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ORIGINAL ARTICLE

The productive effect of transport infrastructures: does road transport liberalization matter?

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Abstract In this paper we analyze the impact of motorway networks on production for a panel of twenty one manufacturing and service sectors of eleven EU countries observed over the period 1980–2003. In particular, we analyze if the degree of regulation of the road transport sector affects the link between productivity and motorway infrastructures. Overall results suggest that output elasticity with respect to motorway is found to be lower for countries characterised by a high degree of entry barriers in the road transport sector. This result is found to be more evident for industries which rely more heavily on transport services.

Keywords Transport · Deregulation · Public infrastructures · Productivity

JEL Classification L43 · L91 · L92 · H54

1 Introduction

In this work we analyze whether the impact of road transport infrastructures (motorway) on industry production is modified by the degree of regulation observed in the road transport sector: given that road transport services are important production inputs we argue that possible inefficiencies in the road transport sector due to the presence of regulatory barriers might weaken the productive impact of transport infrastructure investments. This issue is investigated by estimating a production function model on industry level data for a panel of eleven EU countries observed over the period 1980–2003.

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Public infrastructure in general, and in particular transport networks (such as roads, railways, airports, and waterways) have long been considered important inputs to economic and productivity growth. Economic theory has made it clear that the services provided by transport infrastructures may positively impact industry productivity in different ways,[1] the main channel being a reduction of time and transport costs, which in turn can have different possible implications such as higher productivity of other inputs, lower production costs, greater specialization, growth of trade, more intensive competition, enlargement of relevant markets, changes in location decisions that allow the development of spatial clusters, improvement in the division of labour, better access to foreign intermediates, exploitation of scale economies, more efficient management of inventories, better coordination of decisions among suppliers and consumers, more efficient allocation of resources between firms and sectors.

The importance of the transport system as a possible factor enhancing economic growth has also been widely recognized by policy makers as witnessed by the publication in 2001 of the EU White Paper "European transport policy for 2010: time to decide", which suggests that an efficient transport system is essential for keeping Europe's economy competitive and its Single Market running smoothly. The same White Paper highlights the importance of road infrastructure within the transport system, by showing that road makes up 44% of the goods transport market compared with 41% for short sea shipping, 8% for rail and 4% for inland waterways.[2]

The Aschauer (1989) seminal paper -which ascribed most of the US post oil shock productivity slowdown to a lack of investments in public infrastructures- spurred a wide empirical literature on the effects of transport and, more generally, public infrastructures on production. A necessarily brief survey would not do justice to the variety of theoretical approaches and empirical frameworks that have been adopted. Afraz et al. (2006) and Romp and de Haan (2007) provide excellent surveys of this literature, which generally suggests the existence of small but not-negligible effects of public infrastructures expenditure on production.[3] Limiting our discussion to some of the most recent works which are more closely related to the relationship between production and the road network, we mention Cohen and Morrison (2004) who estimated a variable cost function for the manufacturing sector using US states panel data, and found that an increase of the motorway capital stock of 10% tends to be associated with a reduction of variable costs of about 1.5%. In turn, Bronzini and Piselli (2009) estimated a production function for a sample of Italian regions and found an elasticity of output with respect to the road stock of about 0.15, which is a very similar value to that reported in Boarnet (1998) for a sample of Californian counties. Finally, Fernald

[1] An overview of the theoretical literature on the links between economic growth and public infrastructure may be found in Afraz et al. (2006).

[2] The EU road transport sector has experienced a sustained growth during the last decades for different reasons, such as the abolition of physical frontiers, the implementation of the Single Market Programme, the introduction of measures which encourage cross-border cooperation between firms and businesses, the reduction in heavy bulk transport and the increased importance of door-to-door and just-in-time services.

[3] Most of the recent studies that use state-of-the-art econometric approaches tend to identify elasticities of output with respect to public infrastructure capital stock ranging between 0.10 and 0.20.

(1999), using a panel of US industries, found a positive impact of the motorway stock on TFP growth.

The empirical literature that we have just briefly considered has however mainly focused on the role played by the extent of the network, with only very few authors that have empirically controlled for issues related to congestion (e.g. Fernald 1999) and network quality (Hulten 1996; Calderon and Serven 2005), as far as road infrastructure is concerned.

In this work we add to this literature by arguing that the productive effect of the road transport infrastructure does not rely only on infrastructures' features like network length, roads quality and congestion, but it may also depend on the degree of product market competition in the road freight sector which is the main industry that allows the infrastructure to fully exert its productive potential.[4]

In particular, we suggest that a heavy regulated road transport sector, characterized by high barriers to entry, will be a sector insulated from healthy competition and thereby characterized by low innovation, by the survival of many small and inefficient operators and by lower quality and higher prices.[5]

As road transport services are important inputs in the production process of many industries, higher road transport costs and lower quality of the service may translate into significantly higher production costs so that the expected reduction in transport costs stemming from road infrastructure investments might be partly crowd-out. Moreover, possible higher transport costs associated with high regulation in the road freight sector might translate into a lower utilization of the road network: as a result, some of the channels cited above through which motorway infrastructures could enhance industry productivity might be further weakened.

In the following section we discuss the empirical model and illustrate our empirical strategy. Section 3 presents the data and is followed by the discussion of the empirical results. Section 5 concludes.

2 Model specification

In order to analyze if a high level of regulation in the road transport sector weakens the productive effects of road infrastructures, we assume that firms produce gross output according to the following Cobb-Douglas technology:

$$Y_{ijt} = TFP_{ijt} K_{ijt}^{\alpha} L_{ijt}^{\beta} M_{ijt}^{\gamma} \tag{1}$$

4 Economic theory suggests that anti-competitive regulation in a particular sector may have a direct impact on market conditions in the same sector by reducing allocative efficiency, stifling innovation and reducing investment rates. Schiantarelli (2008) presents a critical overview of the theoretical and empirical contributions that use cross-country data to provide insights on the direct economic effect of product market regulations/reforms.

