Finance Research Letters 58 (2023) 104317

Contents lists available at ScienceDirect

Finance Research Letters

journal homepage: www.elsevier.com/locate/frl

Public and private investments: Long-run asymmetric effects in France and the US

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ARTICLE INFO

JEL classification:
E62
H72
O47

Keywords:
Asymmetric effects
Cointegration
NARDL
Public investments
Private investments

ABSTRACT

We analyse whether private investments are impacted asymmetrically by public investments and how any asymmetry evolves over time. We conduct time series analyses for France and the US within a flexible empirical framework exploiting quarterly data over 1960Q1–2022Q4. The results are summarized as follows: France's private investments are positively impacted by public investments both in the short- and long-run; public investments have a neutral or negative effect on private investments in the US; the asymmetric effects in France become significant and persistent after 10 quarters; some evidence of asymmetry emerges for the US in the long-run.

1. Introduction

Recent economic shocks have been pushing governments worldwide to stimulate growth through unprecedented expansionary measures within environments characterized by high inflation and rising interest rates. The most relevant recent economic stimuli are the US's American Rescue Plan Act of 2021 and the EU's Next Generation Fund. A large share of such financial resources is constrained to the implementation of investment programmes. Moreover, as reported by the IMF's World Economic Outlook (April 2023 Edition), the GDP shares of government expenditure for the US and the EU are expected to remain above pre-pandemic levels at least until 2028. It is therefore pivotal to look at the possible implications for the private sector.

A high degree of complementarity between public and private capital emerged in the seminal works of Aschauer (1989a, 1989b). Nevertheless, the empirical evidence is quite diversified – although crowding-in effects prevail. Recent evidence at the panel level suggests that public investment is a key determinant of private investments (e.g. Marattin and Salotti, 2011; Abiad et al., 2016; Carvelli, 2023). Hence, there are economic and statistical aspects of the phenomenon that are worth further analysis.

We depart from the existing literature by investigating whether private investments respond asymmetrically to positive and negative changes in public investment, and how such asymmetries evolve over time. Any evidence of asymmetric effects would have relevant policy implications as the current massive government purchasing programmes will likely soon be followed by a reduction in the flow of public capital.[1] Using quarterly data from the OECD Economic Outlook database (Edition 2022/2), we implement novel time series techniques to accommodate non-stationarity, asymmetric cointegration and fiscal feedback effects. Such quarterly data are only available for 14 OECD countries, and we focus on France and the US for the following reasons. Firstly, amongst the EU countries

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[1] As stated in the IMF Fiscal Monitor (April 2023), governments are expected to gradually reduce their fiscal stimuli.

https://doi.org/10.1016/j.frl.2023.104317
Received 6 June 2023; Received in revised form 26 July 2023; Accepted 10 August 2023
Available online 18 August 2023
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whose quarterly data are available, France benefits from the highest share of Next Generation EU resources. Secondly, since 2020 the US has been heavily resorting to public expenditure and is the most economically relevant country in the above-mentioned database. Related studies for France and the US – which have mostly employed VAR techniques – provide mixed results (e.g. Voss, 2002; Afonso and Aubyn, 2009, 2010, 2019; Creel et al., 2016), leaving the question open to additional evidence.

The rest of the paper is organized as follows. Section 2 describes the data and implements preliminary tests; Section 3 sketches out the econometric strategy; Section 4 discusses the results; Section 5 performs a set of robustness checks; Section 6 concludes the articles.

2. Data and unit root tests

We employ quarterly data from the OECD Economic Outlook database (Edition 2022/2) spanning the 1960Q1–2022Q4 period. As shown in Table 1, the GDP shares of public and private investments in both countries were similar over the last two decades.

As indicated by Fig. 1, the series appear persistent in levels. We, therefore, implement unit root tests to assess the integration orders of the variables. The main procedure is based on the augmented Dickey–Fuller (ADF) test. In addition, we employ the Phililps–Perron (PP) test to tackle autocorrelation and heteroskedasticity more effectively. Moreover, since our T is large, we implement the Zivot-Andrews unit root test as it accommodates any endogeneous structural break. The outcomes of the tests, reported in Tables A1-A3, suggest that the series are non-stationary in levels and stationary in first differences for both France and the US.

3. Econometric strategy

Let us consider the following long-run asymmetric equation:

$$y_t = \lambda^+ x_t^+ + \lambda^- x_t^- + e_t \tag{1}$$

for $t = 1960Q1, \ldots, 2022Q4$. y denotes private investments and $x^+$and x^- are the positive and negative changes in public investments, respectively. All variables are expressed as a share of GDP. The parameters λ^+ and λ^- measure the asymmetric responses of private investments to public capital formation. The term e represents the innovations in the model.

Following Schorderet (2003), we construct positive and negative variations of the explanatory variable through partial sum decomposition around a null threshold in order to separate expansions and contractions in the flow of public capital:

$$x_t^+ = \sum_{j=1}^{t} \Delta x_j^+ = \sum_{j=1}^{t} \max\left(\Delta x_j, 0\right), \; x_t^- = \sum_{j=1}^{t} \Delta x_j^- = \sum_{j=1}^{t} \min\left(\Delta x_j, 0\right). \tag{2}$$

Since the variables are non-stationary in levels, estimating Eq. (1) would lead to biased results. Considering the economic and statistical features of the phenomenon, the Non-Linear ARDL (NARDL) model proposed by Shin et al. (2014) appears to be the most suitable technique to identify the coefficients as it allows for non-stationarity, asymmetric effects and past feedback of the observables. We can give the following representation of the NARDL model:

$$y_t = \sum_{j=1}^{p} \gamma_j y_{t-j} + \sum_{j=0}^{q} \left(\lambda_j^+ x_{t-j}^+ + \lambda_j^- x_{t-j}^-\right) + e_t \tag{3}$$

where the term γ_j, for $j = 1, \ldots, p$, is the autoregressive parameter. The lag lengths p and q are selected according to the Akaike information criterion (AIC). Since NARDL only requires that the integration order of the variables is strictly lower than two, Eq. (3) can be consistently estimated. However, we aim to distinguish short- and long-run effects as well as ascertain whether and how the system convergences to an equilibrium path. The latter aspect is crucial since, as discussed in depth in Hatano (2010) and Dreger and Reimers (2016), the theoretical mechanisms justifying a cointegrating relationship between private and public capital are solid.[2] Therefore, we reparametrize Eq. (3) into its ECM version:

$$\Delta y_t = \sum_{j=1}^{p-1} \rho_j \Delta y_{t-j} + \sum_{j=0}^{q-1} \left(\beta_j^+ \Delta x_{t-j}^+ + \beta_j^- \Delta x_{t-j}^-\right) + \theta^+ x_{t-1}^+ + \theta^- x_{t-1}^- + \xi y_{t-1} + v_t \tag{4}$$

where the speed of convergence to the equilibrium path is given by ξ. The autoregressive parameter is now ρ_j, for $j = 1, \ldots, p-1$. The short-run asymmetric coefficients are β^+ and β^-, whereas their long-run counterparts are defined as $\lambda^+ = -\theta^+/\xi$ and $\lambda^- = -\theta^-/\xi$. The error term v is assumed to follow an *iid* process with null mean and constant variance.

