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INDUSTRY PRODUCTIVITY, INFRASTRUCTURE AND ROAD TRANSPORT LIBERALIZATION IN EUROPE

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Industry Productivity, Infrastructure and Road Transport Liberalization in Europe.

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Abstract

In this paper we assess the impact of both the highways network and the degree of regulation in the road freight sector on industry productivity by estimating a Cobb-Douglas production function on a panel of twenty one manufacturing and service sectors of eleven EU countries observed over the period 1980-2003. The production function estimates suggest that the highways network elasticity is positive, although we find that there are differences across sectors and countries. Furthermore, we find that the degree of liberalization in the road freight sector might play an equally important role in driving industry productivity: in particular, we find a non-linear effect of deregulation, which seems to be more effective when the process starts from an already more deregulated environment. Our results suggest that policymakers should consider deregulating the road transport sector as the gains in industry production might be as important as those stemming from further extensions in the infrastructure network.

Key Words: Transport, deregulation, public infrastructure.

JEL codes: L43 L91 L92 H54

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1 Introduction.

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. The basic intuition behind this is that improvements in public infrastructure (e.g. better roads) would be expected to raise the productivity of private inputs (say, by reducing the time and cost of transporting goods from factory to retail outlet), reducing the costs of production and raising the rate of total factor productivity growth. Nevertheless, the idea that public capital might be productive, does not necessarily imply that increasing public capital investment spending would lead to higher growth rates of GDP. In fact, conventional growth models à la Solow predict that higher investment (both private and public) would have effects only on the level of GDP, rather than on its rate of growth. However, more recent theories suggest that public investment might have long run effects on the rate of growth of GDP. For instance, a higher stock of public infrastructure might reduce costs of production by allowing greater specialization, thereby generating more output. In addition there may be further changes in factor markets and firm location decisions that allow the development of spatial clusters of economic sectors, thereby affecting innovation and allowing further reduction in costs. More recent theoretical and empirical research has sought to analyze the effects of public infrastructure in general equilibrium models that allow the joint addressing of issues such as the optimal provision of public infrastructure capital, taxation and technological progress. An overview of the theoretical literature on the links between economic growth and public infrastructure may be found in Afraz et al (2006) and in Agenor and Moreno (2006).

The seminal work by Aschauer (1989), which ascribed the US productivity decline of the 1970s to under-investment in infrastructures, has spurred empirical debate on the role of public capital in production[1]: Romp and De Haan (2007) and Afraz et al (2006) provide excellent surveys of this empirical literature.

The first wave of empirical work, based on Cobb-Douglas production functions estimated on time series data, confirmed Aschauer results. These early results (which were mainly based on US samples) were widely regarded as implausibly high and did not however find robust support in the studies that immediately followed.[2] In general, studies that were conducted at a more disaggregated level (such as by sector or state or region) tended to produce smaller estimates than these identified by studies employing national level data and also tended to display an interesting variability both across time and cross-sectionally. The main objections that were raised against the studies supporting the Aschauer findings were related to various weaknesses of the statistical analysis which did not properly address methodological issues related to reverse causality, simultaneity, non stationarity, functional form choice and measurement errors problems. More recent research has attempted to take account of some of these criticisms: production functions have been generally estimated after performing preliminary analysis on stationarity and cointegration, endogeneity and reverse causality have been addressed in several ways; moreover, the use of cost functions has become more common, especially in the case of studies using regional or industry data and vector auto-regressive models (VARs) have been increasingly used in the most recent studies that rely entirely on time series data.

Turning to EU evidence, most studies identify a positive effect of public infrastructure on output and productivity as well on costs. Only a few of them consider international comparisons (Evans and Carras (1994), Kamps (2004a), Kamps (2004b), Kamps (2005), Demetriades and Mamuneas (2000)), while the large majority are based on national samples. Focusing on studies which employ international

[1] Aschauer (1989) estimated a production function using US annual data for the period 1949-1985 and found that a one per cent increase in the stock of public capital infrastructure would have increased output by about 0.35 per cent

[2] See Gramlich's (1994) literature review for a comment on the early contributions.