5 A number of studies have shown, for the US case, that deregulation in the road transport sector led to lower transport prices and to changes in the concentration and organization of the sector (Boyer 1993; Belman and White 2005). In the case of the EU, the empirical evidence is very limited: a recent work by Lafontaine and Malaguzzi Valeri (2008) shows that deregulation in the EU had a positive effect on the rate of growth of the international trucking sector.

where Y_{ijt} is the gross output in sector *i* of country *j* at time *t*, and K, L and M are the associated capital stock, index of labour services and intermediate inputs used in the production process and, finally, TFP_{ijt} represents total factor productivity in sector *i* of country *j* at time *t* and α, β and γ represent the output elasticity of capital, labour services and intermediate inputs, respectively, whose sum is not constrained to equal one. Total factor productivity, in turn, can be represented as in the following equation:

$$TFP_{ijt} = A_{ijt} H_{jt}^{\upsilon} \quad (2)$$

where H_{jt} represents the network of motorway in country j at time t, A_{ijt} represents the other neutral total factor productivity determinants not already accounted for in the model and υ is the elasticity of gross output with respect to the stock of motorway. Substituting Eq. 2 into 1, we get:

$$Y_{ijt} = A_{ijt} H_{jt}^{\upsilon} K_{ijt}^{\alpha} L_{ijt}^{\beta} M_{ijt}^{\gamma} \quad (3)$$

Equation 3 is the standard equation estimated in most studies that seek to evaluate the impact of transport infrastructure capital on productivity levels. In this work we assume that the elasticity of gross output with respect to the motorway infrastructure stock depends linearly on the degree of regulation in the road transport sector:

$$\upsilon = \eta_0 + \eta_1 R_{jt} \quad (4)$$

where R_{jt} represents the degree of regulation in the road transport sector of country j at time t as proxied by the entry barrier index described in Sect. 3.

In turn, the (unobservable) neutral total factor productivity term A_{ijt} can be decomposed as follows:

$$A_{ijt} = \exp\left(t_t + \varepsilon_j + u_{ij} + v_{ijt}\right) \quad (5)$$

where v_{ijt} is the idiosyncratic error term; t_t is a time period effect, common to all cross sections in our sample, which accounts for common macroeconomic shocks or technological improvements and is represented by a set of time dummies; u_{ij} represents country-sector time invariant effects which capture unobservable country and industry specific characteristics, which may be correlated with regressors, such as permanent differences in industrial specialization across countries or differences in technologies or regulatory environments across industry-country combinations; ε_j represents a set of country specific dummy variables which capture unobservable time invariant effects, common to every sector in each country, that might drive productivity, such as the degree of urbanization, or country-specific institutional frameworks related to the motorway sector[6] among other things. Moreover, these country-specific dummy variables might mitigate estimation problems arising from possible residuals correlation

6 At the beginning of this decade, on average about 38% of the European motorway network was under concession and 53% of the concessions were managed by the public sector, although these values show high variability across countries. Moreover, the structure of concession agreements differs across countries, for example in terms of length, risk sharing, remuneration and so on. Since detailed information on these

across sectors belonging to the same country.[7] Furthermore, in the robustness section we will estimate different specification of the baseline models which include country or industry specific time trends. In particular, country trends allow us to control for the effects of country specific policy changes (e.g. deregulation in sectors others than transport) or the existence of different timings in the transmission of macroeconomic shocks. Industry specific trends may instead capture industry specific patterns of growth stemming, for example, from differences in skill biased technical change.

After substituting Eqs. 4 and 5 into Eq. 3 and taking logs, we get our estimated equation (where lower case variables denote natural logs):

$$y_{ijt} = \alpha k_{ijt} + \beta l_{ijt} + \gamma m_{ijt} + \eta_0 h_{jt} + \eta_1 \left(h_{jt} R_{jt}\right) + t_t + \varepsilon_j + u_{ij} + v_{ijt} \quad (6)$$

In order to better investigate whether the productive effect of motorway infrastructure is reduced in countries with a higher level of regulation in the road transport sector, we analyze if this possible effect is more evident for those sectors which rely more intensively on transport services, the intuition being that those sectors that rely more intensively on transport services should also be those where the impact of regulation is higher: if a sector does not use transport services at all, we might then expect that regulation in the road transport sector matters less. This latter approach is related to the seminal paper by Rajan and Zingales (1998) who argued that the impact of cheaper external finance on productivity growth should be stronger in those sectors that rely more extensively on external finance. A similar approach is adopted by Barone and Cingano (2008) who examine whether OECD countries with less anti-competitive regulation in service sectors show a better economic performance of manufacturing industries that use less regulated services more intensively.

Using a methodology described in Appendix 1 we have therefore derived, for each sector, a (country-invariant) transport intensity variable, w_i. Following Rajan and Zingales (1998), we have multiplied the interaction variable $h_{jt} R_{jt}$ with w_i: a negative coefficient for the variable $h_{jt} R_{jt} w_i$ would indicate that a higher degree of regulation in the road transport sector tends to decrease the output elasticity with respect to motorway infrastructure particularly in sectors that rely more intensively on transport services. As an alternative approach, we split the sample in low and high transport intensity industries depending on whether w_i was below or above the sample median and we then allow η_1 to be different for the two groups.

Turning to the estimation strategy, Eq. 6 is estimated using the GMM-SYS approach proposed by Blundell and Bond (1998) and by Arellano and Bover (1995), which appears particularly suitable when estimating production functions with persistent

Footnote 6 Continued
institutional features are not available for all countries over the whole sample period we cannot include them in the model.

[7] The main variables of interest in this paper, the motorway network H and the degree of liberalization R in the road transport sector, are defined at country, rather than at the sector-country level: this might lead to underestimate standard errors. The use of clustered standard errors might correct for possible residuals correlation across sectors belonging to the same country (Moulton 1990). Nevertheless, when the number of groups is small and even lower than the number of industries within each group, as it is the case in our sample, clustered robust standard errors are unlikely to perform reasonably well (Wooldridge 2003).

data and simultaneity issues arising from possible correlation between explanatory variables and disturbances.[8] In particular, both the motorway network and its interaction with the road regulation variable may be considered endogenous. We address this issue by using appropriate instruments for both variables and by comparing industries with different degrees of transport intensity: if road transport regulation is found to reduce motorway elasticity disproportionaly in transport intensive industries, we argue that this is less likely to be generated by a spurious relationship.

Given the length of the time span covered by our sample we decided to investigate the time series properties of our variables. Indeed OLS and fixed effects estimates of autoregressive models displayed autoregressive parameters ranging between 0.95 and 0.99, thus showing the presence of quite large persistence in our series. We therefore decided to undertake a fully-fledged empirical investigation of the time series properties of the series. We run a battery of panel unit root tests, namely the Levin-Lin-Chu, Breitung, Maddala-Wo, Im-Pesaran and Shin as well as Pesaran CADF tests, five of the most popular tests in the panel unit root literature.

For all variables but the motorway network length, in four out of five tests we had to reject the null hypothesis of the existence of a unit root in the series and these results were broadly confirmed when we conducted the panel unit rot tests separately for each country.[9,10] Summarizing, there is not a clear cut evidence on the time series properties of the series, although the majority of tests would suggest that most variables are I(0).

