series models that tend to have low power in short samples. The early generation panel models (such as pooled OLS, or models with fixed and random effects) do not capture the dynamic relationships or distinguish between short- or long-run effects. Such consideration may be necessary in light of previous research (e.g. Nijkamp 1986, who examined the long- and short-run effects of public capital in the context of public investment impact on regional growth). Likewise, the use of certain dynamic panel data models (generalised method of moments) may be unwarranted in long samples ($T > N$). In addition, the variables in consideration may have different order of integration: we expect that capital intensity and GDP gap variables are I(0) while capital variables are I(1). This will make it impossible to use fully modified and dynamic OLS models (FMOLS and DOLS). Instead we adopted a procedure that includes panel autoregressive distributed lag (ARDL) and panel vector autoregressive (PVAR) models.

Firstly, given the particular dimension of the panel where time dimension does not differ substantially from the cross-sectional dimension ($N = 9$, $T = 30$), we verified the presence of cross-sectional dependence using a two tests: Breusch-Pagan (1980) CD_{LM1} test for cross-sectional dependence (with a null hypothesis of no cross-sectional dependence and its presence as an alternative) that is suitable for cases when time dimension of a panel data set exceeded its cross-sectional dimension ($T > N$); and Pesaran (2015) test of weak cross-sectional dependence (with weak dependence as a null hypothesis and strong dependence as an alternative) that is more flexible with regard to N and T combinations.

Secondly, the order of integration of variables is tested using a battery of panel unit root tests: Im-Pesaran-Shin/IPS, Levin-Lin-Chu/LLC, Breitung, ADF-Fisher χ^2 and PP-Fisher χ^2 tests (Maddala and Wu 1999; Breitung 2000; Choi 2001; Levin et al. 2002; Im et al. 2003), and a more robust cross-sectionally augmented IPS test (CIPS) by Pesaran (2007).[7] The CIPS test estimates Augmented Dickey-Fuller (ADF) regression (with cross-section terms) for individual series, averages the t-statistics of the coefficient of the lagged term (b_i) for ADF regression, and contrasts the null hypothesis of $b_i = 0$ for all i (non-stationarity of all series) with the alternative hypothesis of stationarity of at least one of the series.

Thirdly, the presence of cointegration relationships was examined by means of Pedroni (2004), Kao (1999), and Westerlund (2007) tests of cointegration. The statistical significance of the error correction coefficient in panel ARDL model was also considered as an indication of possible cointegration (Kremers et al. 1992).

Pedroni (2004) test involved regression of residuals from the cointegrating equation against residuals' lagged values:

$$\hat{\varepsilon}_{it} = \hat{\gamma}_i \hat{\varepsilon}_{it-1} + \hat{\mu}_{it} \tag{12}$$

[7] The tests proposed by Maddala and Wu (1999) and Choi (2001) involve calculation of ADF-Fisher χ^2 and PP-Fisher χ^2 statistics.

to obtain $\hat{\gamma}_i$, the autoregressive coefficient of the residual $\hat{\varepsilon}_{it}$. The null hypothesis of no cointegration $H_0 : \gamma_i = 1$ was then contrasted with the homogeneous (common γ_i) and heterogeneous alternatives, $H_1 : (\gamma_i = \gamma) < 1$ for all i and $\gamma_i < 1$ for all i, and the respective panel and group mean statistics were constructed under both alternatives.

Kao (1999: 3-7) cointegration test was based on applying DF and ADF tests to the residuals from the least-squares dummy variable (LSDV) regression model:

$$\hat{e}_{it} = \rho \hat{e}_{it-1} + \upsilon_{it} \tag{13}$$

and contrasting the null hypothesis of no cointegration and nonstationarity of residual series ($H_0 : \rho = 1$) against the alternative hypothesis, $H_1 : \rho < 1$.

Westerlund (2007) cointegration test was based on structural estimation, and allowed for individual specific short-run dynamics and cross-sectional dependence. The data-generating process was given as (Persyn and Westerlund 2008: 233):

$$\Delta y_{it} = \delta_i' d_t + \alpha_i (y_{i,t-1} - \beta_i' x_{i,t-1}) + \sum_{j=1}^{p_i} \alpha_{ij} \Delta y_{i,t-j} + \sum_{j=-q_i}^{p_i} \gamma_{ij} \Delta x_{i,t-j} + e_{it} \tag{14}$$

,where d_t represented deterministic components (trend and constant, constant, or the absence of trend and constant). The null hypothesis of no error correction and no cointegration, $H_0 : \alpha_i = 0$ for all i, was compared to the alternative hypotheses of the presence of error correction and cointegration for at least one i, i.e. $H_1^g : \alpha_i < 0$ (group-mean tests), or the presence of cointegration for all i, i.e. $H_1^p : \alpha_i = \alpha < 0$ (panel tests).

Fourthly, we implemented panel ARDL, assuming that there exists a long-run relationship among the variables (as represented by the statistical significance of the error-correction coefficient in the ARDL equation). Three alternative estimators are employed to obtain panel ARDL coefficients (Pesaran and Smith 1995; Pesaran et al. 1999): pooled mean group (PMG), that imposes homogeneity restriction on the long-run coefficients, mean group (MG), that does not impose the restrictions, and dynamic fixed effects (DFE), that pools time series data for each group and allows only the intercepts to differ across groups. The Hausman test was then used to ensure the appropriateness of restrictions and to select the relevant estimator. Common lag structure was set: given the number of observations in each panel, one lag was selected for dependent variables and regressors (i.e. $p = 1, q = 1$).

Notation-wise, the panel ARDL with the error-correction component had the following form:[8]

$$\Delta PR_{it} = \sum_{j=1}^{p-1} \psi_j^i \Delta PR_{it-j} + \sum_{j=0}^{q-1} \phi_j^i \Delta X_{it-j} + \delta^i \left[PR_{it-1} - \left\{ \beta_0^i + \beta_j^i X_{it-1} \right\} \right] + \varepsilon_{it} \tag{15}$$

,where PR is the profit rate, X is one of the regressors from the equations (8) or (9), p and q are the lags of the profit rate and the regressors, ψ and ϕ are the short-run coefficients of the profit rate and the regressors, β is the long-run elasticity, δ is the error-correction term, that indicates the speed of adjustment to the long run equilibrium, i and t are country and time indicators. When the PMG estimator was used, the homogeneity restrictions were imposed on β_j coefficients.

The long-run coefficients were derived from equation (15) as:

$$\Delta PR_{it} = \beta_0^i + \beta_j^i X_{it} + \mu_{it} \tag{16}$$

,where μ_{it} was $I(0)$.