samples, Kamps (2005) estimated an endogenous growth model which allows exploring the non linear link that might exist between economic growth and infrastructure capital and deriving the growth maximizing public capital stock: he found that, for a panel made up of the "EU-15", the elasticity of output with respect to public capital was about 0.20, a very similar result from Kamps (2004a) who reported an elasticity of about 0.22 from a production function estimated on a panel of OECD countries. However, separate regressions showed a large variability across countries: from negative (though insignificant) estimates for Portugal and Ireland, to positive, large and significant ones (Germany and The Netherlands, for instance)[3]. Demetriades and Mamuneas (2000) estimate a dynamic model of production based on an inter-temporal maximization framework using a panel of the manufacturing sector of 12 OECD countries (the G7 countries plus Australia, Belgium, Norway, Sweden and Finland) over the 1972-1991 period. Their specification of the model allows them to specify the effects of public capital on output both in the short and the long run: estimates seems to suggest that public capital tends to increase output, with elasticities varying form 0.36 in the UK to 2.06 in Norway and that these elasticities do not vary much in the long run.[4]

Some authors (Hulten, 1996; Calderon and Serven, 2005) have jointly considered the quality of public infrastructure (e.g. the percentage of paved roads or the electricity transmission and distribution losses) together with the infrastructure stock. Hulten (1996), for instance, argued that "how well you use the infrastructure is much more important than how much you have of it", and he found that, after controlling for the quality of roads in his cross country growth regressions, the impact of the

[3] Kamps (2004b) estimated a vector auto regressive model for 22 OECD countries over the period 1960-2001 to test the relationship which exists between macroeconomic variables like output, private capital, employment and public capital. He found that, in the long run, the elasticity of output with respect to public capital was positive and significant for twelve countries, negative and significant for one and not significantly different from zero for the remaining nine.

[4] Demetriades and Mamuneas (2000) found that public and private capital are substitutes, as labour and public capital. Moreover, they compare rates of returns for public infrastructure to its costs: in the short run, rates of return (gross of depreciation) range from 11 per cent in the UK to 27 per cent in Italy, while in the long run they range from 29 per cent in the US to 39 per cent in Italy. Comparing these figures with estimates for the user cost of public capital, they conclude that, in the long run, public capital had been under-provided in all countries, but that the "public infrastructure gap" had been falling over time for all countries and that for some it was even closed at the end of the sample period.

road stock turned out to be insignificantly different from zero. In this paper we take the Hulten's insight a step further, by considering whether the degree of liberalization of the road freight sector might have a role in driving industry productivity growth. The novelty of this study is that not only do we study the impact of transport infrastructure (highways) on industry production in a panel of eleven EU countries (over the period 1980-2003), but we jointly analyze the effects of regulation of the road transport sector. This industry is a typical industry which, in most countries, was traditionally protected from competition, with a myriad of inefficiently small operators, and that has experienced some form of deregulation in most OECD countries. The liberalization in the road transport sector is in fact just a part of a wider trend in most EU countries towards the introduction of privatization and regulatory reforms aimed to reduce barriers to entry and to stimulate competition in several sectors of the economy (product markets, factor markets and financial sectors).

Economic theory suggests different transmission mechanisms through which product market regulation may induce positive effects on economic performance: the reduction of X-inefficiency, the improvement in allocative efficiency and the incentivation of innovation are some examples. Nevertheless, those predictions need to be supported by empirical investigation, as they may run in the opposite direction if certain models assumptions are modified.

The effects of deregulation reforms on economic performance have been investigated empirically, but most of the existing literature is concerned with the labor market. A substantial literature has also developed on the effect of financial reform on a country's real performance for both developed and developing countries, while studies on the macroeconomic effects of goods market are more limited.

An excellent survey of this literature is found in Schiantarelli (2005) who presents a critical overview of the recent empirical contributions that use cross-country data to provide insights on the effect of product market regulations/reforms on a country's macroeconomic performance. Focusing on those studies which analyze productivity and output growth, the more exhaustive contribution is the work by Nicoletti and Scarpetta (2003) based on cross country data (18 OECD countries) for several industrial sectors (17 manufacturing and 6 service). Authors include a regulatory variable in the total factor productivity (TFP) equation and find evidence of a positive effect of privatization and entry liberalization on TFP growth, particularly for service sectors. Most studies which include regulation variables in TFP regressions find a negative effect of tighter regulation on TFP or per capita output growth (e.g. Cincera and Galgau (2005), Loayza et al. (2005)).[5]

In the following section we explain how we model the impact of transport infrastructure and its degree of regulation on production and illustrate our empirical strategy. Section three presents the data and is followed by the discussion of the empirical results. Section five concludes.