Nevertheless, since stationarity tests may perform weakly when the variables are characterized by near unit root processes, we also performed some robustness analysis in order to account for problems stemming from the possibility of the existence of spurious regressions. In particular, we estimated our preferred model by applying instrumental variable techniques after taking first differences and with the GMM-DIFF approach.

3 Data

The dataset we employ in this paper is made up of industry level data for a sample of eleven EU countries, namely Austria, Belgium, Denmark, Spain, France, Italy, UK,

[8] The GMM difference approach, which removes time invariant heterogeneity (the country-sector effects u_{ij} in Eq. 6) by first differencing, and uses lagged (level) instruments, have usually proven to provide too small estimates of the capital stock coefficient, especially when inputs and output are characterized by strong persistency or by an almost random walk behaviour (as it is the case in our sample). As shown by Blundell and Bond (1998), this is essentially a weak instrument problem: when variables are very persistent, lagged levels are often a poor proxy of the current change in the endogenous variable. Blundell and Bond (1998) therefore suggested an alternative estimator, the GMM-SYS, that exploits more informative moment conditions and provides strong asymptotic and finite sample efficiency gains as well as reductions in the small sample biases which plague the GMM-difference estimator.

[9] For the motorway network length, three tests out of five lead us to reject the null hypothesis of the existence of a unit root. Results are available from the authors upon request.

[10] All panel unit root tests were conducted assuming the existence of individual fixed effects as well as time trends. The lag structure was selected using the Akaike Information Criterion. See Breitung and Pesaran (2005) for an excellent discussion of the relative merits of the different panel unit root tests available in the literature.

the Netherlands, Sweden, Finland and Germany, observed over the period 1980–2003. For each country we consider twenty two sectors (eleven manufacturing, nine service plus agriculture and mining) as shown in Table A1.

Our sample is made up of 217 cross sections observed for 24 years: given the unbalancedness of our dataset (mainly because of Germany and Sweden, for which we have data only after 1991 and 1993, respectively) we end up with about 4620 observations. Different sources have been used to build the dataset. The major one is the recent KLEMS dataset of the University of Groningen. From the KLEMS database we have taken data on gross output (Y), intermediates (M) and hours of work (L), while the capital stock figures are taken by the OECD STAN dataset or derived applying the perpetual inventory method (see Appendix 1 for a detailed description). To take into account capacity utilization issues we have followed Griffith et al. (2004) and employed a capacity utilization-adjusted capital stock series; similarly, following Harrigan (1999), we have corrected row hours worked by the workers skill levels. All monetary figures have been expressed in constant 1995 prices and converted to a common currency using appropriate PPP indices.

As far as the infrastructure variables is concerned, the motorway stock was proxied by the Km of motorway network taken from EUROSTAT, while the public capital stock series (which we have used as a robustness check) was taken from the University of Kiel website and it is described in Kamps (2004).[11,12]

The first two columns in Table A2 in Appendix 1 report the length of the motorway network as of 1980 and 2003. As we can see, the motorway networks increased substantially in most countries, Spain and Finland being the countries that increased it the most and Italy and The Netherlands those that increased it the less (either because most of the motorway network had already been built by 1980, or maybe because they simply failed to extend it in the later part of our sample). In turn, the last two columns in Table A2 report the same information but after normalizing the motorway network by the respective country's area size (measured in km^2): as we can see, there are large differences among countries, reflecting, among the other things, different degrees of urbanization rates and differences in the pattern of population distribution. For instance, the two Scandinavian countries in our sample display very low ratios of motorway network km per km^2, reflecting the fact that most of the inhabitants are concentrated in a few areas. At the other extreme there is France, that by far displays the largest "density" of motorway in the sample.

Turning to the liberalization variables, the main source of data was the OECD regulatory database, which contains liberalization indices for a set of utility sectors, namely air transport, road transport, railways, telecom, gas, electricity and post.[13] The degree of liberalization in the road sector of each country is captured from two sub-indices. The first index, EB, is a variable ranging from 0 (fully liberalized) to 6 (fully

[11] In terms of km^2, the motorway length and the total public capital stock display a correlation coefficient of about 0.95.

[12] Although it would have been interesting to compare the results with a monetary indicator of the stock of motorway, comparable figures across countries do not exist.

[13] For a detailed description of the procedures used to build the indexes on regulation in non manufacturing sectors of OECD countries, see Conway and Nicoletti (2006), pg 31.

regulated) that proxies the extent of entry barriers in the road transport sector and it is built by jointly considering five different issues: (a) existence of a licence or permit to establish a national road transport service; (b) existence of criteria other than safety requirements, technical and financial fitness considered in decisions on entry of new operators; (c) ability of the regulator to limit capacity; (d) existence of professional bodies involved in specifying and enforcing entry regulations; (e) existence of professional bodies involved in specifying or enforcing pricing guidelines or regulations. The index is calculated using a bottom-up approach, in which the regulatory data on the five issues listed above are quantified using an appropriate scoring algorithm (see Conway and Nicoletti 2006) and then aggregated into a summary indicator.

The second one, *PR*, is constructed similarly to the EB index and is again a variable ranging from 0 (full liberalization) to 6 (very high regulation), proxying for the importance of price controls in the road transport sector. It was built by the OECD by considering whether in the road sector of a particular country: (a) the government regulates in some way retail prices of road transport services and, (b) the government provides pricing guidelines to road transport companies. This index has been used as an alternative regulation index as a robustness check.

Table A3 in Appendix 1 shows the values of *EB* and *PR* in the eleven countries in our sample observed in some selected years from 1980 to 2003. The entry barrier index shows a high variability across countries, with a clear tendency to fall over time from values as high as 6 in most countries, although the timing of liberalization has not been uniform. Italy stands up as the country with the highest barriers to entry in 2003, with a value of 5.01 compared to an average (excluding Italy) of 2.29.

Price regulation in the road transport sectors was relaxed in virtually all countries over the sample period, the only exception being, on one side, Italy, which still had a value of 6 for *PR* in 2003 and the UK, which had already abolished any form of price control as early as 1980.

For a description of the other variables that have been used in this study as additional instruments or control variables we refer to Appendix 1.

4 Empirical results

We estimated our baseline specification as in Eq. 6 with the GMM-SYS[14] approach and we report estimates results in Table A4. All estimates include country-sector fixed effects, time effects as well as a full set of country dummies. Given simultaneity issues affecting production functions estimates, private inputs have been instrumented with their own appropriate lags. The same approach has been adopted for the motorway network variable, as there might be a reverse causality problem: since transport infrastructure investment might depend on the level of output, a productivity shock might be associated with a variation of the motorway network, thereby causing biased estimates of the elasticity of output with respect to the motorway network.[15] In particular,

[14] Standard errors are two-step robust and include the Windmeijer (2005) correction.