Fifthly, given the inconsistency of cointegration test results, I(1) order of the series as well as contemporaneous correlation between cross-sections (as demonstrated further), Chudik-Pesaran (2015)

[8] The description of panel ARDL model is adopted from Elveren and Hsu (2016: 564).

dynamic common correlated effects (DCCE) estimator that is robust in the case of (non-)stationary data, as well as presence (or absence) of cointegration (NguyenHuua and Orsal 2019: 13: 14; Kapetanios et al. 2011) was applied. DCCE extends common correlated effects (CCE) estimator, developed by Pesaran (2006), that augments static panel data model with the cross-sectional averages ($\overline{PR_t}$, $\overline{LK_t}$, $\overline{PK_t}$, $\overline{GKPK_t}$ and $\overline{GFCEPK_t}$) and thereby removes cross-sectional dependence. The assumption of the common β coefficients in Equations (8) and (9) is relaxed and the average of the respective coefficients is estimated as $\beta = N^{-1}\sum_{i=1}^{N}\beta_i$, reflecting the possible presence of cointegration, while the short-run dynamics and adjustment in the long-run are reflected in a multifactor error structure $\varepsilon_{it} = \omega_i^{'} f_t + e_{it}$ (Sakyi et al. 2012: 28). To ensure consistency of CCE in a dynamic framework, Chudik and Pesaran (2015) further augment CCE with the p_T lags of the cross-sectional averages.

As a last step, given that all variables were I(1), group mean fully modified OLS (FMOLS) and group mean dynamic OLS (DOLS) estimators were used to establish long-run relationships between the variables (Perdoni 2000, 2001; description of the estimators follow Sakyi et al. 2012: 30-31). Following Pedroni (2001), among the alternative FMOLS and DOLS estimators (pooled and weighted), the grouped estimators were considered to perform better in small samples, have smaller size distortion and yield the true value of coefficients.

The group mean FMOLS estimator is a modification of the conventional pooled OLS estimator. Assuming for presentational purposes two variables y_{it} and x_{it} , the estimator is given as:

$$\hat{\beta}_{FM} = N^{-1}\sum_{i=1}^{N}\left(\sum_{t=1}^{T}(x_{it}-\bar{x}_i)^2\right)^{-1}\left(\sum_{t=1}^{T}\left(x_{it}-\bar{x}_i\right)\right)y_{it}^{*} - T\hat{\tau}_i \tag{17}$$

,where $\hat{\tau}_i \equiv \hat{\Gamma}_{21i} + \hat{\Omega}_{21i} - \frac{\hat{L}_{21i}}{\hat{L}_{22i}}\left(\hat{\Gamma}_{22i} - \Omega_{22i}^{0}\right)$ and $y_{it}^{*} = \left(y_{it} - \bar{y}_i\right) - \frac{\hat{L}_{21i}}{\hat{L}_{22i}}\Delta x_{it}$. The group mean FMOLS is then constructed in a simplified form using the pooled FMOLS estimator as:

$$\hat{\beta}_{gmFM} = N^{-1}\sum_{i=1}^{N}\hat{\beta}_{FM_i} \tag{18}$$

The group mean DOLS estimator is obtained from the dynamic panel model:

$$y_{it} = \alpha_i + \beta_i X_{it} + \sum_{k=-K_i}^{K_i}\gamma_{ik}\Delta X_{it-k} + \mu_{it} \tag{19}$$

,where K_i is the order of lags and leads of the differenced regressors that are used to deal with autocorrelation and to suppress endogeneity. The estimator is then derived as a cross-country average of individual pooled DOLS estimators as follows:

$$\hat{\beta}_{gmDOLS} = N^{-1}\sum_{i=1}^{N}\hat{\beta}_{DOLS_i} \tag{20}$$

To capture the short-run dynamics and to consider the possibility of absence of cointegration (as suggested by some of the tests and statistics), the panel VAR model that involves estimation of the panel regression of each variable in VAR system on its own lags and lags of other variables was set as:

$$z_{it} = \Gamma_0 + \Gamma_1 z_{it-1} + f_i + e_t \tag{21}$$

,where Γ_0 are deterministic components of the data, Γ_1 is a matrix polynomial that captures relationships between the endogenous variables and their lags (Arachi and Assisi 2017: 13), z_{it} is a vector of variables (regressors as well as dependent variables), and f_i are country specific fixed effects.

The order of the VAR system was determined to minimize Schwarz Bayesian Information Criterion and was set to one. The least square dummy variable estimator, devised by Kiviet (1995) and adopted by Cagala and Glogowsky (2014) for panel VAR models was used, arguably being the most appropriate for panels with a large number of observations relative to the small number of cross-sections.

The VAR identification scheme and impulse-response functions' interpretation was based on lower-triangular Choleski decomposition that allows each variable in the vector to react simultaneously only to all variables above itself, and treats the first variable in the causal ordering of variables as having contemporaneous effects on all other variables, but not allowing any of other variables to have contemporaneous effects on the first one (Kireyev 2000: 13; Arachi and Assisi 2017: 11).

We followed the standard approach in the public capital effects literature (Pereira and Roca-Sagales, 2003: 244; Mitze 2012) and put public capital variables (specifically the ratio of the value of the public capital stock to the private capital stock) as the first in the identification scheme, followed by private capital, the labour-capital ratio and profit rate $[LNGKPK_{it}; LNPK_{it}; LNLK_{it}; LNPR_{it}]$. Thus, public capital contemporaneously affected private capital and private sector profit (but not the other way around), while private capital contemporaneously influenced the labour-capital ratio and profit rate (with no reverse causality).

The relevant impulse response functions (IRFs) were plotted to portray the temporal effects of a one standard deviation shock of each respective variable; the forecasting period was set at 8 years, and the confidence intervals are set at 95%. Two types of the confidence intervals were obtained, standard normal confidence intervals and the intervals generated through Monte Carlo simulation with 1000 replication, the latter used in the presentation of IRFs in this paper.

4. Empirical results

As a first step, we considered the common sample descriptive statistics of the variables and the bivariate relationships between profit rate and each regressor. The findings are reported in Figure 1 and Table 1.

The distributions of the public capital (adjusted for capacity utilisation) and the inverse of capital intensity are approximately symmetric around the mean, while the distribution of the ratio of government final consumption expenditure to private capital has high negative skewness and long left tail. The distributions of the profit rate and the ratio of public to private capital are skewed to the right (moderately in the former case, substantially in the latter). The distribution of the inverse of capital intensity is mesokurtic, while the distribution of the private capital (adjusted for capacity utilisation) is moderately platikurtic. All other variables are moderately leptokurtic. For all variables perhaps except the inverse of the capital intensity, the distributions are not normal, as attested by the Jarque-Bera result.