2 Model Specification and Empirical Strategy.

Let us assume that firms produce gross output according to the following Cobb-Douglas technology:

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

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:

[5] Griffith and Harrison (2004) suggest that reductions in the mark-up induced by regulatory reforms are associated with higher productivity growth in a panel of developed countries.

$$TFP_{ijt} = A_{ijt} TS_{jt}^{\eta} \quad (2)$$

where TS_{jt} represents the transport services in country j at time t and A_{ijt} represents the other neutral total factor productivity determinants not already accounted for in the model and η is the elasticity of gross output to the supply of transport services. Assuming that $TS_{jt} = H_{jt}^{\upsilon} \exp(R_{jt})^{\phi}$, and substituting this into equation 2, we get:

$$TFP_{ijt} = A_{ijt} TS_{jt}^{\eta} = A_{ijt} H_{jt}^{\upsilon\eta} \exp(R_{jt})^{\phi\eta} = A_{ijt} H_{jt}^{\varphi} \exp(R_{jt})^{\chi} \quad (3)$$

where H is the network of motorways and R is the degree of regulation in the road freight sector and φ and χ are parameters to be estimated. The basic idea underlying equation 3 is that transport services provided by the highways network reduce transport costs and make trade easier. This, in turn, making imports cheaper, tends to expose more sectors to foreign competition (the geographical scope of the "relevant" market tends to increase), favouring specialization and exploitation of economies of scale. However, alongside a well developed transport infrastructure network, of which the highways are an essential component in advanced countries,[6] an efficiently managed road transport sector can play an equally important role. In fact, a heavy regulated road freight sector characterized by high barriers to entry will be a sector insulated from healthy competition, which in turn might lead to low innovation and productivity growth and to the survival of many small and inefficient operators, depressing productivity in all sectors of the economy.

Let us further assume that the remaining component of total factor productivity is given by:

[6] See also Fernald (1999) and Cohen and Morrison (2004) who focused on the impact of the US highway stock on productivity dynamics.

$$A_{ijt} = \exp(t_t + \epsilon_j + u_{ij} + v_{ijt}) \tag{4}$$

where t_t is a time period effect, common to all cross sections in our sample that may account for common macroeconomic shocks or technological improvements and that is represented by a set of time dummies; u_{ij} represents a set of country-sector fixed effects representing, among other things, time invariant country-sector efficiency, v_{ijt} is an error term and ϵ_j represents a set of country specific effects.

After substituting equations 4 and 3 into equation 1 and taking logs, we get our estimated equation (where lower case variables denote natural logs):

$$y_{ijt} = \alpha k_{ijt} + \beta l_{ijt} + \gamma m_{ijt} + \varphi h_{jt} + \chi R_{jt} + t_t + \epsilon_j + u_{ij} + v_{ijt} \tag{5}$$

Before turning to the estimation strategy followed in this paper, we think we should spend a few words on ϵ_j, the country specific fixed effects that we have included in equation 5. Its role is to proxy for every time invariant effect, common to each sector in each country, that might drive productivity, like the degree of urbanization, the spread of population across the country, population density and so forth. Furthermore, the main variables of interest in this paper, the highways network h and the degree of liberalization R in the road transport sector, are "aggregate" variables, as they are defined at the country, rather than at the sector-country level.[7] It is well known that using "aggregate" variables when the dependent variable is defined at a lower level of aggregation (at the sector-country level in our case), although delivers unbiased and consistent estimates of the coefficients of interest, might lead to underestimate standard errors.[8] A possible solution would be to use standard errors robust to

[7] In other words, they display a variability through time and across country, but not across sectors in a single country.
[8] See Schiantarelli (2005) for comments in the case of cross country-sector level data when the liberalisation variable

some form of correlation across sectors belonging to the same country (see Moulton (1990) and Hoxby (2005) for a clear discussion of this issue) which are however unlikely to perform reasonably well when the number of groups is small, as it is the case in our sample, where not only is the number of groups very low (eleven), but it is even lower than the number of individuals within each group. As a sort of compromise, and to try to capture some within cluster correlation arising from omitted country-level variables, we decided to include a full set of country-specific dummy variables in equation 5.