[15] Cohen and Morrison (2004) argue that the reverse causality issue might be less important in studies which employ sectorial level data.

h_{jt} has been instrumented with h_{jt-2}, which is a valid instrument provided that h_{js} is uncorrelated with u_{ijt}, with $s < t$: in other words, our identification assumption is that there is no correlation between current network extensions and future productivity shocks; on the other hand, our estimates are fully robust to possible correlation between past and present shocks to productivity and current decisions to extend the network. A similar reasoning has been applied for the entry barrier index which might be endogenous, as a productivity shock might be correlated with contemporaneous change in the regulatory environment: again, we have instrumented R_{jt} with R_{jt-2}.[16] For all estimated models, both the Hansen and Sargan tests confirm the validity of the instruments set employed and estimates do not exhibit problems of serial correlation (as shown by the Arellano-Bond test).

Estimates reported in Column 1 are based on a simple model which includes the motorway network together with the standard production function inputs. Private input elasticities show reasonable values (suggesting weak decreasing returns to scale); the elasticity of output with respect to the motorway stock is positive, statistically significant, and with a magnitude of about 0.14. Cohen and Morrison (2004) report an elasticity of manufacturing costs with respect to the motorway stock of about 0.15 for the US case which, although not directly comparable to a production function elasticity, is remarkably similar to ours; furthermore, most studies which focus on the impact of public capital on productivity found output elasticities approximately ranging between 0.10 and 0.20.[17] We also obtained very similar results by including a relative measure of the motorway infrastructure computed as the ratio of network kms and country's area size or as the ratio of network kms and population.

If we allow for country-specific motorway elasticities we note that there are some differences across countries, with the Scandinavian countries displaying the largest elasticities and Italy, Spain and France the smallest ones.[18] In part these differentials across countries might stem from differences in the motorway network extension, but other effects (like regulation in the road transport sector) might play an equally important role. In particular Italy displays the lowest elasticity (0.08) as opposed to Finland which has an elasticity of 0.18: this finding might be related to the different degrees of road transport regulation which characterize those two countries (see Table A3).

In Model 2 we checked whether the impact of the motorway network on output is nonlinear: it might be possible that once the main network has been laid out, further extensions might prove to be less productive. Parameter estimates show that this is indeed the case as the square term is negative and statistically different from zero so that motorway elasticity declines as the motorway network increases.[19]

[16] The regulation variable and the motorway network have been alternatively instrumented with external instruments such as the Freser index of trade barriers, a dummy representing the head of government's political orientation and an Herfindhal index of government fragmentation: main results are confirmed.

[17] If, instead of the motorway network, we include the total public capital stock we obtain an elasticity of about 0.17 (with a p value of 0.09), which is quite in line with previous literature.

[18] Average motorway elasticity is about 0.12. We do not report results for reason of space.

[19] The elasticity ranges from 0.12 at the 25th percentile, to 0.10 at the median and to 0.06 at the 75th percentile, all significantly different from zero with the exception of the value computed for the 75th percentile.

In order to check if the elasticity of output with respect to motorway network is affected by the degree of liberalization in the road transport sector we estimated Model 3 where we include an interaction variable between the motorway network and the entry barriers index. Estimates show that an higher level of regulation in the road transport sector reduces the impact of the motorway infrastructure on productivity[20] and this result holds when we use the price regulation index as an alternative to the entry barriers index. This finding is consistent with the hypothesis that high regulation in the road freight sector may induce inefficiencies and high transport prices, which in turn leads to an increase in transport costs for those sectors which use transport services as intermediate inputs; as a result, the productive effect of the motorway network, which operates by lowering trade and transport costs, may be weakened. Furthermore, it may also be the case that higher transport costs induce a lower utilization of the road network and, as a result, some of the channels through which motorway infrastructures could enhance industry productivity might be further weakened: for example, there might be a narrowing of relevant markets, worse access to foreign intermediates, lower scope for exploitation of scale economies and a less efficient allocation of resources between firms and sectors.

Further empirical evidence is obtained by estimating Model 4 where we allow the interaction variable to vary according to the different degrees of sectorial dependence on land transport.[21] In principle, we expect that industries which rely more heavily on transport services should be more affected by the intensity of regulation in that sector: if this is the case we can reasonably argue that motorway elasticity differentials might be attributed to the different degrees of liberalization in the road transport sector.

Punctual estimates of Model 4 suggest that regulation in the road transport sector affects negatively motorway elasticity for those sectors that rely more intensively on transport services. In particular, motorway elasticity is about 0.10 for countries with a very high degree of entry barriers while it is about 0.13 for countries with the lowest levels of regulation.

An alternative approach to discriminate sectors on the basis of their transport intensity has been adopted in Model 5 where we weight the interaction variable between the motorway stock and the entry barrier index with weights that proxy for the degree of dependence of each sector on transport services (see Appendix 1). Estimates show that the coefficient of the new interaction variable is negative and statistically significant. Punctual estimates suggest that for sectors with the highest transport intensity the motorway elasticity falls from about 0.15 for countries with the lowest levels of regulation to about 0.13 for countries with the highest levels of road transport sector regulation. This effect is however considerably reduced for countries with a medium-low level of transport services dependency.[22]

We think that contrasting industries with different transport intensities allows us to reduce biases stemming from omitted variables and endogeneity issues which would not allow us to identify a causal effect from road regulation to output; since we expect

[20] This result is confirmed when we include the regulation term lagged one period.

[21] See the Model Section.

[22] Even if the aim of this paper is to better understand the link between motorway investment on productivity, our empirical model also allows us to evaluate the impact of deregulation in the road freight sector

that industries which rely more heavily on transport services would be those which are more affected by the intensity of regulation in that sector, we argue that estimated motorway elasticity differentials might be linked to different degrees of liberalization in the road transport sector.

4.1 Robustness analysis

We now turn to discuss some robustness analysis conducted mainly on Model 5, which we consider as our preferred specification.

As a first robustness check, we have estimated Model 5 after dropping each country, one at a time: estimates suggest that main results are not driven by the inclusion of a specific country and the same conclusion can be drown by estimates obtained after dropping each sector. In both cases the coefficient of $h_{jt} R_{jt} w_i$remained negative, statistically significant and with a magnitude broadly comparable to that reported in Table A4.