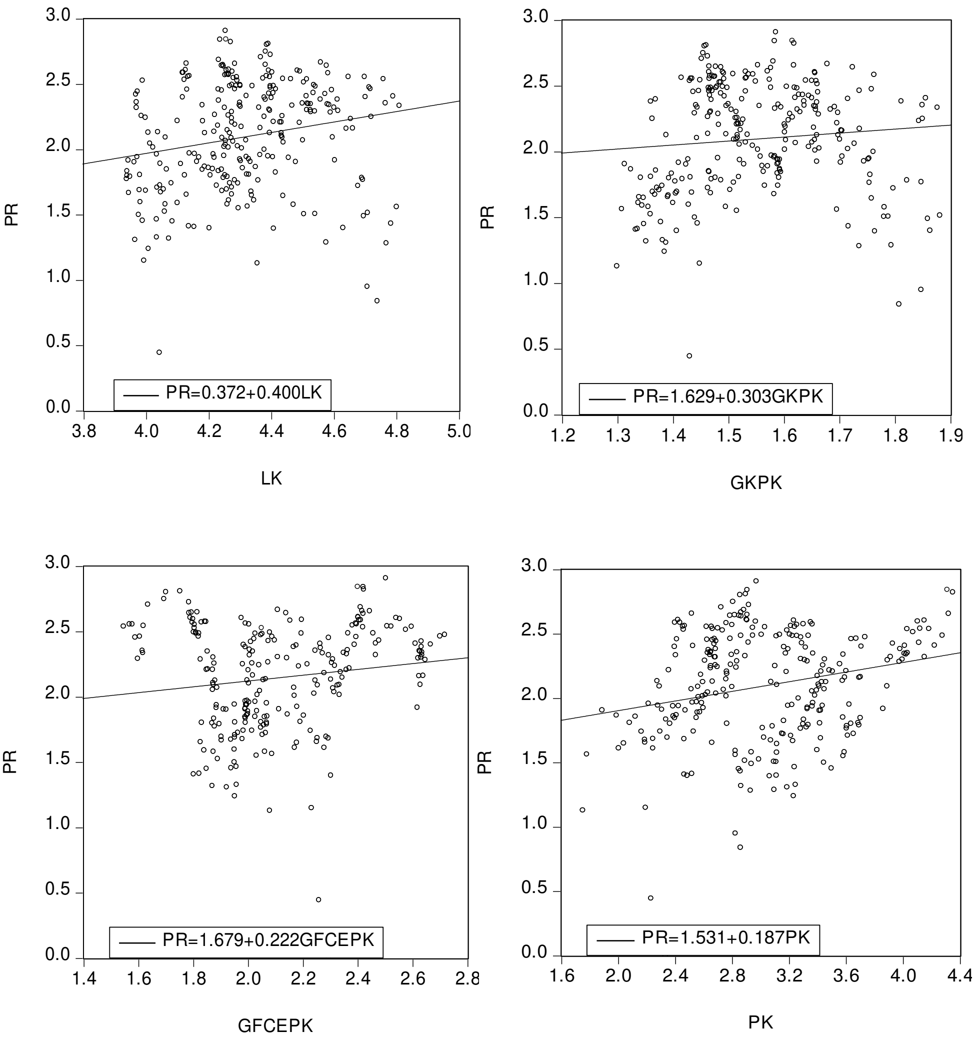
Over the 1977-2006 period, the ratio of public to private capital was positively related to the profit rate. This pattern also persisted over the shorter periods, when the length of the sample was curtailed (e.g., 1977-1990 and 1990-2006). A positive correlation between profit rate and the ratio of government consumption expenditure to private capital was observed throughout. Denmark had a particularly high level of the ratio, reflecting the extent of the welfare state and the degree of the state involvement in the economy (Brons-Petersen 2015). Excluding Denmark did not alter the results: the two variables continued to be positively related. Positive correlations were also present between the capacity-adjusted private capital and the rate of profit (reflecting not only the steady growth of public capital in line with profit rate restoration, but also the countercyclical behaviour of the labour share (Andolfatto 1996; Boldrin and Horvath 1995). A positive relationship was likewise demonstrated between the inverse of the capital intensity and the profit rate.

Table 1 - Descriptive statistics (common sample)

Statistics	PR	PK	GFCEPK	GKPK	LK
Mean	2.065	7.408	1.912	3.661	-0.300
Median	2.106	7.622	2.041	3.600	-0.327
Maximum	2.909	10.081	2.717	4.781	0.204
Minimum	0.001	5.200	0.046	3.068	-0.668
St. deviation	0.431	1.304	0.677	0.369	0.209
Skewness	-0.862	0.069	-1.837	1.103	0.295
Kurtosis	4.726	1.833	5.366	4.349	2.547
Jarque-Bera	62.471	14.491	200.519	70.151	5.798
Prob. (Jarque-Bera)	0.000	0.001	0.000	0.000	0.055

Note. All variables are expressed in logarithms. PK is the logarithm of the product of public capital and GDP gap.

Figure 1. Bivariate relationships: scatter plots of the stacked cross-sections



Note. All variables are expressed in logarithms.

Cross-sectional dependence was likely present in the panel, arising from the degree of economic integration between the economies in question (high volume of trade and foreign investment),[9] as well as from the similarities in institutional structures and the stage of economic development. Table 2

[9] See Akhtar (2018).

contains the results of the Breusch-Pagan (1980) and Pesaran (2015) cross-sectional dependence tests. The null hypothesis of cross-sectional independence was rejected for all variables in question under both tests, indicating contemporaneous correlation between cross-sections (strong form of it under Pesaran 2015, test).

Table 2 - Cross-sectional dependence test results

Test	PR	PK	GFCEPK	GKPK	LK
Pesaran CD	32.297	32.834	31.724	32.794	27.555
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Breusch-Pagan LM	188.988	858.605	529.996	630.271	681.410
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)

Note. All variables are expressed in logarithms. p-values are indicated in parentheses.

The unit root tests present a consistent picture for most variables. The tests of individual effects and no trends run on the levels of each variable (Table 3) indicate that all variables contained unit root, while there was certain evidence of stationarity in the inverse of capital intensity (according to LLC and PP Fisher χ^2 tests and in the capacity-adjusted private capital (according to LLC test). When the tests included both individual effects and individual time trends, the results did not change substantially. All variables in question continued to have unit root, albeit few of the tests pointed to trend stationarity in the profit rate (IPS and ADF-Fisher) and in the ratio of public to private capital (PP-Fisher). In the case of capacity-adjusted private capital, four tests pointed to stationarity, and we therefore conclude that this variable was trend stationary in levels. The test run on the first difference of each variable (Table 4) found that the first differences of all variables were (trend) stationary according to all tests and both specifications (with and without trends). Overall, we conclude that the variables in question had a mixed order of integration: I(1) for profit rate, the ratio of government final consumption expenditure to private capital, the ratio of public to private capital, and the inverse of capital intensity; and either I(1) or I(0) for the capacity-adjusted private capital.