The existence of short- and long-run asymmetries can be verified through the Wald test on linear restrictions, based on the null hypothesis of no asymmetric effects. The ECM model specified in Eq. (4) differs from those previously implemented in related studies, as these implicitly assume symmetric effects. However, if the data-generating process is characterised by asymmetries, neglecting them could lead to an omitted-variable bias (Granger and Yoon, 2002).

A typical econometric issue when estimating the macroeconomic effects of fiscal policy is endogeneity, as government expenditure

[2] Recent evidence of cointegration can also be found at the empirical level (e.g. Matvejevs and Tkacevs, 2023; Monastiriotis and Randjelovic, 2023).

Table 1
Private and public investments, average values as a share of GDP.

	France				USA			
	1960–80	1981–2000	2001–2022	1960–2022	1960–80	1981–2000	2001–2022	1960–2022
Private investments	-	0.094	0.126	0.112	0.068	0.092	0.123	0.095
Public investments	0.058	0.038	0.038	0.044	0.049	0.040	0.037	0.042

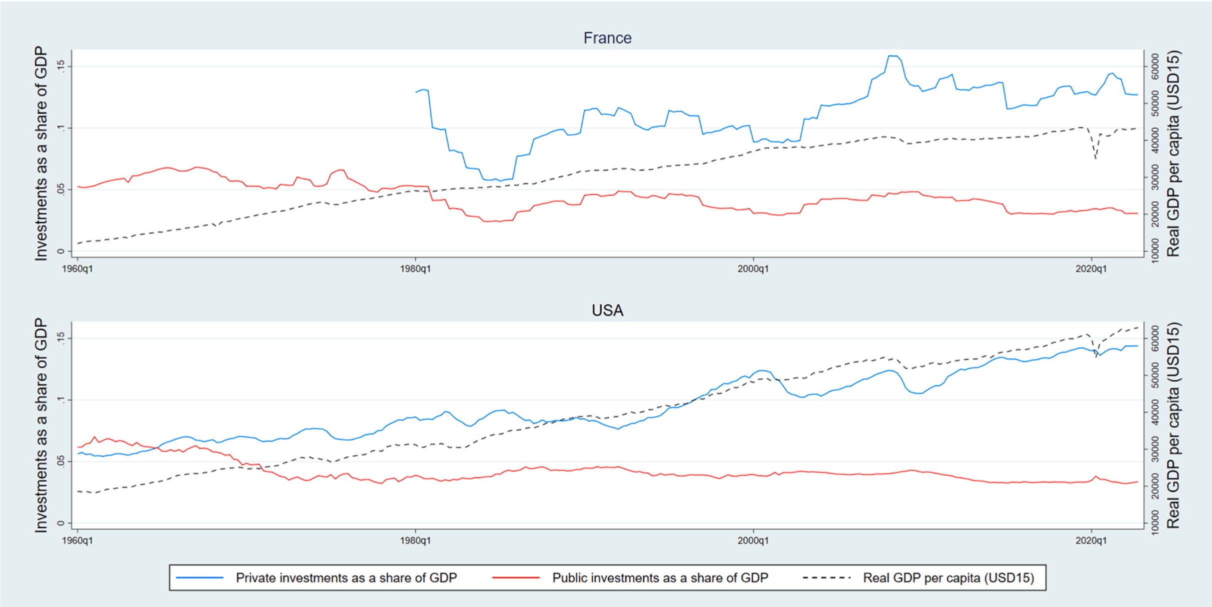

Fig. 1. Private investments, public investments and output 1960Q1–2022Q4. Authors' elaboration on OECD Economic Outlook data.

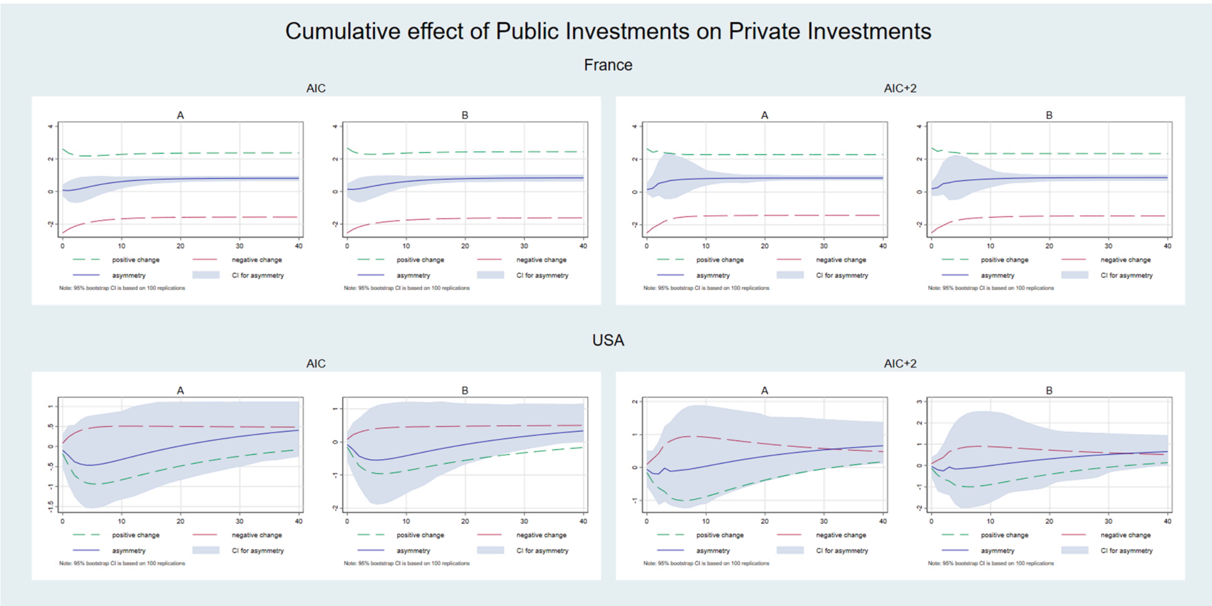

Fig. 2. Dynamic multipliers. The 95% confidence intervals for asymmetry are built through bootstrap procedures based on 100 replications. Model A does not include additional controls. Model B controls for the GDP growth rate.