All estimated models that include regulation in the road transport sector assume that the latter influences production mainly by altering the productive effect of motorway infrastructure investments. However, we believe it may be important to check whether our result is confirmed when we control for a possible direct effect of road transport regulation on downstream industry output.[23] We have therefore included in Model 5 the entry barrier index as an additional regressor. Estimates shown in column 1 of Table A5 suggest that both the entry barrier index and its interaction with motorway infrastructure are negatively signed although individually noisily estimated. An F test on the joint significance of both variables leads us to reject the null hypothesis that they are jointly equal to zero (with a p value of 0.03) thus reflecting the possible existence of multicollinearity problems which do not allow us to obtain correct inference on the significance of individual parameters. Nevertheless, the magnitude of the interaction term is broadly consistent with estimates of Model 5, thus suggesting that even after controlling for the possible direct effect of road transport regulation on production we still find that the elasticity of output with respect to the motorway stock is higher in countries with a more liberalized road transport sector.

Another possible issue which may be raised is that the entry barrier variable might pick up, among other things, the effects of the introduction of deregulatory reforms in other sectors of the economy. In order to capture these possible effects we have augmented Model 5 with three variables proxing for the degree of liberalization in the

Footnote 22 continued
on industry productivity. When we compute the marginal effect of deregulation on industry production we find that, for low transport intensity sectors (25th percentile), a unit decrease in the entry barriers index is associated with a one for all 0.6% increase in gross output; in turn, for high transport intensity sectors (75th percentile), the effect on production is about 1.9%.

23 Barone and Cingano (2008) and Arnold et al. (2008a,b)) found a positive impact of deregulation in service sectors (transports, electricity and telecommunication) on the productivity of downstream industries which use the liberalized services more intensively as an intermediate input.

labour (lab_{jt}) and credit ($credit_{jt}$) markets as well as with a variable proxing for the degree of entry barriers in the utility sectors other than road transport ($Eothers_{jt}$): estimates shown in column 2 of Table A5 confirm our main results. Moreover, given that the time span covered by our data includes the period of implementation of the European Single Market Program, it might be useful to control whether our results are affected by omitted variables related to the abolition of tariffs and non-tariffs barriers which have contributed to create a more liberalized economic environment. In order to tackle this issue we estimated a model where the coefficient of the interaction variable ($h_{jt}R_{jt}w_i$) has been allowed to vary before and after 1992 and we could not identify significant differences between coefficients estimated over the two sub-periods.[24]

In column 3 of Table A5 we report estimates of our baseline model augmented with a full set of country trends which account, among other things, for country specific policy changes (e.g. changes in motorway charges policies), different implementation timings of deregulatory reforms, different productivity patterns at country level, specific trends in network quality features, such as the amount of motorway with three or more lanes or the importance of maintenance expenditures. Overall results confirm our main findings.[25]

The productive impact of the motorway stock might also be affected by the degree of congestion which characterizes the network, an issue that has not been sufficiently investigated in the relevant literature (Fernald 1999). In order to take into account possible effects of road congestion we have modified Eq. 4 by letting the motorway elasticity to depend not only on regulation but also on the degree of congestion. As a result, the empirical Model 5 is augmented with an additional regressor represented by the interaction between the motorway network and congestion ($h_{jt}cong_{jt}$), the latter measured as the ratio of national vehicle-kms to the total road-kms in each country.[26] Estimates reported in the fourth column of Table A5 confirm our main results and suggest that congestion tends to significantly reduce motorway elasticity; in particular, motorway elasticity for a network characterized by high levels of congestion (75th percentile) is found to be about two percentage point lower with respect to a less congested network (25th percentile).

This finding suggests that congestion may impose significant negative externalities on production that should be internalized in order to improve economic efficiency of the motorway system. As congestion might be a more serious problem in some parts of the motorway network (e.g. around big cities) or during some periods (e.g. peak hours, weekends, etc.), a structure of tolls that varies according to the degree of

[24] Results are available upon request.

[25] When we add to the model a full set of industry specific trends the coefficient of motorway stock elasticity and its interaction with road regulation are equal to 0.07 and −0.08, with p values of 1 and 10%, respectively. However, the labour and capital coefficients turn out to be poorly estimated.

[26] Vehicle km is the unit of measurement representing the movement of a vehicle over one km. Data are taken from EUROSTAT. Missing values have been linearly extrapolated and interpolated.

congestion might be an appropriate tool to address this externality issue, as required by the Directive 2006/38/EC (the so-called Eurovignette Directive).[27]

Another issue which deserves some discussion is related to possible spillovers effects arising from transport infrastructure stocks in neighbouring countries. A few empirical contributions suggest (Boarnet 1998 and Cohen and Morrison 2004) that when analyzing the effects of transport infrastructure on production, neglecting possible spillovers effects might result in biased estimate of infrastructures elasticity. In order to take into account this issue we have extended Model 5 by adding a variable which represents a weighted measure of the stock of motorway infrastructure in other countries (G_{jt}).[28] Estimates shown in column 5 of Table A5 suggest that positive spillovers might indeed exist: when the stock of motorway in the other countries rises by 10%, production rises on average by 1.7% (with a p value of 0.09), which is a quite large effect, but not uncommon in the literature (see Bronzini and Piselli 2009). The possible existence of non-negligible spillover effects suggests that transport infrastructures may be underprovided because individual member states do not internalize the full benefits arising from public infrastructure investments: this in turn lends empirical support to the active role played by EU institutions in financing major European transport infrastructure projects such as the so called Trans-European Networks.

As a further robustness check we applied an instrumental variable method after first differencing the model in order to eliminate country-sectors fixed effects (IV-FD)[29] and the GMM difference (GMM-DIFF) method of Arellano and Bond (1991). The last two columns of Table A5 report estimates results which confirm our previous finding: the elasticity of output with respect to the motorway stock is positive and is lower in countries with a higher level of entry barriers in the road transport sector.[30] We note that, by estimating our preferred model in first differences, we partly address issues deriving from the possible existence of unit roots in our series as the first difference specification minimizes concerns of spurious regression.

A final discussion is related to the possible issue of dynamic misspecification of our model. As a matter of fact, all estimated models are based on the assumption that production instantaneously adjusts, although there might be reasons to expect that the adjustment process could take some time to complete. In order to take into account this possibility, we have estimated an autoregressive distributed lag model (ARDL(1,1)) which can be considered an explicit dynamic approximation to an adjustment process:

[27] Unfortunately the Directive has established that the rules will apply to motorway under concession (which in some countries such as Italy and France account for most of the motorway network) only when the concession contract will be renewed, which in some cases may take many years.

[28] The construction of the other countries' motorway stock (G_{it}) is described in the Appendix.

[29] In practice, we have used the Anderson-Hsiao estimator. Standard errors are robust to heteroskedasticity and to arbitrary correlation within groups. Estimates are performed with the command IVREG2 for Stata 10 by Baum et al. (2007) which allows us to perform the Kleibergen Paap test which suggests that our equation is not underidentified.