Table 3 - Unit root tests' results (levels)

Test (constant)	PR	GKPK	GFCEPK	LK	PK
Levin, Lin & Chu t*	-0.016	-0.773	-1.184	-2.961	-2.440
	(0.494)	(0.220)	(0.118)	(0.002)	(0.007)
Im, Pesaran and Shin W-stat	-0.624	1.211	0.865	-1.006	0.585
	(0.266)	(0.887)	(0.807)	(0.157)	(0.721)
ADF - Fisher Chi-square	24.138	9.574	12.021	25.356	16.107
	(0.151)	(0.945)	(0.846)	(0.115)	(0.585)
PP - Fisher Chi-square	14.665	12.780	8.780	36.695	55.967
	(0.685)	(0.805)	(0.965)	(0.006)	(0.000)
Pesaran CIPS	-1.519	-1.443	-2.504	-1.743	-1.198
	(0.775)	(0.836)	(0.011)	(0.520)	(0.959)
Test (constant plus trend)	PR	GKPK	GFCEPK	LK	PK
Levin, Lin & Chu t*	-0.772	-1.084	0.229	-1.278	-4.689
	(0.220)	(0.139)	(0.591)	(0.101)	(0.000)
Breitung t-stat	-0.894	1.856	1.147	2.419	-1.908
	(0.186)	(0.968)	(0.874)	(0.992)	(0.028)
Im, Pesaran and Shin W-stat	-1.299	-0.506	0.140	-0.058	-3.548
	(0.097)	(0.306)	(0.556)	(0.477)	(0.000)
ADF - Fisher Chi-square	26.955	24.168	17.934	19.997	43.713
	(0.080)	(0.150)	(0.460)	(0.333)	(0.001)
PP - Fisher Chi-square	14.494	26.681	8.441	6.419	30.862
	(0.696)	(0.085)	(0.971)	(0.994)	(0.030)
Pesaran CIPS	-2.049	-1.882	-2.325	-1.849	-1.886
	(0.800)	(0.942)	(0.480)	(0.932)	(0.914)

Note. All variables are expressed in logarithms. p-values are indicated in parentheses.

Table 4 - Unit root tests' results (first differences)

Test (constant)	PR	GKPK	GFCEPK	LK	PK
Levin, Lin & Chu t*	-6.301	-3.528	-3.501	-3.498	-8.286
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Im, Pesaran and Shin W-stat	-8.023	-4.031	-4.294	-3.540	-8.077
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
ADF - Fisher Chi-square	96.346	49.963	52.462	42.300	96.597
	(0.000)	(0.000)	(0.000)	(0.001)	(0.000)
PP - Fisher Chi-square	132.889	53.888	67.369	30.059	90.587
	(0.000)	(0.000)	(0.000)	(0.037)	(0.000)
Pesaran CIPS	-2.518	-3.384	-2.589	-2.625	-2.343
	(0.009)	(0.000)	(0.005)	(0.004)	(0.036)
Test (constant plus trend)	PR	GKPK	GFCEPK	LK	PK
Levin, Lin & Chu t*	-4.659	-2.589	-2.550	-3.259	-7.013
	(0.000)	(0.005)	(0.005)	(0.001)	(0.000)
Breitung t-stat	-7.235	-1.665	-1.552	-4.001	-3.872
	(0.000)	(0.048)	(0.060)	(0.000)	(0.000)
Im, Pesaran and Shin W-stat	-7.009	-2.229	-2.994	-3.210	-6.269
	(0.000)	(0.013)	(0.001)	(0.001)	(0.000)
ADF - Fisher Chi-square	80.277	31.671	38.420	38.628	70.254
	(0.000)	(0.024)	(0.003)	(0.003)	(0.000)
PP - Fisher Chi-square	126.764	37.851	48.228	19.421	93.312
	(0.000)	(0.004)	(0.000)	(0.366)	(0.000)
Pesaran CIPS	-3.727	-3.632	-2.710	-2.691	-3.325
	(0.000)	(0.000)	(0.098)	(0.109)	(0.001)

Note. All variables are expressed in logarithms. p-values are indicated in parentheses.

The analysis of cointegration is inconclusive, as indicated in Tables 5 and 6. On one hand, the coefficient of the error correction term in the ARDL model (PMG estimator) is statistically significant and falls into the 0 to -1 range (Table 6), providing an indirect indication of cointegration and showing the adjustment to the long-run equilibrium and restoration of the cointegration relationship following an exogenous shock. In all specifications in Table 5, the probabilities of the panel and group ADF statistics (and certain other statistics) in Pedroni cointegration tests are smaller than the critical values, rejecting the null hypothesis and indicating the presence of cointegration. These statistics, according to Pedroni (2004), are the most reliable. Kao test t-statistics likewise reject the null and point to cointegration in both specifications that include constant. On the other hand, in a far greater number of cases, the Pedroni cointegration test does not reject the null hypothesis of no cointegration. Regarding the Westerlund test, the null hypothesis of no error correction and no cointegration is rejected only twice by P_t statistics in Specifications 1 and 2.

The empirical analysis therefore proceeded along two avenues. If the long-run relationship among the variables is indeed present, the estimates are obtained from the panel ARDL, FMOLS, DOLS and DCCE models. In contrast, in the absence of cointegration (and long-run relationship) the short-run dynamics in the panel VAR framework that does not require an assumption of cointegration is examined.

The long-run estimates were obtained from the group mean FMOLS and DOLS models with a constant or constant plus trend (given that a number of variables trended over time). The DOLS equation included one lead and one lag of the first difference of the logarithm of profit rate. In both models the long-run covariance was computed based on Bartlett kernel and Newey-West fixed bandwidth. The DCCE model was implemented with three lags of the cross-sectional averages. The panel ARDL model was implemented with a pooled mean group (PMG) estimator that was, according to the joint Hausman test, deemed to be the most efficient compared to mean group (MG) and dynamic fixed effects (DFE) estimators. The lag structure was set at $p = 1, q = 1$.

In Specification 1 (that had the ratio of public to private capital as regressor), Table 6, the long-run coefficients of this ratio were positive in every model (and significant in all models except DCCE). The long-run coefficient of the inverse of capital intensity was positive and significant only in FMOLS model, while DCCE and panel ARDL models yielded negative coefficients. The effects of capacity-adjusted public capital on profit rate were positive and significant in all models except panel ARDL. The short-run coefficients were positive and significant (with the exception of the ratio of public to private capital) according to the panel ARDL model. The error correction term was negative and significant and falling within $(0;-1)$ range, indicating reasonably high speed of adjustment to equilibrium after a shock (30% annual speed of correction form a previous year disequilibrium). In the DCC model, the cross-sectional dependence (CD) test statistic indicates reduced cross-sectional dependence. The models with constant plus trend as deterministic components (panel FMOLS, DOLS and ARDL) delivered similar significant and positive estimates of the ratio of public to private capital, while the coefficients of the inverse of capital intensity and capacity-adjusted public capital were negative: for the former variable in panel ARDL and DOLS, for the latter in panel ARDL. Alternative lag structures for the ARDL model were tried: the sign and significance of the long run coefficients in the model with $p = 3; q = 1$ (selected using Akaike information criterion) were similar to the model with constant plus trend. To conserve space the results of these additional estimates are not reported here.

In Specification 2 (the ratio of government final consumption expenditure to private capital as regressor), the long-run coefficient of this ratio was positive in all models (and significant in panel FMOLS and ARDL), while long-run coefficients of the other two variables were negative (except for the case of capacity-adjusted private capital in DCCE model). Consideration of models with constant plus trend reveals negative sign of the ratio of government final consumption expenditure to private capital coefficient in both short- and long-run estimates: negative sign of all three coefficients in panel ARDL models, and a negative sign of the ratio of government final consumption expenditure to private capital and the inverse of capital intensity in panel DOLS. A panel ARDL model with alternative $p = 3; q = 2$ structure delivered similar results as ARDL with constant plus trend.