Turning to the estimation strategy, equation 5 was 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 data and simultaneity issues. Popular approaches to estimate production function relationships like that in equation 5 have traditionally relied either on the conventional fixed effects approach or, more recently, on the Difference GMM approach of Arellano and Bond (1991). Both methods, however, have often been found to produce very unsatisfactory results in terms of parameters estimates and economies of scale:[9] in particular, the GMM difference approach, which removes time invariant heterogeneity by first differencing, and uses lagged (level) instruments to take into account simultaneity problems, have usually proven to provide too small estimates of the capital stock coefficient, especially when inputs and output are characterized by strong persistency or an almost random walk behavior (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 by using lagged first differences for the equation in levels on top

considered is country specific. See also Wooldrige (2003).

[9] The fixed effects approach has also been commonly found to exacerbate measurement error problems, resulting in parameter estimates possibly more biased than simple OLS (see Griliches and Mairesse, 1999).

of the usual lagged levels for the equation in differences. This method provides strong asymptotic and finite sample efficiency gains as well as reductions in the small sample biases which plague the GMM-difference estimator and allows researchers to incorporate time-invariant variables into the regression equation, which, in our case, is likely to be quite important given the potentially relevant role played by the country dummies in equation 5.

The use of datasets that pool together country as well as sector data over time is not new in the applied industrial organization literature (see, for instance, Griffith et al (2004) and Brandt (2007)). The main advantage is that we can increase the efficiency of the estimates by pooling together country as well as sector level information: the gain in efficiency often may be so large that it can often be worth the cost of some bias in the estimates.[10] Our approach is to treat the heterogeneity by allowing for country-sector fixed effects as well as for country specific dummies.[11] In addition to this, while we maintain homogeneity for the coefficients of the conventional production function inputs (capital, labour and intermediates), we have allowed, as a robustness check, for some degree of heterogeneity in the parameter φ, the elasticity of the motorways network, and χ, the marginal effect of the road sector liberalization index. In particular, the former was allowed to vary at the sector or country level. Following Alesina et al (2005), χ was allowed to vary according to the size of the regulatory reform.

As we said above, the series considered in equation 5 display quite large persistence: OLS and fixed effects estimates of autoregressive models for each of the series considered in equation 5 displayed autoregressive parameters ranging between 0.95 and 0.99. We therefore decided to undertake a fully-fledged empirical investigation of the time series properties of the series considered in equation 5. We run a battery of panel unit root tests, namely the Levin-Lin-Chu, Breitung, Maddala-Wo and

[10] Baltagi (2005) argues that MonteCarlo studies seem to suggest that estimators that do not confine the heterogeneity in the regression equation to the constant only tend to provide worse out of sample forecasts.

[11] For robusteness check we also verified that the main results hold also by inserting in the regression equation a full set of sector dummies.

Im-Pesaran and Shin tests, four of the most popular tests in the panel unit root literature. For all variables but the motorways network length, in all four 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.[12] The common feature of all these tests is that they all assume cross sectional independence: as a robustness check, we performed the Pesaran (2007) CADF test, and we could never reject the null hypothesis of the existence of a unit root in the case of m, l, y, k and R, while we had to reject it in the case of the motorways length network.[13] 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); therefore we proceeded on the assumption that the series are I(0) and that the GMM-SYS method is a valid estimation approach for our sample.

3 The 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, 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 built by researchers at the University of

[12] For the highways network length, two tests out of four lead us to reject the null hypothesis of the existence of a unit root. Results are available from the authors upon request.

[13] 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.

Groningen. From the KLEMS database we have taken data on gross output (Y), intermediates (M) and hours of work (H). Unfortunately, the KLEMS database reports capital stock data only for a limited number of countries. Therefore, for capital stock data we turned to 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, following Griliches and Mairesse (1999), we have exploited the long-run relationship between investment and capital stock to construct a benchmark capital stock estimate: $K_{io} = I_{it}/(\delta + g_i)$, 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[14], and to values ranging between 7.6% and 9.9% from the remaining sectors.[15] 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

[14] 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%.