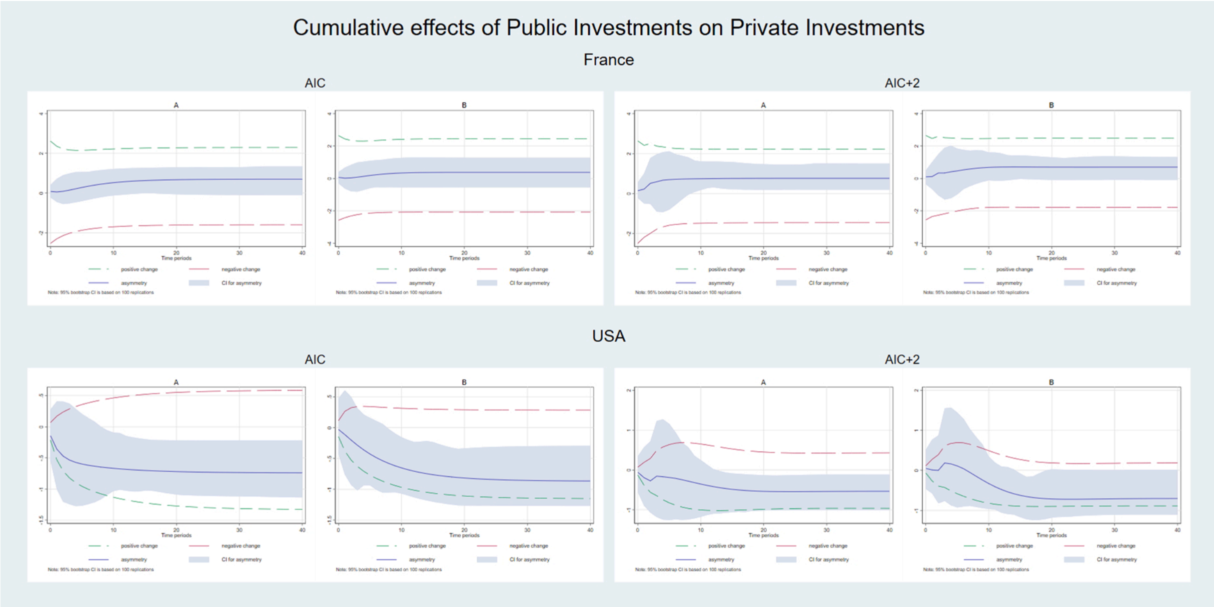

Fig. 3. Dynamic multipliers with additional controls. The 95% confidence intervals for asymmetry are built through bootstrap procedures based on 100 replications. Model A controls for GDP per capita. Model B controls for the real GDP per capita, the long-term interest rate on government bonds and the openness index.

could be jointly determined with the macroeconomic environment (Fatás and Mihov, 2003). However, things change when considering measures of public spending that are net of the automatic stabilizers and of government actions whose implementation is not immediate – as in the case of public investments. Detailed discussions on the exogeneity of public investment can be found in Mittnik and Neumann (2001), Deleidi et al. (2020) and Afonso and Rodrigues (2023). Furthermore, when using quarterly data – as in the present study – the risk of a simultaneous response of public investments to macroeconomics fluctuations is even lower (e.g. Beetsma et al., 2009; Born and Muller, 2012) since fiscal measures typically require at least half a year to enter into force (Erenburg and Wohar, 1995; Blanchard and Perotti, 2002; Ramey, 2011; Kilian and Lütkepohl, 2017). In addition, the NARDL estimates are unbiased even when the regressors are weakly endogenous (Shin et al., 2014).

A significant estimated value of ξ ranging in the interval $[-1; 0]$ can be interpreted as evidence of asymmetric cointegration. However, we also perform two formal tests for cointegration for each estimated model. As suggested by Shin et al. (2014), we implement the t-test developed by Banerjee et al. (1998) and the F-test of Pesaran et al. (2001), denoted as t_{BDM} and F_{PSS}, respectively. Both the t_{BDM} and the F_{PSS} testing procedures are largely employed in the ARDL frameworks. The t_{BDM} test verifies the null hypothesis H_0: $\xi = 0$ against the alternative H_a: $\xi < 0$. The F_{PSS} is more restrictive as it tests for the joint significance of the long-run parameters and the error correction term, which in our case is H_0: $\xi = \theta^{+} = \theta^{-} = 0$. Under the null hypothesis, the asymptotic distributions of both the t_{BDM} and the F_{PSS} test statistics are nonstandard. Accordingly, the critical values – derived numerically – are functions of the number of regressors k entering the cointegrating vector.[3]

4. Results

Since T is large, we assess the sensitivity of the estimates to the lag structure by further estimating regressions in which the lags are increased by two. In addition, we estimate regressions that include the growth rate of GDP in the cointegrating vector, as the output fluctuations may significantly affect the investment decisions of the agents. To test for parameter stability, for each regression we plot the cumulative sum of recursive residuals (CUSUM) and the cumulative sum of recursive residual squares (CUSUM-Q), with the related 95% confidence interval around zero (Figs. A1-A4). The estimates are reported in Table 2.

Let us begin with the findings related to France. We note that the short- and long-run effects of public investments are positive and significant across all the specifications, aligning with Creel et al. (2016) and Afonso and Aubyn (2019). The effects are symmetric in the short-run but turn out to be asymmetric in the long run, implying that the investment decisions of private firms are more sensitive to expansions of public capital than to its reliefs. The coefficients associated with the error correction term are negative and significant across all the specifications, suggesting that private and public investments are asymmetrically cointegrated. The evidence of cointegration is strengthened by the outcomes of the t_{BDM} and the F_{PSS} tests, as both lead to the rejection of the null hypothesis of no

[3] As suggested by Shin et al. (2014), we set the value of k considering the number of regressors prior to the decomposition of the asymmetric variable to avoid the overestimation of the test. In fact, the critical values for each level of significance decrease as the value of k increases.