Previous research on public capital - private profit rate relationships (Aschauer 1988; Lynde 1992) that focused on the US economy covered the periods of the concurrent fall in profitability and public capital (specifically the ratio of public to private capital), and thereby established positive effects of public capital on private profit. The study period considered in this paper is 1977-2006, and it is well known and empirically verified that profit rates began to rise across the developed economies in the late 1970s and early 1980s (Basu and Vasudevan 2013), driven by the restoration of the profit share at the expense of labour share, increase in capital productivity and decline in the relative price of capital goods (Dumenil and Levy, 2001: 145: Chou et al. 2016).

Likewise, the decline in public capital that was the most pronounced in the 1970s (driven in part by privatisation of public enterprises, but also by fiscal consolidations in the post-recession periods) slowed down in the 1980s. Certain revival in public capital (as growth rate in capital stock,as a share of GDP and private capital stock, or in per capita) was documented in the US in the 1990-2000s (Bivens 2012), in the UK in the late 1980s-early 1990s (Toigo and Woods, 2006), and across the OECD in 1995-2011 (Allain-Dupré et al. 2012: 6-7). The likely explanations of the phenomenon that were advanced included the growth of public investment of sub-national level (Allain-Dupré et al. 2012: 9-10), the slow-down of privatisation process, with the majority of privatisations having taken place in the earlier period (Toigo and Woods 2006: 75-76); introduction of specific measures to preserve public investments in times of fiscal consolidations, or public investment in infrastructure recovery programs (Schuldes 2011: 28). The revival, however, was short-lived and in many instances limited (e.g. in the US it was confined to the early 1990s and the early 2000s, while further declines were experienced in the second halves of both decades, Bivens 2012).

Whether such moderate revival in public capital and investment has been sufficient to put public capital restoration in line with profit restoration is an open-ended question, and in this regard the empirical results that were documented above (positive long-run relationship/cointegration between the variables) should be taken 'with a grain of salt'. This, however, should not be considered as indication of absent short-run positive effects of public capital on private profitability.

On the other hand, the documented negative and effects of government final consumption expenditure on private profitability are consistent with Aschauer's findings and the theoretical arguments. These arguments propose that it is not the overall level of government spending that matters for economic outcomes, but rather the particular type: namely, spending on non-military capital goods that directly affect the private sector and productive economy, as opposed to transfer payments, social expenditure, and spending on military capital goods (Aschauer 1989).

Regarding the negative effects of private capital and the inverse of capital intensity on the private profit, these can be attributed (apart from the possibility of crowding-out on the part of public capital) to variations in technological change and distributional drivers of profitability across the OECD economies during the recent decades (Basu and Vasudevan 2013; Vasudevan 2013: Ch. 5). The period covered in this paper (1977-2006) encompasses three periods with distinct behavior of these drivers. In the US, The period until the early 1980s witnessed decline in the profit share and the Marx-biased technical change (growth of labour productivity and capital intensity and the fall in capital productivity, Foley and Michl 1999). The period from the early 1980s till the late 1990s was associated with the recovery in profit share and technological change not conforming to Marx-biased pattern (the concurrent growth of capital and labour productivities and sharp decline in capital intensity). This period was characterised by the expansion of both financial and nonfinancial corporate profits and capital saving technological change, in turn driven by the information technology revolution (Mohun 2010). The period from the early 2000s was described by a steep recovery of capital intensity ahead of labour productivity growth, and hence a drastic decline in capital productivity (after the information technology stimulus reached its limits). The profit share, however, was rising steadily in the early 2000s (driven by ongoing income redistribution and financialisation) and this offset capital productivity decline. In this regard, the negative effects of capital intensity (and conversely positive effect of capital intensity inverse) on private profit in the 1980-90s were expected. In contrast, the negative effects of the inverse capital intensity in the early 2000s (fall in L/K while PR rising) were due to acceleration of offshoring that increased K/L and reduced its inverse; but also to stabilisation of the price of capital goods after a long period of decline (in the previous period the cheapening of capital goods attested to rapid technological change in capital goods production, the trend that was reversed in the early 2000s). This latter factor was also a likely explanation of simultaneous increase in capital intensity, growth in profitability and slowdown in private capital accumulation (and hence negative relationship between private capital and profit documented in this paper), the phenomenon that Basu and Vasudevan (2013: 80) termed paradoxical. It is unclear, however, whether similar specific patterns and periodisation were also observed in other OECD economies, given the limited studies on the topic.

Table 5 - Panel cointegration results

Test	Specification							
	1		2		3		4	
Pedroni test								
Panel v-Statistic	1.427	(0.077)	0.191	(0.424)	**1.578**	**(0.057)**	0.342	(0.366)
Panel rho-Statistic	0.592	(0.723)	2.137	(0.984)	1.336	(0.909)	2.347	(0.991)
Panel PP-Statistic	-1.084	(0.139)	**-1.447**	**(0.074)**	-0.063	(0.475)	0.632	(0.736)
Panel ADF-Statistic	**-5.226**	**(0.000)**	**-6.956**	**(0.000)**	**-4.976**	**(0.000)**	**-4.007**	**(0.000)**
Panel v-Statistic (weight.)	0.646	(0.259)	-0.380	(0.648)	0.238	(0.406)	-0.913	(0.819)
Panel rho-Statistic (weight.)	0.618	(0.732)	2.007	(0.978)	0.978	(0.836)	1.677	(0.953)
Panel PP-Statistic (weight.)	-0.180	(0.429)	-0.168	(0.433)	0.355	(0.639)	0.457	(0.676)
Panel ADF-Statistic (weight.)	**-2.634**	**(0.004)**	**-3.882**	**(0.000)**	**-1.614**	**(0.053)**	**-2.351**	**(0.009)**
Group rho-Statistic	1.978	(0.976)	3.278	(1.000)	2.427	(0.992)	3.054	(0.999)
Group PP-Statistic	0.186	(0.574)	0.459	(0.677)	1.146	(0.874)	1.404	(0.920)
Group ADF-Statistic	**-2.957**	**(0.002)**	**-3.643**	**(0.000)**	**-1.358**	**(0.087)**	**-1.950**	**(0.026)**
Kao test (t-stat)	**-2.895**	**(0.002)**			**-2.736**	**(0.003)**		
Westerlund test								
Gt	-1.707	(0.953)	-2.150	(0.972)	-1.028	(1.000)	-2.087	(0.983)
Ga	-4.063	(0.998)	-4.790	(1.000)	-2.158	(1.000)	-3.867	(1.000)
Pt	**-7.424**	**(0.058)**	**-8.633**	**(0.093)**	-5.703	(0.522)	-8.039	(0.245)
Pa	-8.717	(0.283)	-9.223	(0.877)	-5.300	(0.839)	-7.006	(0.979)

Note. Specifications 1 and 2 include the logs of PR, GKPK, PK and LK and constant; and the logs of PR, GKPK, PK and LK and constant plus trend (as per Equation 8). Specifications 3 and 4 include the logs of PR, GFCEPK, PK and LK and constant; and the logs of PR, GFCEPK, PK and LK and constant plus trend (as per Equation 9). Statistically significant results (indicating cointegration) are highlighted in bold. p-values are indicated in parentheses.