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

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: $Y_{ijt} = \alpha_{ij} + t$, where t is a time trend. The adjusted capital stock series is thus given by: $(K * CU)_{ijt} = K_{ijt}(1 + \frac{Y_{ijt} - \widehat{Y_{ijt}}}{Y_{ijt}})$.[16]

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: $L_{ijt} = (HH_{ijt})^{s_{hijt}}(HM)^{s_{mijt}}(HL_{ijt})^{(1-s_{hijt}-s_{mijt})}$, 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.[17] 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.

As far as the infrastructure variables is concerned, the highway stock was proxied by the Km of highways network, which was taken from various publications by EUROSTAT,[18] while the public capital stock series was taken from the University of Kiel website and it is described in Kamps (2004).[19]

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

[17] 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.

[18] For Austria we could not find the relevant data for the 1981-1989 period, which were therefore reconstructed by linear interpolation.

[19] In terms of squared Kms, the motorways lenght and the total public capital stock display a correlation coefficient

Although it would have been interesting to compare the results with a monetary indicator of the stock of highways, comparable figures across countries do not exist. While there exist some data on gross investment in the overall road sector published by the ECMT, they are available only for shorter time spans and with significant gaps for some of the countries used in this study.

The first two columns in Table A2 report the length of the motorways network as of 1980 and 2003. As we can see, the motorways 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 motorways 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 motorways network by the respective country's area size (measured in squared Km): 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 across the country. For instance, the two Scandinavian countries in our sample both display very low ratios of motorways network Kms per squared Km, 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 motorways 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 utilities sectors, namely air transport, road transport, railways, telecom, gas, electricity and post. We will describe in some more detail the road transport indicators that we have used in this paper, while we refer to Conway et al (2006) for an exhaustive description of the OECD regulatory database. The degree of liberalization in the road sector of each country is derived from two subindices that are the main regulatory variables considered in this work.

of about 0.95.

The first index (PR), is a variable ranging from 0 (full liberalization) to 6 (very high regulation), that seeks to proxy for the importance of price controls in the road transport sector of each country. 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 freight services and, b) the government provides pricing guidelines to road freight companies. The second index, EB, is also a variable ranging from 0 to 6 that seeks to estimate the extent of barriers to entry in the road transport sector, and it was built by jointly considering five different issues: a) existence of a licence or permit to establish a national road freight 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; existence of professional bodies involved in specifying and enforcing entry regulations; d) existence of professional bodies involved in specifying or enforcing pricing guidelines or regulations.

Table A3 in the Appendix shows the behavior of EB and PO in the eleven countries in our sample in 1980 and 2003. As we can see, price regulation in the road freight transport sectors was relaxed in virtually all countries, with values of the index which in most countries fell to 0 over the 1980-2003 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.

The entry barrier index shows a higher variability across countries, although in most of them there is a clear tendency for EB to fall over time from values as high as 6. Again, 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.

In order to control for possible effects of liberalization occurred in other sectors, we used other variables from the OECD regulatory database. In particular, we have built the variable E_oth, which is a simple average of the degree of entry barriers in all the other sectors mentioned above (utilities) other than road transport.

Other variables used in this paper were taken from the Fraser Index on Economic Freedoms. In particular, the following variables were used: *Tariffs*, which is a proxy for the existence of tariffs barriers to trade; *Credit* and *Lab*, which try to capture the extent of regulation in the credit and labour markets. These three variables were available on a yearly basis only after the year 2000, while before that they were available only on a five years basis (1980, 1985, 1990, 1995) and therefore we were forced to use linear interpolation for the missing years.