Table 2
NARDL estimates.

	FRANCE				USA			
	AIC		AIC+2		AIC		AIC+2	
$\widehat{\rho_1}$	0.487***	0.465***	0.465***	0.419***	0.343***	0.328***	0.333***	0.314***
	(0.0660)	(0.0625)	(0.0756)	(0.0712)	(0.0633)	(0.0644)	(0.0657)	(0.0671)
$\widehat{\beta_0^+}$	2.616***	2.673***	2.637***	2.684***	−0.176	−0.154	−0.148	−0.126
	(0.138)	(0.131)	(0.143)	(0.134)	(0.155)	(0.156)	(0.174)	(0.175)
$\widehat{\beta_0^-}$	2.531***	2.526***	2.492***	2.503***	−0.0836	−0.0797	−0.0933	−0.0903
	(0.115)	(0.109)	(0.127)	(0.118)	(0.134)	(0.134)	(0.138)	(0.138)
$\widehat{\lambda^+}$	0.168***	0.159***	0.188***	0.179***	0.00421	0.00302	0.0105	0.00979
	(0.0504)	(0.0476)	(0.0560)	(0.0523)	(0.0194)	(0.0194)	(0.0205)	(0.0205)
$\widehat{\lambda^-}$	0.110***	0.104***	0.118***	0.112***	−0.00832	−0.00872	−0.00658	−0.00723
	(0.0378)	(0.0357)	(0.0404)	(0.0377)	(0.0144)	(0.0144)	(0.0148)	(0.0148)
$\widehat{\omega}$		0.0336***		0.0356***		0.00976		0.0117
		(0.00733)		(0.00733)		(0.00835)		(0.00859)
$\widehat{\xi}$	−0.0706***	−0.0650***	−0.0828***	−0.0767***	−0.0176*	−0.0167*	−0.0197**	−0.0193**
	(0.0191)	(0.0181)	(0.0225)	(0.0211)	(0.00916)	(0.00919)	(0.00962)	(0.00961)
T	172	172	170	170	251	251	249	249
SR Wald p-value	0.919	0.933	0.702	0.711	0.445	0.379	0.794	0.763
LR Wald p-value	0.000	0.000	0.000	0.000	0.030	0.042	0.008	0.008
t_{BDM} stat.	−3.6936	−3.5931	−3.6790	−3.6443	−1.4917	−1.4029	−2.2154	−2.1499
F_{PSS} stat.	4.6727	4.5891	4.5666	4.5889	1.3117	1.1813	1.4338	1.3900
CUSUM stat.	0.8188	0.9297	0.7785	0.9161	0.8509	0.6445	0.7560	0.5784
Port. p-val.	0.1224	0.1935	0.1294	0.3033	0.4369	0.4388	0.7238	0.7858
BP p-val.	0.0825	0.0640	0.0901	0.0659	0.0033	0.0048	0.0041	0.0055
RESET p-val.	0.7278	0.7918	0.6398	0.5473	0.1812	0.2628	0.0991	0.1915
JB p-val.	0.1202	0.2587	0.0981	0.2613	0.0421	0.6070	0.5918	0.7011
Adj. R^2	0.8721	0.8861	0.8748	0.8908	0.2455	0.2466	0.2399	0.2427
RMSE	0.00197	0.00186	0.00197	0.00184	0.00146	0.00146	0.00147	0.00147

Notes. The asterisks ***, **, and * denote significance at the 1%, 5%, and 10% level, respectively. The standard errors are reported in square brackets. ω is the long-run coefficient associated with GDP growth. SR (LR) Wald test verifies the null hypothesis $\widehat{\beta_0^+} = \widehat{\beta_0^-}$ ($\widehat{\lambda^+} = \widehat{\lambda^-}$). Under the null hypothesis of no cointegration, the lower bound (upper bound) critical values at 10%, 5%, and 1% of the t_{BDM} test are the following: *i)* for $k = 2$, −3.13 (−3.63) −3.65 (−4.20) −3.96 (−4.53); *ii)* for $k = 3$, −3.13 (−3.84) −3.41 (−4.16) −3.96 (−4.73). Under the null hypothesis of no cointegration, the lower bound (upper bound) critical values at 10%, 5% and 1% of the F_{PSS} test are the following: *i)* for $k = 2$, 4.19 (4.45), 4.87 (5.85) and 6.34 (7.52) *ii)* for $k = 3$, 3.47 (4.45), 4.01 (5.07) and 5.17 (6.36). CUSUM test is constructed on the null hypothesis of no structural breaks, and the related critical values at 10%, 5%, and 1% are, respectively, 0.8499, 0.9479, and 1.1430. The p-values of the following tests are reported: Portmanteau test (Port); Breusch/Pagan heteroskedasticity test (BP); Ramsey test (RESET); Jarque-Bera test on normality (JB). RMSE is the root mean squared error.

cointegration. No evidence of parameter instability arises – as suggested by the CUSUM and CUSUM-Q tests on structural breaks and the recursive cumulated sum of residuals plotted in Fig. A1 and Fig. A2, respectively. The dynamic multipliers over 40 quarters (Fig. 2) highlight that the effects are permanent over the whole period and that the impact of positive changes in public investment is always greater in absolute value than the impact of negative variation. The asymmetry quickly becomes positive, and the related bandwidth gets narrower over time. The existence of permanent effects of changes in public investment underlines the relevance of policy decisions for the private sector and, therefore, for the macroeconomy as a whole, as the effects of the variation in the flow of public capital are not offset by economic adjustments, neither in the short- nor in the long-run.

The results for the US are substantially different. The main contrasting outcome with France lies in the sign of the estimated coefficients associated with increases (decreases) in public investment, as these are negative (positive) both in the short- and the long-run. Nevertheless, the estimates are statistically insignificant, suggesting that public investments have neither crowding-in nor crowding-out effects on private capital formation as its impact is rather neutral – partially aligning with Erenburg and Wohar (1995). While the effects are symmetric in the short-run, there is some evidence of asymmetry in the long-run, with the bandwidth associated with the dynamic multipliers of the asymmetric effects becoming smaller as the time horizon increases. In addition, the asymmetries are negative in the first periods and become slightly positive after around 5 quarters. However, the interpretation of any evidence of significant long-run asymmetries could be misleading in the presence of insignificant coefficients. Moreover, the evidence of cointegration appears to be weak as the t_{BDM} and the F_{PSS} tests fail to reject the null hypothesis of no cointegration. However, further factors might significantly affect the dynamics of private investments. We address such an issue in the next section.