[30] We have also estimated Model 5 with GMM-DIFF after including either a full set of industry trends or industry-years fixed effects. We do not report parameter estimates for reasons of space but main results do not change.

$$y_{ijt} = \delta y_{ijt-1} + \alpha_0 k_{ijt} + \alpha_1 k_{ijt-1} + \beta_0 l_{ijt} + \beta_1 l_{ijt-1} + \gamma_0 m_{ijt} + \gamma_1 m_{ijt-1} + \eta_0 h_{jt} + \eta_1 h_{jt-1} + \eta_2 \left(h_{jt} R_{jt}\right) + \eta_3 \left(h_{jt-1} R_{jt-1}\right) + t_t + \varepsilon_j + u_{ij} + v_{ijt} \quad (7)$$

In particular, the inclusion of the lagged level of output also allows for additional persistence in the adjustment process related to lagged responses of production to exogenous shocks due to the existence of adjustment costs that firms have to bear in order to reach the new equilibrium path. Moreover, the dynamic specification in Eq. 7 allows for the possibility that both the introduction of deregulatory reforms and an extension of the motorway network may take time to display their effects.

In Table A6 we report estimates of the long run production function elasticities computed from the parameter estimates of the ARDL(1,1) specification[31] which confirm previous results obtained with the static production function model.[32] In particular, the long run elasticity of motorway is about 0.12, while the interaction term suggests that such elasticity is found to be lower for countries characterized by heavier regulated road transport sectors: reassuringly, estimation results stemming from the dynamic specification are consistent with those obtained from the static model.

5 Conclusion

In this paper we add to the literature on the role of transport infrastructure on productivity growth as we analyze the impact of motorway networks on industry production on a panel of eleven European countries observed over the period 1980–2003. As suggested by Hulten (1996), we argue that "how well you use the infrastructure is much more important than how much you have of it" and we believe that road transport sector liberalization is an important factor which might affect the impact of motorway networks on industry production. In fact, a road transport sector characterized by a high level of regulation might be associated with higher inefficiency and higher prices: since transport services are important intermediate inputs in several production processes, regulatory barriers might weaken those channels through which motorway infrastructures could positively affect industries production.

The liberalization of the road transport sector is part of a larger program of regulatory reforms which have been introduced by most EU countries during the last two decades and whose effects on productivity have not been investigated extensively. To the best of our knowledge, this is the first study which jointly analyzes the effect of road transport infrastructure and of road transport sector liberalization on industry production.

Our estimates of a production function show that the average elasticity of output with respect to the motorway stock is about 0.14 and that it slightly declines as the motorway network increases. Overall results suggest that improvements in transport

[31] The ARDL(1,1) model has been estimated with the GMM-SYS method. Instruments used for the difference equation are l, m, k, h and y, all dated T-3; and Δl, Δm, Δk, Δh and Δy, dated T-2 for the level equation. Additional instruments are a full set of country and year dummies plus Por and HHI.

[32] The long run elasticity of private capital has been calculated as $(\alpha_0 + \alpha_1)/(1 - \delta)$ and similarly for the other variables.

infrastructure (as proxied by motorway network) seems to rise the productivity of private inputs by reducing the costs of production via a reduction in transport costs; this in turn might expand the relevant product markets thereby encouraging competition, stimulating specialization and exploitation of economies of scale. However the positive effect of transport infrastructure investments on output seems to be depressed by the lack of a liberalized road transport sector.

The elasticity of output with respect to motorway is found to be lower for countries characterized by a high degree of entry barriers in the road transport sector. This result is reinforced by estimates obtained after discriminating industries on the basis of their different degree of dependency on transport services; for those industries which rely more heavily on transport services, the link between production and transport infrastructures is more affected by the intensity of regulation in the road transport sector. Given all countries examined, with the exception of Italy, have introduced regulatory reforms aimed at reducing entry barriers in the road transport sector, our findings suggest that those reforms had a positive impact on the productive effect of infrastructure investments over the period 1980–2003.

These results are robust to different specifications of the regulation index and to the inclusion in the model of a variable proxing for congestion on the network, which in turn is found to negatively affect motorway elasticity. Moreover we found that motorway infrastructures may generate significant spillover effects, thus supporting the opportunity of a common EU transport infrastructure policy.

Our findings suggest that investments aimed at developing the motorway network might result to be more productive if accompanied by interventions which reduce congestion and by regulatory reforms in the road transport sector designed to reduce entry barriers.

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Appendix 1

In this section we provide a detailed description of how we derived some variables that have been used in this paper.

While data for gross output, labour hours and intermediate consumption were taken from the Groningen KLEMS dataset, capital stock data were taken from the OECD STAN dataset which however reports net capital stock data (in constant prices) at sector level (with the same level of aggregation of the KLEMS database) only for Denmark, Spain, France, Italy and Germany. For the UK we use data from the UK Office of National Statistics, while for the remaining countries the capital stock was built using data on gross fixed capital formation using a perpetual inventory method. In particular, the following PIM formula was used: $K_{it} = I_{it} + (1 - \delta)K_{it-1}$, where K is the capital stock and I the gross fixed capital formation. The PIM requires an initial (benchmark) level for the capital stock. Following Demetriades and Mamuneas (2000), we have first regressed the gross investment on a constant and a time trend in order to derive a predicted initial level of investment; then we have exploited the long-run relationship between investment and capital stock to construct a benchmark capital stock estimate: $K_{io} = \hat{I}_{it}/(\delta + g_i)$, where g_i is the growth rate of capital which was derived from the investment regression and δ is the depreciation rate, which was set equal to 8.5% for the utilities and the manufacturing sectors[33] and to values ranging between 7.6 and 9.9% for the remaining sectors.[34] To check the plausibility of our estimated capital stock figures, we followed the same procedure for those countries for which the STAN reported capital stock data, and we found out that the correlation coefficients between the STAN and our capital stock figures turned out to be pretty large. To take into account the possibility that countries might have experienced different economic cycles and that, as a consequence, during recessions the capital stock might not be fully utilized and, conversely, during booms it might be overused, we have computed a capacity utilization-adjusted capital stock series. In particular, following Griffith et al. (2004), we adjusted the capital stock series for capacity utilization, by regressing the gross output series on country-sectors fixed effects and a time trend: $Y_{ijt} = \alpha_{ij} + t$. The adjusted capital stock series is thus given by: $K^*_{ijt} = K_{ijt}\left[1 + \left(y_{ijt} - \hat{y}_{ijt}/\hat{y}_{ijt}\right)\right]$[35].