4 Empirical results.

We estimated our baseline specification as in equation 5 with the GMM-SYS[20] approach and we report estimates results in Table A4. All estimates include country-sector fixed effects, time fixed effects as well as a full set of country dummies. Given the simultaneity issues affecting production functions estimates, private inputs have been instrumented with their own appropriate lags. The same issue applies to the highways network variable which has been instrumented with its own past values;[21] in fact, there might be a reverse causality problem between output and the highway network: since transport infrastructure investment might depend on the level of output, a productivity shock might be associated with a variation of the highways network, thereby causing biased estimates of the elasticity of output with respect to the highways network.[22] Moreover, the regulation index might be endogenous, due to the possibility that a productivity shock might be correlated with contemporaneous change

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

[21] Results presented in Table A4 used the highways network lagged one period as its own instrument, that is valid under a weak exogeniety assumption (i.e the highways network correlated with past shocks, but not with future and contemporaneous shocks to productivity). Results are robust using the highways network lagged two periods as its own instrument (which allows for correlation of the highways network with past and current productivity shocks).

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

in the regulatory environment; to tackle this issue we instrumented it with different variables. The instrumental variable that performed better was the level of the Freser index of Trade barriers;[23] nevertheless, the main results were not affected when we alternatively included both a dummy representing the head of government's political orientation, *POR* (leftwing, centre, rigthwing) and an Herfindhal index of government fragmentation, *HHI,* both taken from the Database of Political Institutions. 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, based on a model which considers the entry barrier index (*EB*) as a proxy for the degree of regulation in the road freight sector, show reasonable results for private input elasticities (suggesting weak decreasing returns to scale) and an elasticity of output with respect to the highway stock which is positive, statistically significant, and with a magnitude of about 0.12. Cohen and Morrison (2004) report an elasticity of manufacturing costs with respect to the highway stock of about 0.15 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.[24] We also obtained very similar results by including a relative measure of the highway infrastructure computed as the ratio of network kms and country's area size or as the ratio of network kms and population. Moreover, we checked whether the impact of the highway network on output is nonlinear: in fact it might be possible that once the main network has been laid out, further extensions might prove to be less productive. By augmenting the model with the square of the highways variable we found out that the coefficient of the

[23] A Difference Sargan statistics did not reject the validity of *Tariffs* as an additional instrument.

[24] See Afraz et al (2006). If, instead of the highways 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.

square term resulted to be negative but not statistically different from zero, as estimates in Column 2 show. However, by testing the significance of the elasticity of output with respect to the highways network at different percentiles of h, we find that such elasticity slightly declines as the highways network increases.[25] This implies that those countries that experienced substantial increases in the highways network over the sample period, such as Spain, exhibit a declining elasticity through time. The magnitude of the average country level elasticities that could be derived from the specification shown in Column 2 are broadly confirmed by a regression where we allowed for a country-specific elasticity of the highways network: the country specific elasticities are displayed in Table A5 and show that the average elasticities turned out to be higher in Finland and Sweden, with Italy, France and Spain displaying the smallest ones. Finally, in order to investigate whether the impact of the highways network on output differs across sectors we let the coefficient of the highways variable vary by interacting it with sector dummies: Table A6 reports the estimated elasticities. In particular, Table A6 displays some degree of heterogeneity across industries in the elasticity of output with respect to the highways network. For instance, the elasticity displays the highest values for sectors such as Transport and Storage, Electricity, Gas and Water, Finance, Wholesale and Retail Trade and slightly lower values for sectors such as Chemicals, Textiles, Post and Telecoms, Construction and Real Estate and Business Activities. The comparison with earlier literature is not easy as only a few papers assessed the impact of transport (or public) infrastructure using sector level data and reported the relevant industry-level elasticities;[26] furthermore, the methodology, as well as the industries considered, differed substantially across those few studies. Aviles Zugasti et al (2001) estimated a variable translog cost function for a panel of Spanish industries and found that the output elasticity of public capital had the highest values

[25] The elasticity ranges from 0.12 at the 25^{th} percentile, 0.11 at the median and 0.10 at the 75^{th} percentile, all significant at 10% or less.

[26] See Afraz et al (2006).

in sectors such as Construction, Metal and Non Metal Products and Chemicals, although authors failed to report the standard errors.[27] Although it is not straightforward to compare our results with the previous literature, our findings appear reasonable since we find the highest elasticities in those sectors where a priori one would expect a higher impact of highways network on output, such as Transport and Wholesale and Retail Trade.