The different outcomes for France and the US reflect a high degree of cross-country heterogeneity related to the macroeconomic effects of fiscal policy, consistent with Afonso and Sousa (2012) and Carvelli (2023). The main explanations for the contrasting results for the two economies considered in this study could be a greater market efficiency in the US and/or a more prominent efficiency in the allocation of public resources in France. In addition, the historical dissimilarity in terms of public finance structure between the two countries, as well as the ways private firms borrow financial resources, might translate into different macroeconomic effects of fiscal policies. As suggested by Mittnik and Neumann (2001) and Erden and Holcombe (2005), the economic and financial channels behind the crowding-in and crowding-out effects may coexist with different degrees of intensity, and thus the mechanism that prevails drives the sign of the impact. The short- and long-run positive effects of public investment in France could be due, to a large extent, to demand effects and enhancement in the marginal productivity of private capital, as the latter would incentivise private firms to increase their stock of capital. If such mechanisms do characterise the data-generating process, it means that in the US the expansionary channels of public investments are offset by distortions related to increased competition for inputs or higher (current or expected) taxation. Given the complexity of the phenomenon under analysis and its implications at both the economic and financial levels, the existence of additional channels linking public investments to private capital formation is plausible. Whether and how these mechanisms are determinants for the relationship under study and its heterogeneous cross-country behaviour represents a useful question to be addressed in future research.

5. Additional robustness checks

In this section, we conduct additional empirical exercises related to relevant inferential and economic issues to assess the estimates' robustness.

As a first step, we augment the baseline models with further covariates to rule out the risk of model misspecification. However, considering that we have built an ECM equation in which the effects of public investments are allowed to be asymmetric both in the short- and in the long-run, we attempt to balance the bias-variance trade-off carefully, therefore maintaining a parsimonious approach in the choice of the number of controls. Following the recent related empirical literature (e.g. Ashraf and Herzer, 2014; Afonso and Aubyn, 2019; Carvelli, 2023), we consider as relevant determinants of the long-run dynamics of private investments the real GDP per capita, the long-term real interest rate on government bonds and the openness index. Unlike the GDP growth rate, the output level is more suitable for establishing a long-run equilibrium and accommodating any differences in capital profitability when the economies experience new steady-state levels. The long-term real interest rate on government bonds proxies financial conditions, as well as the current and expected cost of capital and its profitability. The openness index – defined as the ratio of imports and exports to GDP – accommodates the interlinkages between private investments and the international economic environment (see, for instance, Levine and Renelt (1992) and Servén (2003)). The covariates mentioned above entering the cointegrating vector are nonstationary in levels and stationary in first differences.[4] The related estimates and dynamic multipliers are reported in Table 3 and Fig. 3, respectively.

As it concerns France, the results closely align with the ones discussed in the previous section in terms of magnitude and statistical significance, except for a slight increase in the speed of adjustment to the long-run equilibrium. The dynamic multipliers show that the way the asymmetric effects evolve from the short- to the long-run are almost identical to those reported in the previous section, except for the enlargement in the confidence intervals due to the inclusion of additional controls.

Unlike the case of France, controlling for output levels, interest rate and international openness improves the model's overall performance for the US and exerts some differences in terms of long-run effects and cointegration. In fact, while the results on the short-run effects and the significance of the asymmetries remain almost unchanged, the long-run crowding-out effects turn out to be statistically significant. Moreover, the t_{BDM} and the F_{PSS} cointegration tests become significant across all the specifications, thus providing robust evidence against the null hypothesis of no cointegration. Consistently, the dynamic multipliers plotted in Fig. 3 show that the confidence intervals on asymmetry are narrower compared to the ones plotted in the previous section – although the way the effects and the asymmetries evolve from the short- to the long-run are closely aligned to those reported in Fig. 2. Considering the statistical significance of the long-run parameters, the speed of adjustment term and the Wald test on linear restrictions, we can argue that the evidence of asymmetric cointegration also emerges for the US, provided that the model is conditioned to output levels, interest rate and international openness.

The joint interpretation of the estimates for France and the US provides some important econometric insights. While for the US the model's overall performance improves when output levels, interest rate and openness index are controlled for, the opposite arises for France. In addition to the fiscal heterogeneous effects – as emerged in the existing literature and this study – it appears likely that the phenomenon differs across the countries also in terms of data-generating processes. This would imply that when modelling the empirical equation, one should consider that the population functional form linking public investments to private capital formation might differ across the countries – at least as it concerns the factors that need to be controlled for.

As a second step, we have carefully considered the direction of causality to assess whether more than one cointegrating relationship exists between the two key variables. While both the t_{BDM} and F_{PSS} tests provide strong evidence of causal (asymmetric) effects going from public investments to private capital formation, any cointegrating relationship in the opposite direction would challenge the estimates previously obtained. Therefore, we have estimated the models (with all the sets of controls considered in this article) by interchanging the dependent and the independent variables. The related estimates provide robust evidence of a lack of causal effects going from private capital formation to public investments. Further confirmation is provided by the outcomes of the Vector Error Correction (VECM), as it highlights the existence of only one cointegrating relationship. Such results are consistent with a stream of the literature that we have referred to in Section 3 to justify the assumption of orthogonality of public investments to the private capital formation (e.g. Mittnik and Neumann, 2001; Beetsma et al., 2009; Born and Muller, 2012; Deleidi et al., 2020; Afonso and Rodrigues, 2023). As a final step, we have estimated the models by considering the fiscal variables in absolute terms (USD 2015) in place of their ratio to GDP. The related outcomes do not lead to different conclusions.[5]

[4] The outcomes of the unit root tests for the control variables are not reported in the manuscript to save space but are available upon request.

[5] The results of the second and third steps discussed in this section, not reported in order to save space, are available upon request.

Table 3
NARDL estimates with additional controls.