Table 6 - Estimates of long- and short-run relationships

Specification 1 (PR as dependent variable)					Specification 2 (PR as dependent variable)				
Variable	FMOLS	DOLS	DCCE	PMG	Variable	FMOLS	DOLS	DCCE	PMG
GKPK	1.460	2.934	4.502	1.613	GFCEPK	1.734	0.260	0.273	0.898
	(0.087)	(0.028)	(0.166)	(0.000)		(0.000)	(0.798)	(0.732)	(0.026)
LK	0.926	1.065	-0.960	-2.956	LK	-1.048	-3.397	-0.471	-2.921
	(0.034)	(0.109)	(0.343)	(0.000)		(0.022)	(0.000)	(0.598)	(0.000)
PK	0.485	0.718	1.264	-0.516	PK	-0.234	-0.450	1.175	-0.670
	(0.001)	(0.001)	(0.027)	(0.003)		(0.108)	(0.095)	(0.016)	(0.003)
d(GKPK)				1.051	d(GFCEPK)				-1.814
				(0.425)					(0.136)
d(LK)				2.474	d(LK)				3.105
				(0.008)					(0.013)
d(PK)				1.708	d(PK)				1.720
				(0.000)					(0.000)
CD			-1.520		CD			-0.480	
			(0.123)					(0.632)	
Error correction				-0.301	Error correction				-0.291
				(0.000)					(0.000)
Hausman				0.660	Hausman				4.560
				(0.883)					(0.207)
Observations	261	243	270	261	Observations	261	243	270	261

Note. All variables are expressed in logarithms. p-values are indicated in parentheses.

Finally, given certain evidence against cointegration, panel VAR estimates were provided. Due to a strong evidence of stationarity in first differences of all variables (including capital variables) and in line with other public capital studies (Pereira and Roca-Sagales 2001: 377; Pereira and Roca-Sagales 2003: 243), panel VAR was estimated in first differences of log-levels.

Two panel VAR models (both having order of one, as per Schwarz Bayesian Information Criterion) were estimated for the following vectors of endogenous series: $[LNGKPK_{it}; LNPK_{it}; LNLK_{it}; LNPR_{it}]$ to examine the effects of public to private capital ratio on profit rate (Model 1), and $[LNGFCEPK_{it}; LNPK_{it}; LNLK_{it}; LNPR_{it}]$ to examine the effect of the ratio of government final consumption expenditure to private capital (Model 2). Both models were stable, with no roots of characteristic polynomial falling outside the unit circle (roots and moduli in Table 7). Panel VAR model has atheoretical nature with no a priori scheme that indicates how variables affect each other, hence, the interpretation of PVAR coefficients (Table 7) is complemented by the analysis of IRFs (Figures 2 and 3) and FEVD (Table 8).

The results of PVAR(1) for Model 1 are presented in Table 7 and Figure 2. In the profit rate equation, the rate responded positively to its own lag and to the first lag of the public to private capital ratio and responded negatively to the first lag of capital intensity inverse and capacity-adjusted private capital. All coefficients in the equation were significant, as indicated by t-statistics. Impulse-response functions, as a more direct means of results interpretation, indicate that the logarithm of profit rate responded positively to its own shock, the effect attenuating after two periods. The signs of the effects of the capacity-adjusted private capital and the inverse of capital intensity on profit rate switched from positive to negative after one period, eventually converging to their stationary states (in periods seven and nine respectively). The positive and significant effect of the ratio of public to private capital on profitability was relatively small and limited to period two, quickly converging to zero by period four. For relative contribution of regressors to profit rate variability (Table 8), the fluctuations in profit rate were principally due to its own shocks (profit rate explained 60.69% of own variation, eight periods ahead), while the capacity-adjusted public capital and the inverse of the capital intensity explained a larger percentage of profit rate variation than public to private capital ratio (25.12% and 11.41% versus 2.77% respectively).

The delays in the positive influence of public capital on private variables (in this paper on private profitability) can be attributed to implementation lags (due to the limits in authorised spending per period) and time-to-build effects, i.e. the gaps that exist between the time the investment in physical capital is made by the government and the time when public capital stock is ready for use (Leeper et al. 2010; Atolia et al 2017: 10; Elekdag and Muir 2014: 19). The moderate size of the public capital effect can in turn be related to the composition of public capital and specifically to those public investments that do not stimulate private sector (e.g. defence procurement) or that are relatively less productive (Al-Faris 2002), as well as to potential inefficiency of public investment management institutions (Crivelli 2017).

The results of PVAR(1) for Model 2 are in most respects similar to those for Model 1 (Table 7 and Figure 3). In the profit rate equation, the coefficient of the lagged profit rate was positive and significant, while the coefficient of the lagged inverse of capital intensity and of the lagged capacity-adjusted private capital were negative (significant in both cases). The lagged ratio of government final consumption expenditure to private capital was negative but insignificant. The IRFs confirm these findings: the effects of the capacity-adjusted private capital and the inverse of capital intensity on profit rate change the signs from positive to negative and converge after six and seven periods respectively. The influence of $LNGFCEPK_{it}$ is negative in the first four periods and convergence to the steady state happens by period five. Eight periods ahead, profit rate explained 60.11% of own shocks, while $LNPK_{it}$, $LNLK_{it}$ and $LNGFCEPK_{it}$ explained 18.15%, 19.60% and 2.15% of profit rate variance respectively (Table 8).

As part of robustness checks, the panel VARs were run with alternative lag orders (two, as suggested by Hannan-Quinn Information Criterion, and three, as suggested by Akaike Information Criterion). Additionally, trivarite or bivariate VAR were estimated (without either $LNPK_{it}$ or $LNLK_{it}$ or both), and the sample was curtailed, with the US and/or Denmark removed (in the former case, due to the relative size of the US economy, in the latter due to the higher levels of $GKPK$ and $GFCEPK$). Finally, alternative orderings were tried, as long as profit rate remains the most endogenous variable in the identification scheme. While the size of the effects in these alternative estimates differed to a certain extent from those in Tables 7 and 8 and Figures 2 and 3, the signs of the effects and the general shapes of IRFs were similar. (The results of these checks are not reported here to conserve space.) Overall, the hypothesis of positive effects of public capital and insignificant and negative effects of overall government expenditure is confirmed.