The positive impact of the highways network on production is contrasted by the negative effects of barriers to entry in the road freight sector as the negative coefficient of the EB variable shows. In particular, a unit increase in the entry barriers index would reduce output by about 1.3 percent. As Table A4 shows, all countries in our sample,with the exception of Italy, have substantially deregulated the road freight sector by considerably reducing the barriers to entry: in particular, Denmark and Finland stand up as the countries that have almost eliminated barriers to entry; by way of contrast, the Italian road freight sector has maintained a high degree of barriers to entry. If Italy had to reduce entry barriers from its current (2003) level of about 5 to a value of 1 (as in the UK, Finland and Denmark), it could experience a transitory boost to its production level of about 5 per cent, once the regulatory change has fully displayed its effects. By estimating the baseline model augmented with the entry barriers index lagged one period (see Column 3) we found out that the contemporaneous and the lagged value of the index were jointly significant (with a p value of 0.07), thus suggesting that the effect of a regulatory change might need time to take place; in particular, the long run effect of entry barriers turned out to be about 1.3 percent (statistically significant at the 5% level).

In order to control for the possibility that the entry barrier variable might pick up the effects of the degree of regulation and liberalization in other sectors of the economy, we augmented the baseline

[27] Moreno et al (2003), in turn, found that, for a panel of manufacturing industries in Spain, the output elasticity of core public infrastructure was highest in sectors such as textiles, food, and electrical machinery, and lowest in sectors such as chemistry and metallic and non metallic products.

model with three variables, representing the degree of liberalization in the credit and labour market (see Data section) and a variable which proxies for the degree of entry barriers in the whole economy, which was built as an average of the OECD entry barriers indices in the gas, electricity, telecom, air and railways transport. Our basic results remain unchanged in terms of both magnitude and statistical significance of the main coefficients.[28] Furthermore, we experimented a different index of regulation, namely the price control index in the road freight sector (PR) and obtained estimates shown in Column 4: output elasticity with respect to the highways network is found to be slightly higher (0.16) and the effect of the price control index is very similar to that of the entry barriers index.[29]

We further investigated possible non-linearities of the effect of entry barriers on output by including the square of the entry barriers index: as Column 5 shows, the coefficient on the linear term remains significantly negative (with a value of about 5 per cent) while the coefficient on the squared term is positive (with a value of 0.6 per cent) and statistically significant. This result implies that the marginal effect of a reduction in entry barriers is higher for those countries characterized by a relative lower starting level of regulation, while it is negligible at the very beginning of the deregulation process.[30] In particular the impact of deregulation on output becomes statistically significant for values of the entry barrier index lower than 3.6: given that most countries (with the exception of Italy) have already reached this level of deregulation by the end of 2003, this result suggest that further deregulation interventions are expected to have positive impact on output.

Following Alesina et al (2005) we investigated whether the effect of a regulatory change in the road freight sector on production differs according to the size of the regulatory change. We constructed two dummy variables, *Weak* and *Strong*, the first equal to one for those countries (Austria, Belgium,

[28] Estimation results are available from the authors upon request.

[29] Similar results stems from a regression were we included the OECD composite road transport index derived as an average of the entry barriers and price control indices.

[30] See Alesina et al. (2005) for a similar result.

France, Italy, UK and Sweden) for which the regulatory index less than halved over the sample period, and the second equal to one for the remaining countries. We then estimated the model including the interaction between the regulatory index and the two dummies and obtained results shown in Column 6. According to estimates, although the magnitude of the coefficients is very similar, only the interaction with the dummy *Strong* is statistically significant. However, we can not reject the hypothesis that the two coefficients are equal to each other at conventional confidence level.

Finally we let the coefficient of the entry barriers index vary across country and we found results that were broadly in line with those suggested by the model displayed in Column 5.[31]

As a final robustness check, we estimated our baseline regression with an IV method and the results are shown in Column 7: the coefficient of the highways network drops somewhat, although it retains its statistical significance. More interestingly, the coefficient of the *EB* variable is virtually identical to that reported in Column 1, which gives further support to the estimates we obtained with the GMM-SYS approach.[32]

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 highways 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 drive industry productivity together with the road network itself. The liberalization of the road transport sector is part of a larger

[31] Results are available upon request.

[32] Furthermore, the standard errors used to compute statistical tests in Column 7 are robust to the existence of arbitrary forms of within country correlation. Using non-cluster robust standard errors would not appear to matter.