	FRANCE				USA			
	AIC		AIC+2		AIC		AIC+2	
$\widehat{\rho_1}$	0.485***	0.476***	0.463***	0.450***	0.418***	0.403***	0.323***	0.319***
	(0.0667)	(0.0653)	(0.0762)	(0.0745)	(0.0571)	(0.0572)	(0.0641)	(0.0640)
$\widehat{\beta_0^+}$	2.613***	2.666***	2.634***	2.660***	−0.118	−0.0756	−0.119	−0.0780
	(0.139)	(0.137)	(0.144)	(0.141)	(0.170)	(0.171)	(0.170)	(0.172)
$\widehat{\beta_0^-}$	2.530***	2.564***	2.491***	2.563***	−0.0631	−0.105	−0.0496	−0.0872
	(0.116)	(0.115)	(0.127)	(0.132)	(0.139)	(0.141)	(0.138)	(0.140)
$\widehat{\lambda^+}$	0.162***	0.264***	0.183***	0.265***	−0.0831***	−0.110***	−0.0684**	−0.0937**
	(0.0542)	(0.0667)	(0.0600)	(0.0702)	(0.0290)	(0.0372)	(0.0304)	(0.0382)
$\widehat{\lambda^-}$	0.113***	0.240***	0.120***	0.237***	−0.0363**	−0.0486***	−0.0323**	−0.0424**
	(0.0393)	(0.0670)	(0.0416)	(0.0724)	(0.0157)	(0.0167)	(0.0161)	(0.0171)
$\widehat{\vartheta}$	3.69e-08	−4.58e-07**	3.23e-08	−4.88e-07**	1.96e-07***	1.64e-07	1.92e-07***	1.80e-07*
	(1.37e-07)	(2.09e-07)	(1.38e-07)	(2.23e-07)	(5.77e-08)	(1.01e-07)	(6.09e-08)	(1.01e-07)
$\widehat{\iota}$		−0.00018**		−0.00021***		8.02e-05		7.14e-05
		(7.92e-05)		(8.02e-05)		(5.45e-05)		(5.60e-05)
$\widehat{\psi}$		0.0363***		0.0376***		0.0140		0.0109
		(0.0123)		(0.0140)		(0.00899)		(0.00915)
$\widehat{\xi}$	−0.0707***	−0.107***	−0.0827***	−0.109***	−0.0613***	−0.0704***	−0.0680***	−0.0768***
	(0.0192)	(0.0246)	(0.0226)	(0.0264)	(0.0165)	(0.0172)	(0.0175)	(0.0183)
T	172	172	170	170	252	252	250	250
SR Wald p-value	0.926	0.993	0.694	0.498	0.315	0.643	0.747	0.822
LR Wald p-value	0.008	0.000	0.046	0.001	0.004	0.010	0.033	0.033
t_{BDM} stat.	−3.6887	−4.3448	−3.6639	−4.1381	−3.7350	−4.1053	−3.8899	−4.2009
F_{PSS} stat.	4.5362	6.3916	4.4860	5.7449	5.0121	6.1635	5.2550	6.1945
CUSUM stat.	0.7227	0.6502	0.7991	0.6104	0.0619	0.5832	0.0523	0.6888
Port. p-val.	0.1172	0.3480	0.1259	0.4643	0.1479	0.3851	0.3230	0.5796
BP p-val.	0.0552	0.0863	0.0907	0.0958	0.0587	0.0720	0.0611	0.0774
RESET p-val.	0.7393	0.6810	0.6527	0.7522	0.1418	0.2451	0.1140	0.1590
JB p-val.	0.2506	0.3149	0.2036	0.3204	0.1595	0.2107	0.1852	0.1988
Adj. R^2	0.8714	0.8734	0.874	0.8801	0.264	0.2719	0.2861	0.2891
RMSE	0.00198	0.00196	0.00197	0.00192	0.00145	0.00144	0.00143	0.00143

Notes. The asterisks ***, **, and * denote significance at the 1%, 5%, and 10% level, respectively. The standard errors are reported in square brackets. ϑ, ι and ψ are the long-run coefficients associated with real GDP per capita, long-term real interest rate on government bonds and openness index, respectively. SR (LR) Wald test verifies the null hypothesis $\widehat{\beta_0^+} = \widehat{\beta_0^-}$ ($\widehat{\lambda^+} = \widehat{\lambda^-}$). Under the null hypothesis of no cointegration, the lower bound (upper bound) critical values at 10%, 5%, and 1% of the t_{BDM} test are the following: *i)* for $k = 3$, −3.13 (−3.84) −3.41 (−4.16) −3.96 (−4.73); *ii)* for $k = 5$, −3.13 (−4.21) −3.41 (−4.52) −3.96 (−5.13). Under the null hypothesis of no cointegration, the lower bound (upper bound) critical values at 10%, 5% and 1% of the F_{PSS} test are the following: *i)* for $k = 3$, 3.47 (4.45), 4.01 (5.07) and 5.17 (6.36); *ii)* for $k = 5$, 2.75 (3.79), 3.12 (4.25) and 3.93 (5.23). CUSUM test is constructed on the null hypothesis of no structural breaks, and the related critical values at 10%, 5%, and 1% are, respectively, 0.8499, 0.9479, and 1.1430. The p-values of the following tests are reported: Portmanteau test (Port); Breusch/Pagan heteroskedasticity test (BP); Ramsey test (RESET); Jarque-Bera test on normality (JB). RMSE is the root mean squared error.

6. Conclusions

The global turmoil that began in 2020 has been inducing governments to allocate huge amounts of financial resources to capital purchases. Employing quarterly data over 1960–2022 for France and the US, we attempt to ascertain whether public investment impacts asymmetrically private capital formation in the short- and the long-run.

The main findings of the present study can be summarised in the following points. First, public investments in France have expansionary and persistent effects on private capital over a horizon of 10 years, likely due to demand effects and increases in the marginal productivity of private capital. Second, consistent with some previous related studies for the US, our estimates indicate that variations in the flow of public capital have neutral or negative effects on private investment, depending on whether the models are conditioned to additional observable economic and financial factors. Such a result suggests that the positive pressure on the demand-side of the economy and any increase in the marginal productivity of capital are likely offset by distortive mechanisms, such as higher input prices, higher tax rates or the expectation of future fiscal adjustments (see, for instance, Cochrane (2011)). However, such findings are compatible with expansionary effects on aggregate output since fiscal policy might generate heterogeneous effects across the economic sectors (Blanchard and Perotti, 2002). Third, the asymmetric effects of public investment in France are insignificant in the short-run but turn out to be significant and permanent after approximately ten quarters. Such asymmetry is due to a greater sensitivity of the private sector to positive changes in public investment, implying that the negative long-run effects of a one-unit reduction in the flow of public capital can be compensated by a rise in public investments ranging from 0.62 to 0.85 units. Fourth, no robust evidence of asymmetric effects emerges for the US in the short-run. Still, they become significant in the long-run – provided that the cointegrating vector includes the output per capita, the real long-term interest rate on government bonds and the international openness index. In such a case, the asymmetries are persistently negative after ten quarters following the fiscal impulse and the related ratio between positive and negative long-run effects ranges from 2.24 to 4.83. However, compared to France, the statistical significance of the asymmetries is lower.