In order to correct for cross country differences in labour skills, we followed Harrigan (1999) and adjusted the hours worked in each country-sector by computing a translog index of three types of labour inputs, namely low, medium and high skilled workers:

$L_{ijt} = (HH_{ijt})^{s_{hijt}}(HM_{ijt})^{s_{mijt}}(HL_{ijt})^{(1-s_{hijt}-s_{mijt})}$, where *HH*, *HM* and *HL* stand for the hours worked in each country-sector-year combination by high, medium and low skilled workers, respectively; while s_h, s_m and $(1 - s_h - s_m)$ stand for the share of high, medium and low skilled labour, respectively, in the total labour share.

All monetary figures have been expressed in constant 1995 prices and converted to a common currency using appropriate PPP indices. In particular, for gross output we have used a set of industry 1997 PPPs provided by the University of Groningen.[36] For comparison, and as a robustness check, we also converted the data in national currencies by simply using an aggregate GDP PPP taken from the OECD and, reassuringly, none of the main results of the paper were driven by the particular PPP used. The capital stock data were converted into a common currency by using an investment PPP for 1995 taken from the EU AMECO database.

Other variables have been used in this paper as additional instruments, such as the head of government's political orientation, *POR* (leftwing, centre, rightwing), an Herfindhal index of government fragmentation, *HHI,* both taken from the Database of Political Institutions and *Tariffs*, which is a proxy for the existence of tariffs barriers to trade which is taken from the database of the Fraser index of Economic Freedom. As controls for possible effects of liberalization occurred in other sectors of the economy,

33 Lynde and Richmond (1993), using UK manufacturing data, used an yearly depreciation rate of about 7.2%; Brandt (2007), using cross country sector level data taken from the STAN database assumed a depreciation rate of 9%.

34 The results are however not sensitive on the exact depreciation rate assumed in building the capital stock.

35 We also experimented using the unadjusted capital stock series and our main results were unaffected.

36 As our national currency data were expressed in 1995 constant prices, but the PPP referred to 1997, we have modified the 1997 PPP by considering the relative sectorial output price inflation in each country with respect to the benchmark country (Germany) which occurred between 1995 and 1997.

we used other variables from the OECD regulatory database, such as *Eothers*, which is a simple average of the degree of entry barriers in all the other sectors mentioned in the Data section (utilities) other than road transport, *Credit* and *Lab*, taken from the Fraser Index of Economic Freedom database which try to capture the extent of regulation in the credit and labour markets.

Some regression equations rely on the transport intensity variable w_i. The use of a country-specific transport sector intensity is not ideal, as cross country differences might stem from country specific determinants, such as the development of the road network and/or regulation in the road transport sector, rather than from technological differences across sectors. In order to build a sector specific country-invariant road transport intensity variable, we follow a procedure outlined in Barone and Cingano (2008), and that was originally proposed by Ciccone and Papaioannou (2007), that helps alleviating possible endogeneity issues linked to the transport intensity variable.

First of all, we have taken the Eurostat 1995 Input-Output tables that were used to derive, for each country and each sector, the ratio between expenditure on land transport services and gross output (which is a proxy for land transportation costs as a percentage of the total),[37] w_{ij}. In order to control for the effect of both regulation and motorway network development on the transport intensity, we regressed w_{ij} on country dummies, sector dummies and sector dummies interacted with both country-level regulation in road transport and the motorway network infrastructure: $w_{ij} = a_i D_j + b_i D_i + c_i R_j D_i + d_i H_j D_i$, where D_j and D_i represent a set of country and sector dummies, respectively. To compute the sector specific country-invariant predicted values for the land transport intensity we set the value of the regulation index at its minimum sample level, the motorway stock at its median sample level as of 1990[38] and the country-specific averages to zero: $\hat{w}_i = \hat{b}_i D_i + \hat{c}_i R_{\bar{j}} + \hat{d}_i H_{\bar{j}}$, where $R_{\bar{j}}$ denotes the level of road transport regulation in the most liberalized country and $H_{\bar{j}}$ the median length of motorway network per km^2. In this way the sector specific fitted weights $\hat{w}_i$ that we have used in our regressions do not represent a land transport intensity that is influenced by regulation, by the motorway network or by other omitted country-specific variables.

To check the robustness of our results we simply computed for each sector a cross country average ($\bar{w}_i$) of the raw transport intensity measure (w_{ij}) and we have employed it as an alternative with respect to the fitted weights computed according to the above procedure ($\hat{w}_i$): reassuringly overall results are confirmed.

However, average weights ($\bar{w}_i$) are unlikely to provide a correct measure of each sector technological dependence on land transport because they differ from "true" weights by a idiosyncratic component: in this case their use could generate standard attenuation bias in our parameter estimates if the idiosyncratic component is unrelated to motorway infrastructure or regulation in the road transport sector; alternatively, if that component does depend on motorway infrastructure or regulation, then the bias is not clear a priori (see Ciccone and Papaioannou 2007, for a proof). To deal with

[37] Ideally, the expenditure on road transport services would have been preferable, but that was not available and therefore we turned to the expenditure on land transport (which includes also expenditure on railways services).

[38] Using a value different from the median would not alter our results.

this problem we have instrumented average sector weights ($\bar{w}_i$) with their predicted measures $\hat{w}_i$ (Wooldridge 2002) and we found that main results are virtually unaltered.

The motorway infrastructure stock in other countries G_{it} is relative to all countries belonging to the sample other than i. This is because there are countries that do not share borders (e.g. Germany and Italy) but are very close geographically and important trade partners. The spillover variable G_{it} has been built following a procedure suggested by Cohen and Morrison (2004): $G_{it} = \sum_j H_{jt} \frac{GDP_{it}}{GDP_{jt}}$, where H_{jt} is motorway Km in country *j* in year *t*, GDP is gross domestic product and f_{ij} is the ratio between the sum of imports and exports between country *i* and *j* and the sum of imports and exports between country *i* and all other countries in the sample; in other words, f_{ij} is a proxy for the importance that the motorway stock in country *j* has on country *i*, which we have assumed to depend positively on the (relative) importance of their trading relationship. The data for imports and exports are average values for the period 1988–2004 and are taken from the OECD STAN Bilateral Trade Database.

As suggested by Cohen and Morrison (2004), the ratio $\frac{GDP_{it}}{GDP_{jt}}$ is a multiplicative factor which "*reflects the relatively large effect that highway infrastructure stocks in state j, which shares a large value of goods shipments with state i, will have on state i's costs. Also, if state j, say, has a high level of economic activity relative to state i, it will constitute an overly large portion of* G_{it} *, unless this size effect is counteracted through the multiplication by* $\frac{GDP_{it}}{GDP_{jt}}$".