Table 7 - Panel VAR estimates

Model 1				
Response of	Response to			
	$LNGKPK_{(t-1)}$	$LNPK_{(t-1)}$	$LNLK_{(t-1)}$	$LNPR_{(t-1)}$
$LNGKPK_{(t)}$	0.837	-0.013	-0.098	-0.002
	(22.219)	(-3.046)	(-3.990)	(-1.032)
$LNPK_{(t)}$	1.055	0.181	0.579	0.079
	1.729)	(2.546)]	(1.460)	2.021)
$LNLK_{(t)}$	0.288	-0.012	0.651	0.019
	(3.863)	(-1.375)	(13.432)	(4.055)
$LNPR_{(t)}$	3.074	-0.424	-1.350	0.248
	(2.974)	(-3.509)	(-2.008)	(3.753)
Root	Modulus			
0.680 - 0.169i	0.701			
0.680 + 0.169i	0.701			
0.278 - 0.137i	0.310			
0.278 + 0.137i	0.310			
No. observations	252			
No. groups	9			
Model 2				
Response of	Response to			
	$LNGFCEPK_{(t-1)}$	$LNPK_{(t-1)}$	$LNLK_{(t-1)}$	$LNPR_{(t-1)}$
$LNGFCEPK_{(t)}$	0.354	-0.007	0.101	0.011
	(5.663)	(-0.233)	(1.406)	(1.568)
$LNPK_{(t)}$	-0.141	0.747	0.143	0.018
	(-1.631)	(18.388)	(1.448)	(1.862)
$LNLK_{(t)}$	0.032	-0.090	0.638	0.021
	(0.815)	(-4.899)	(14.288)	(4.921)
$LNPR_{(t)}$	-0.950	-1.308	-1.749	0.210
	(-1.753)	(-5.134)	(-2.818)	(3.468)
Root	Modulus			
0.657	0.657			
0.497	0.497			
0.398 - 0.099i	0.410			
0.398 + 0.099i	0.410			
No. observations	252			
No. groups	9			

Note. t-statistics are indicated in parentheses.

Figure 2 - Impulse-response functions (Model 1)

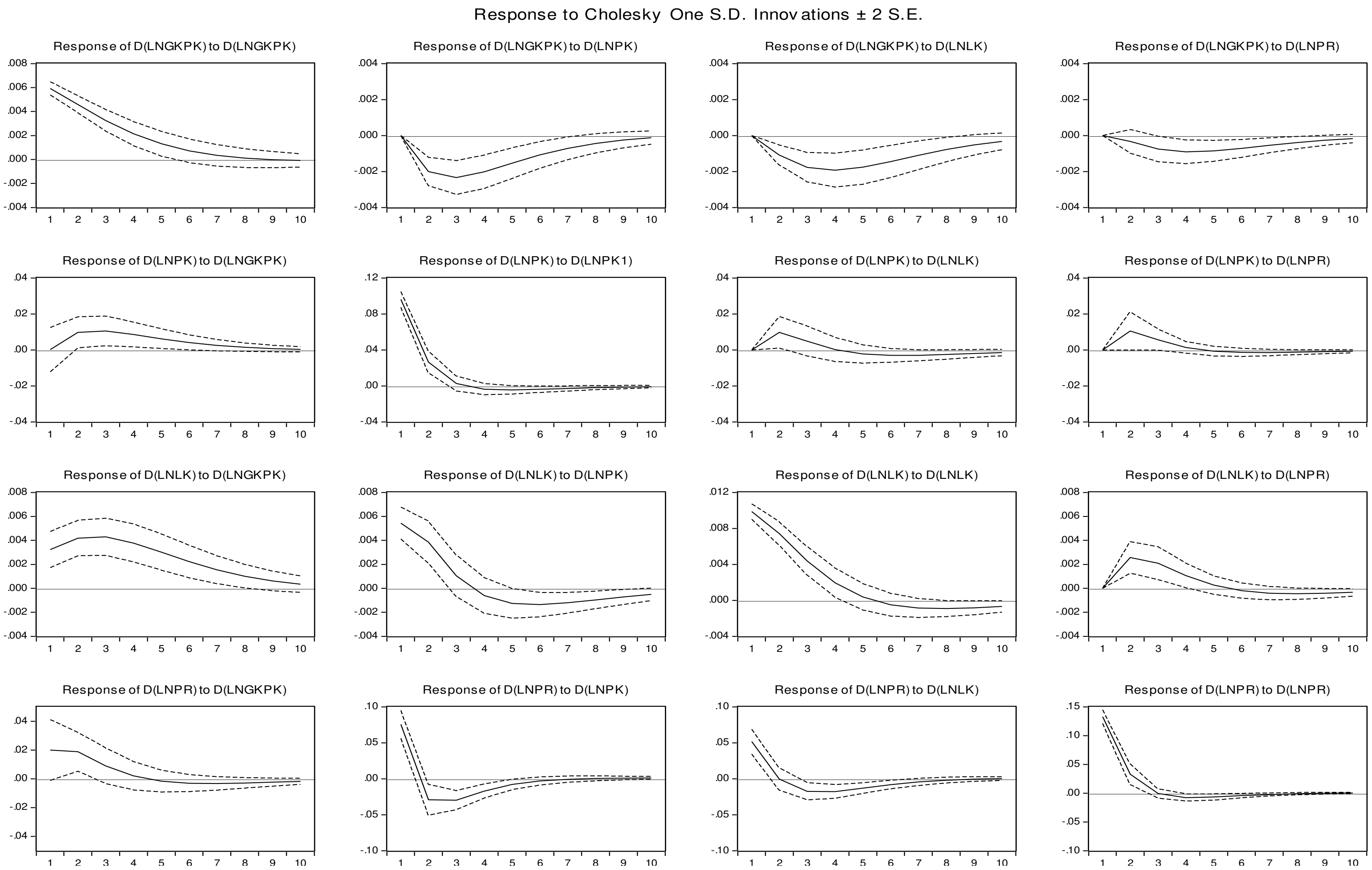


Figure 3 - Impulse-response functions (Model 2)