The different findings for France and the US suggest that the macroeconomic effects of fiscal policy are characterized by high degrees of heterogeneity, consistent with recent evidence (e.g. Afonso and Sousa, 2012; Augustine and Rafi, 2023; Carvelli, 2023). Accordingly, policy recommendations should be parsimonious and consider the country-specific macroeconomic characteristics. Moreover, it appears worth suggesting that future research should accommodate any asymmetric effects in the public-private investments nexus, especially when modelling long-run equations.

Declaration of Competing Interest

None.

Data availability

Data will be made available on request.

Appendix

Fig. A1. Plots of CUSUM tests for parameter stability, baseline models. Cumulative sum of recursive residuals with 95% confidence bands around zero. Model A does not include additional controls. Model B includes the GDP growth rate in the vector of controls.

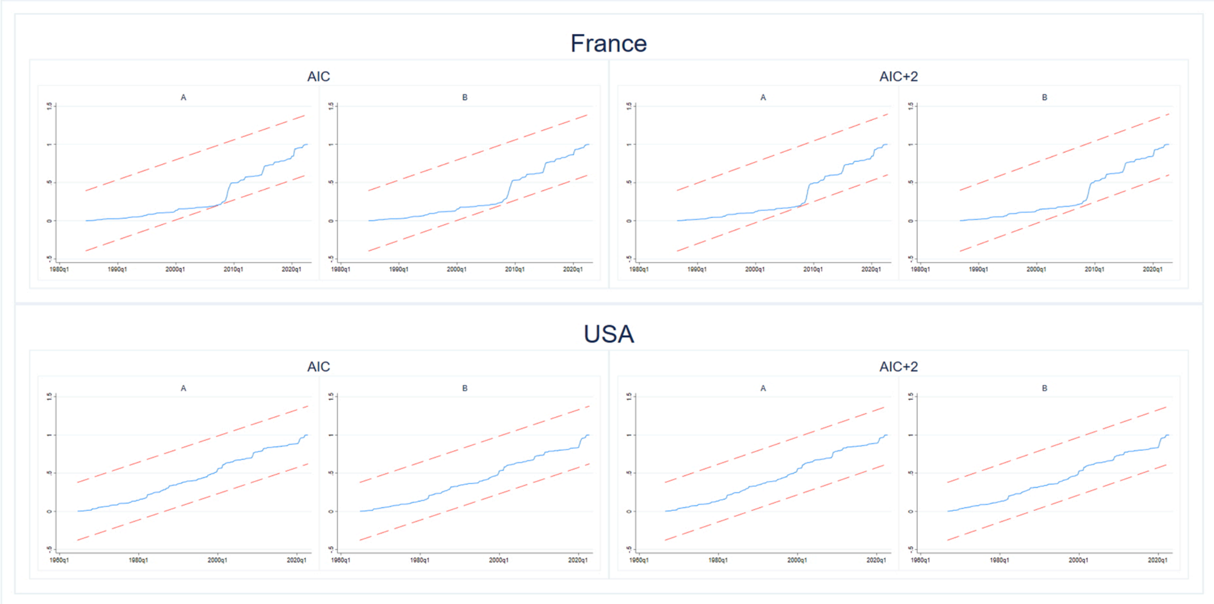

Fig. A2. Plots of CUSUM-Q tests for parameter stability, baseline models. Cumulative sum of recursive residual squares with 95% confidence bands around zero. Model A does not include additional controls. Model B includes the GDP growth rate in the vector of controls.

Fig. A3. Plots of CUSUM tests for parameter stability, models with additional controls. Cumulative sum of recursive residuals with 95% confidence bands around zero. Model A includes the output per capita in the vector of controls. Model B includes the output per capita, the real interest rate and the openness index in the vector of controls.

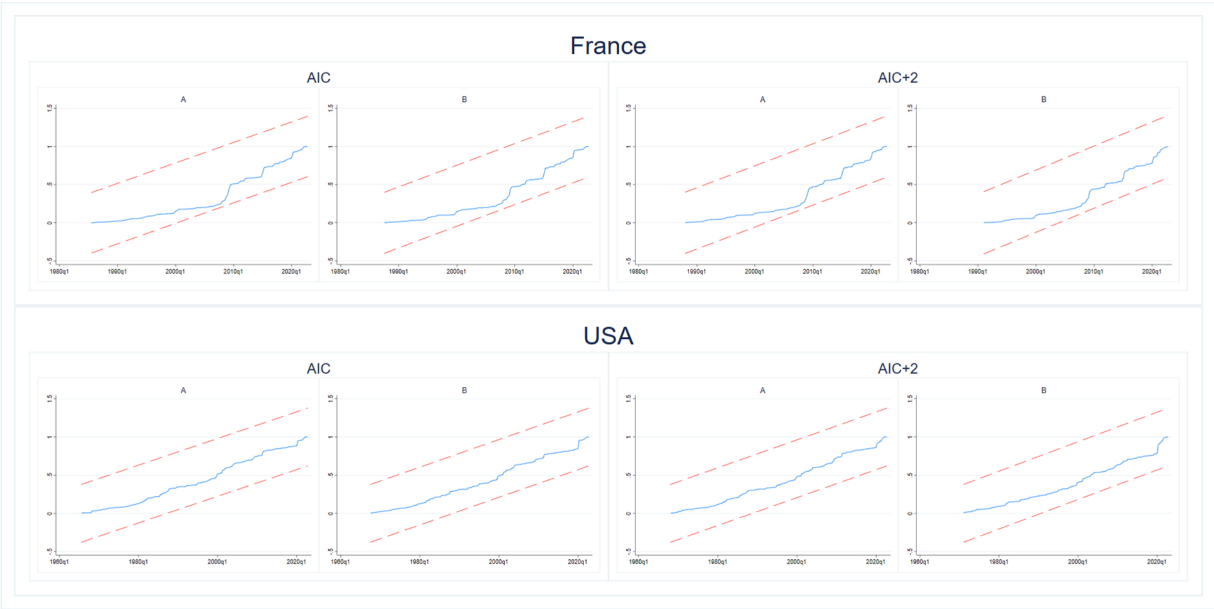

Fig. A4. Plots of CUSUM-Q tests for parameter stability, models with additional controls. Cumulative sum of recursive residual squares with 95% confidence bands around zero. Model A includes the output per capita in the vector of controls. Model B includes the output per capita, the real interest rate and the openness index in the vector of controls.