Table A1 Industries

Industry	Isic-rev 3
Agriculture, hunting, forestry & fishing (AHF)	01–05
Mining & quarrying (MIN)	10–14
Food, beverages & tobacco (FBT)	15–16
Textile, leather and footwear (TEXT)	17–19
Wood and wood products (WOOD)	20
Pulp, paper, printing &publishing (PULP)	21–22
Chemical, rubber, plastics & fuel (CHEM)	23–25
Other non metallic mineral (NON-MET)	26
Basic metal & fabricated metals (MET)	27–28
Machinery (MACH)	29
Electrical and optical equipment (ELE-EQ)	30–33
Transport equipment (TRANSP-EQ)	34–35
Manufacturing nec, recycling (MAN-REC)	36–37
Electricity, gas & water (EGW)	40–41
Construction (CONS)	45
Wholesale & retail trade (WRT)	50–52
Hotels & Restaurants (HR)	55
Transport and storage (TS)	60–63
Post & telecommunications (PT)	64
Financial intermediation (FIN)	65–67
Real estate, renting & business activities (REBA)	70–74

Table A2 Motorway network

	Km		Km/000Km^2	
	1980	2003	1980	2003
Austria	938	1670	30.7	54.7
Belgium	1251	1729	14.9	20.6
Denmark	504	1010	11.6	23.4
Spain	1923	10286	3.8	20.3
France	5287	10379	9.7	19.01
Italy	5900	6487	19.5	21.5
UK	2694	3611	11.03	14.7
Netherlands	1798	2308	43.3	55.6
Sweden	850	1591	1.8	3.5
Finland	185	653	0.5	1.9
Germany	9225	12044	25.8	33.7

Table A3 Price controls and barriers to entry

	Entry barriers						Price regulation					
	1980	1985	1990	1995	2000	2003	1980	1985	1990	1995	2000	2003
Austria	6	4.74	3.49	3.49	3.49	3.49	3	1.5	0	0	0	0
Belgium	6	6	5.01	3.49	3.49	3.49	–	–	–	–	0	0
Denmark	6	6	0.98	0.98	0.98	0.98	6	6	0	0	0	0
Finland	6	4.47	4.47	0.98	0.98	0.98	6	4.5	0	0	0	0
France	6	6	3.49	3.49	3.49	3.49	6	6	0	0	0	0
Germany	6	4.47	4.47	4.47	3.69	2.51	6	6	6	2.25	0	0
Italy	6	6	6	6	5.6	5.01	6	6	6	6	6	6
Netherlands	6	6	6	2.5	2.5	2.51	6	6	6	0	0	0
Spain	6	6	6	6	4.6	2.51	6	6	5.18	3.81	1.8	0
Sweden	2.95	2.95	1.97	1.97	1.97	1.97	0	0	0	0	0	0
UK	0.98	0.98	0.98	0.98	0.98	0.98	0	0	0	0	0	0

Table A4 Production function estimates (GMM-SYS)

	Model 1	Model 2	Model 3	Model 4	Model 5
l	0.163**	0.163*	0.13***	0.154***	0.144***
k	0.133***	0.124***	0.163***	0.140***	0.154**
m	0.674***	0.686***	0.626***	0.598***	0.619***
h	0.145***	0.355***	0.115***	0.132***	0.160***
h^2	–	−0.017*	–	–	–
hR	–	–	−0.002***	–	–
hwR	–	–	–	–	−0.133**
hR_{high}	–	–	–	−0.006***	–
hR_{low}	–	–	–	0.003*	–
m1 (p-value)	0.13	0.09	0.05	0.01	0.03
m2 (p-value)	0.61	0.18	0.16	0.28	0.24
Hansen J (p-value)	0.13	0.17	0.11	0.42	0.11

Table A4 Continued

	Model 1	Model 2	Model 3	Model 4	Model 5
Diff. Sargan (*p*-value)	0.87	0.89	0.82	0.93	0.65

***, ** and * stand for statistically significant at 1, 5 and 10%, respectively
m1 and m2 are Arellano–Bond tests for first and second order serial correlation
Instruments: l, m, k, h and R all dated T-2 for diff. eq
Instruments: Δl, Δm, Δk, Δh, ΔR, dated T-1 for level eq
Hansen: over identifying restrictions test; Diff. Sargan: validity test moment conditions for level equation
Country-sector fixed effects, time dummies, country dummies in all models

Table A5 Production function estimates: Robustness analysis

	GMM-SYS					IV-FD	GMM-DIFF
l	0.129**	0.143***	0.144***	0.182***	0.126***	0.301***	0.072*
k	0.164***	0.154***	0.153***	0.157***	0.166***	0.217**	0.158***
m	0.626***	0.625***	0.612***	0.585***	0.624***	0.217	0.522***
h	0.095***	0.134***	0.160***	0.159***	0.084***	0.170***	0.094***
hwR	−0.094	−0.111***	−0.130**	−0.117**	−0.125**	−0.440***	−0.079**
wR	−0.157	–	–	–	–	–	–
credit	–	−0.038	–	–	–	–	–
lab	–	0.012	–	–	–	–	–
Eothers	–	0.006	–	–	–	–	–
hcong	–	–	–	−0.035**	–	–	–
G	–	–	–	–	0.173*	–	–
m1 (*p*-value)	0.04	0.01	0.04	0.03	0.07	0.00	0.02
m2 (*p*-value)	0.37	0.12	0.24	0.36	0.38	0.48	0.45
Hansen: J (*p*-value)	0.12	0.09	0.10	0.31	0.12	0.46	0.15
Diff. Sargan (*p*-value)	0.78	0.61	0.86	0.83	0.82	–	–

***, ** and * stand for statistically significant at 1, 5 and 10%,
m1 and m2 are Arellano–Bond tests for first and second order serial correlation,
Hansen: over identifying restrictions test; Diff. Sargan: validity test moment conditions for level equations
GMM SYS instruments: l, m, k, h and R all dated T-2 for difference equation; Δl, Δm, Δk, Δh, ΔR, dated T-1 for level equations
GMM SYS: additional instruments. cong dated T-2 in column 4; G dated T-2 in column 5
IV-FD instruments: k and l dated T-2, m and h dated T-2 and T-3, POR, Tariffs and HHI
GMM-DIFF instruments: l, m, k, h and R all dated T-2 and T-3
Country-sector fixed effects and time dummies in all models + country dummies in column 1–2–4–5, country trends in column 3

Table A6 ARDL model: long run elasticities

ε_{yl}	0.122
ε_{yk}	0.202***
ε_{ym}	0.377***
ε_{yh}	0.124*
ε_{yhwr}	−0.44**

***, ** and * stat. sign. at 1, 5 and 10%

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