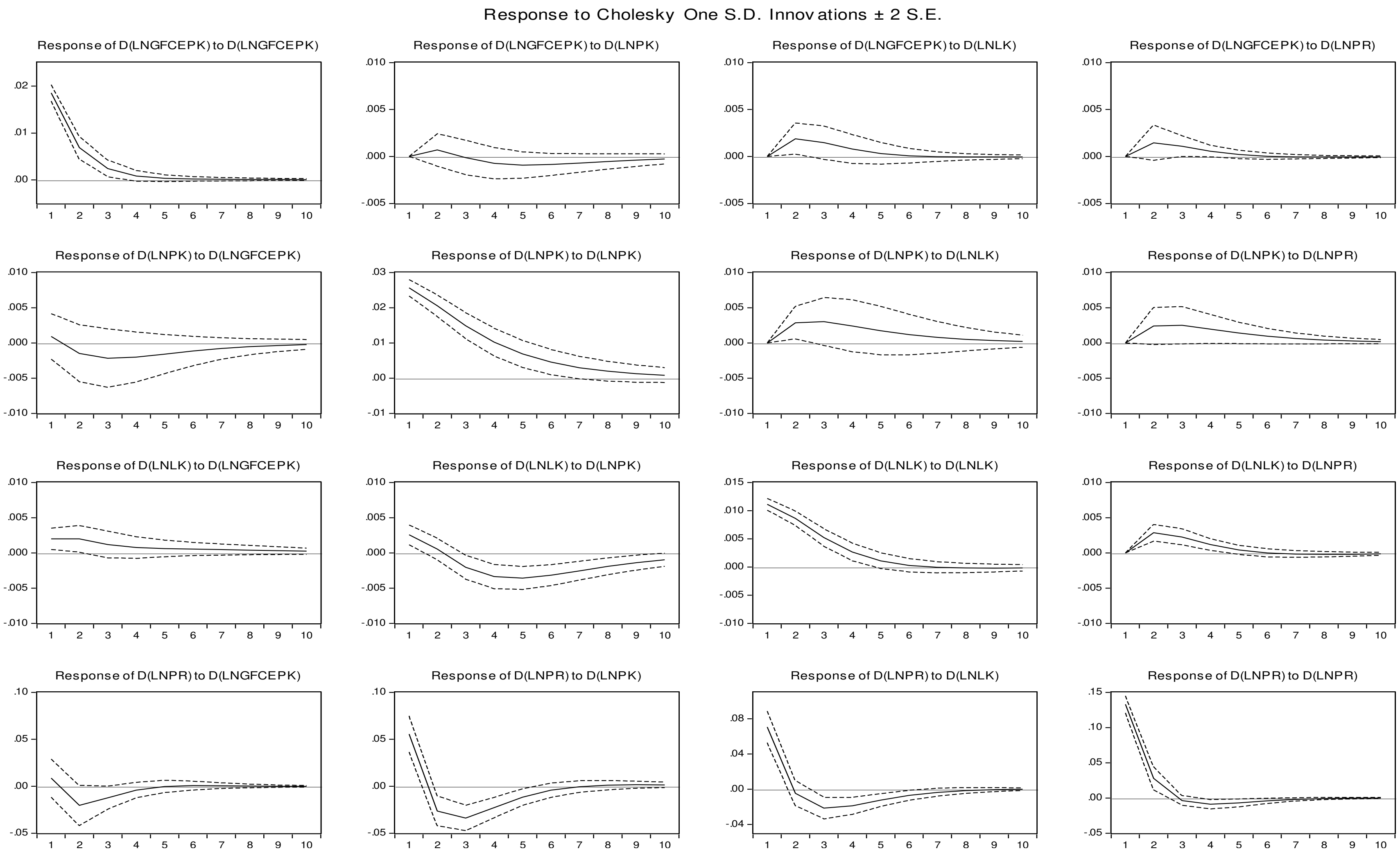


Table 8 - Forecast error variance decompositions

Model 1								
Periods ahead	Responses: impulses							
	LNGKPK:LNGKP	LNGKPK:LNPK	LNGKPK:LNL	LNGKPK:LNPR	LNPK:LNGKP	LNPK:LNPK	LNPK:LNLK	LNPK:LNPR
2	91.343	6.539	1.940	0.178	0.923	97.064	0.938	1.076
4	75.558	14.348	8.515	1.579	2.648	94.839	1.142	1.372
6	68.991	16.018	12.414	2.577	3.138	94.211	1.270	1.381
8	67.145	16.203	13.725	2.927	3.207	93.969	1.412	1.412
Periods ahead	Responses: impulses							
	LNLK:LNGKPK	LNLK:LNPK	LNLK:LNLK	LNLK:LNPR	LNPR:LNGKPK	LNPR:LNPK	LNPR:LNLK	LNPR:LNPR
2	12.065	19.192	65.890	2.853	2.597	22.833	9.271	65.299
4	20.606	15.599	59.684	4.111	2.708	25.236	10.726	61.330
6	23.914	15.804	56.372	3.910	2.714	25.161	11.354	60.771
8	24.399	16.174	55.498	3.929	2.772	25.121	11.414	60.693
Model 2								
Periods ahead	Responses: impulses							
	LNGFCEPK:LNG	LNGFCEPK:LN	LNGFCEPK:LN	LNGFCEPK:LNP	LNPK:LNGFCE	LNPK:LNPK	LNPK:LNLK	LNPK:LNPR
2	98.445	0.120	0.902	0.533	0.279	98.450	0.750	0.522
4	97.309	0.250	1.555	0.885	0.811	96.507	1.596	1.086
6	96.900	0.634	1.574	0.891	1.022	95.962	1.800	1.217
8	96.716	0.821	1.572	0.891	1.073	95.847	1.839	1.241
Periods ahead	Responses: impulses							
	LNLK:LNGFCEP	LNLK:LNPK	LNLK:LNLK	LNLK:LNPR	LNPR:LNGFCE	LNPR:LNPK	LNPR:LNLK	LNPR:LNPR
2	3.560	3.066	89.688	3.686	1.785	13.669	18.092	66.455
4	3.512	8.047	83.214	5.227	2.175	17.911	19.161	60.753
6	3.434	14.959	76.760	4.847	2.147	18.154	19.564	60.135
8	3.431	17.710	74.151	4.708	2.146	18.146	19.600	60.108

5. Conclusion

The paper explored the impact of public capital and government final consumption expenditure (expressed as a proportion of private capital) on the profitability of the productive economies of the OECD in 1977-2006. The profit rate of a private economy is considered a good approximation of the private sector profit rate (data not available in the relevant databases for all the economies in question), given that, in these economies, private enterprises own most of the means of production and perform most of the economic activity in the productive sectors. The empirical strategy was to apply the panel cointegration models and tests to establish long-run relationships and use the panel VAR model to capture short-run dynamics, in addition to the usual robustness checks. The findings across the models and various specifications were consistent and non-contradictory.

The public capital provided by the government in the form of tangible physical capital (such as public infrastructure) positively stimulated the profitability of the productive economy, while overall government expenditure (final consumption expenditure) had negative and generally insignificant consequences. The study therefore confirms the early insight of Aschauer (1988, 1989) that the type of government expenditure is non-neutral with respect to economic outcomes (growth, productivity, and profitability).

This paper did not unambiguously prove the presence of cointegration (a long-run equilibrium relationship) among the variables in levels: certain evidence of cointegration was indicated by Pedroni and Kao tests, but not by Westerlund test. The positive effect of public capital on private profits did not appear large, given a more moderate restoration of public investment in the 1980-90s than restoration of profitability over the same period. The panel VAR model (that captures short-run dynamics and does not assume a long-run equilibrium relationship between the variables) demonstrated, in a similar vein, the positive shocks of public capital on profitability. The magnitude of these shocks was likewise moderate and short-lived.

Future research should consider a longer sample and greater number of economies, examine the public capital effects on profitability in other contexts (e.g., developing economies), and estimate alternative profitability indicators (private or corporate rates). It could also consider public capital alongside other variables or processes that contribute to rising profitability, such as a decline in the bargaining power of labour, export of capital in the form of foreign direct investment, trade liberalisation, financialisation and the growth of indebtedness, and technological changes that revive capital productivity (Toporowski 1999; Baragar and Chernomas 2012; Basu and Vasudevan 2013).

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