Table A1
Unit root tests, France.

	Levels				First differences			
	Without trend		With trend		Without trend		With trend	
	Z(t)	*P*-value	Z(t)	*P*-value	Z(t)	*P*-value	Z(t)	*P*-value
ADF								
Private invest., Lag(1)	−1.753	0.4040	−3.491	0.0403	−8.333	0.000	−8.354	0.000
Private invest., Lag(2)	−1.873	0.3450	−3.789	0.0171	−7.110	0.000	−7.137	0.0000
Private invest., Lag(3)	−1.832	0.3647	−3.867	0.0135	−5.622	0.0000	−5.549	0.0000
Private invest., Lag(4)	−2.059	0.2614	−3.201	0.0841	−5.340	0.0000	−5.285	0.0001
Public invest., Lag(1)	−1.360	0.6012	−2.267	0.4524	−9.804	0.000	−9.791	0.000
Public invest., Lag(2)	−1.502	0.5324	−2.454	0.3512	−8.193	0.000	−8.182	0.000
Public invest., Lag(3)	−1.546	0.5104	−2.485	0.3353	−6.519	0.0000	−6.508	0.0000
Public invest., Lag(4)	−1.775	0.3927	−2.764	0.2106	−6.219	0.0000	−6.207	0.0000
PP								
Private invest., Lag(1)	−1.673	0.4452	−3.230	0.0785	−12.115	0.000	−12.123	0.000
Private invest., Lag(2)	−1.730	0.4159	−3.279	0.0698	−12.135	0.000	−12.142	0.000
Private invest., Lag(3)	−1.758	0.4016	−3.300	0.0662	−12.140	0.0000	−12.146	0.0000
Private invest., Lag(4)	−1.828	0.3668	−3.365	0.0563	−12.193	0.0000	−12.198	0.0000
Public invest., Lag(1)	−1.291	0.6334	−2.141	0.5229	−14.346	0.000	−14.324	0.000
Public invest., Lag(2)	−1.360	0.6014	−2.230	0.4728	−14.382	0.0000	−14.361	0.0000
Public invest., Lag(3)	−1.404	0.5801	−2.290	0.4395	−14.407	0.0000	−14.385	0.0000
Public invest., Lag(4)	−1.462	0.5522	−2.365	0.3982	−14.467	0.000	−14.445	0.000

Notes. Augmented Dickey–Fuller (ADF) and Phillips–Perron (PP) tests. The *p*-values are approximated following MacKinnon (1996).

Table A2
Unit root tests, US.

	Levels				First differences			
	Without trend		With trend		Without trend		With trend	
	Z(t)	*P*-value	Z(t)	*P*-value	Z(t)	*P*-value	Z(t)	*P*-value
ADF								
Private invest., Lag(1)	−0.415	0.9075	−3.099	0.1066	−6.855	0.000	−6.843	0.000
Private invest., Lag(2)	−0.811	0.8160	−3.783	0.0174	−6.003	0.000	−5.995	0.0000
Private invest., Lag(3)	−0.825	0.8118	−4.076	0.0068	−6.111	0.0000	−6.097	0.0000
Private invest., Lag(4)	−0.822	0.8127	−3.813	0.0159	−6.472	0.0000	−6.464	0.0001
Public invest., Lag(1)	−2.026	0.2753	−1.631	0.7801	−11.307	0.000	−11.412	0.000
Public invest., Lag(2)	−2.372	0.1499	−1.907	0.6510	−8.685	0.000	−8.820	0.000
Public invest., Lag(3)	−2.546	0.1048	−2.087	0.5535	−7.415	0.0000	−7.673	0.0000

(*continued on next page*)

Table A2 (*continued*)

	Levels				First differences			
	Without trend		With trend		Without trend		With trend	
	Z(t)	*P*-value	Z(t)	*P*-value	Z(t)	*P*-value	Z(t)	*P*-value
Public invest., Lag(4)	−3.513	0.0076	−2.886	0.1671	−5.250	0.0000	−5.441	0.0000
PP								
Private invest., Lag(1)	−0.110	0.9484	−2.221	0.4779	−9.541	0.000	−9.536	0.000
Private invest., Lag(2)	−0.244	0.9330	−2.507	0.3244	−9.678	0.000	−9.674	0.000
Private invest., Lag(3)	−0.336	0.9202	−2.720	0.2281	−9.876	0.0000	−9.873	0.0000
Private invest., Lag(4)	−0.396	0.9108	−2.866	0.1738	−10.024	0.0000	−10.021	0.0000
Public invest., Lag(1)	−2.012	0.2813	−1.684	0.7577	−17.942	0.000	−18.001	0.000
Public invest., Lag(2)	−2.012	0.2813	−1.688	0.7562	−17.903	0.0000	−17.964	0.0000
Public invest., Lag(3)	−2.013	0.2810	−1.703	0.7494	−17.864	0.0000	−17.927	0.0000
Public invest., Lag(4)	−2.016	0.2798	−1.737	0.7342	−17.828	0.000	−17.886	0.000

Notes. Augmented Dickey–Fuller (ADF) and Phillips–Perron (PP) tests. The *p*-values are approximated following MacKinnon (1996).

Table A3
Zivot-Andrews unit root test.

	Levels			First differences		
	Intercept	Trend	Both	Intercept	Trend	Both
France						
Private invest.	−3.733	−3.564	−3.845	−5.968***	−5.595***	−5.932***
Public invest.	−3.437	−3.128	−3.438	−6.967***	−6.642***	−7.193***
USA						
Private invest.	−2.907	−3.755	−3.119	−7.032***	−6.839***	−7.041***
Public invest.	−2.452	−4.300*	−3.941	−8.680***	−8.124***	−8.920***

Notes. The test verifies the null hypothesis of unit root in the time series against the alternative hypothesis of stationarity.

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