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. In this paper we add also to the literature on the impact of regulatory reform on economic performance by analyzing the impact of road transport sector liberalization on production on a wide range of industries across Europe. To the best of our knowledge, this is the first study on EU countries which jointly analyze 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 highway stock is about 0.12 and that it slightly declines as the highways network increases. This result suggests that those countries which undertook significant increases in the highways network over the sample period, such as Spain, exhibit a slightly declining elasticity through time. Moreover we found higher elasticity values in those sectors, such as Transport, Wholesale and Retail Trade, where it is more likely that an improvement in the highway network might have an higher impact.

Thus, improvements in transport infrastructure (as proxied by highways 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 products markets thereby encouraging competition, stimulating specialization and exploitation of economies of scale. However, as we expected, the positive effect of transport infrastructure investments on output might be depressed by the lack of a liberalized road transport sector.

In fact, the effect of road transport sector liberalization on production is found to be positive in all countries examined: given that most of them, with the exception of Italy, have introduced regulatory reforms aimed at reducing entry barriers in the road freight sector, our findings suggest that those reforms had a positive impact on industry production over the period 1980-2003. This result is found to be robust to different specifications of the regulation index and to the inclusion in the model of different measures of deregulation introduced in other industries. Moreover, we found that the deregulation of the road transport sector starts to display positive and significant effects on output when the regulation reform process has reached intermediate levels of implementations, as it is the case in all observed countries with the exception of Italy as of 2003: this implies that further deregulation of the road transport sector may deliver positive effects on output by stimulating competition in a sector which has been sheltered for a long time .

Overall results suggest that investments aimed at developing the highways network might result to be more productive if accompanied by regulatory reforms in the road transport sector designed to reduce entry barriers and price controls.

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6 Appendix

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 (WR)	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: Highways network

	Km		Km/Kmsq	
	1980	2003	1980	2003
Austria	938	1670	30.76	54.75
Belgium	1251	1729	14.91	20.61
Denmark	504	1010	11.69	23.82
Spain	1923	10286	3.80	20.82
France	5287	10379	64.62	103.05
Italy	5900	6487	19.58	21.53
UK	2694	3611	11.03	14.79
Netherlands	1798	2308	43.32	56.43
Sweden	850	1591	1.88	3.53
Finland	185	653	0.55	1.93
Germany	9225	12044	25.84	34.10

Table A3: Price controls and barriers to entry

	PR		EB	
	1980	2003	1980	2003
Austria	3	0	6	3.49
Belgium	–	0	6	3.49
Denmark	6	0	6	0.98
Spain	6	0	6	2.51
France	6	0	6	3.49
Italy	6	6	6	5.01
UK	0	0	0.98	0.98
Netherlands	6	0	6	2.51
Sweden	0	0	2.95	1.96
Finland	6	0	6	0.98
Germany	6	0	6	2.51

Table A4: Production function estimation results

	Model 1	Model2	Model3	Model 4	Model 5	Model6	Model7
l	0.130**	0.131**	0.131***	0.173***	0.172***	0.134**	0.122**
k	0.176***	0.172***	0.176***	0.166***	0.111**	0.174***	0.129***
m	0.603***	0.610***	0.602***	0.570***	0.633***	0.601***	0.678***
h	0.117***	0.183*	0.115***	0.159***	0.094***	0.119***	0.06**
h^2	-	−0.004	-	-	-	-	-
EB_t	−0.013**	−0.012*	−0.008	-	−0.052***	-	−0.014***
PR	-	-	-	−0.011**	-	-	-
EB_{t-1}	-	-	−0.006	-	-	-	-
EB_t^2	-	-	-	-	0.006**	-	-
$EB_t{*}Weak$	-	-	-	-	-	−0.014**	-
$EB_t{*}Strong$	-	-	-	-	-	−0.020	-
Fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	No
Country dummies	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Industry dummies	No	No	No	No	No	No	Yes
Time dummies	Yes	Yes	Yes	Yes	Yes	Yes	Yes
m1 (p-value)	0.03	0.03	0.08	0.01	0.06	0.02	-
m2 (p-value)	0.20	0.25	0.30	0.09	0.21	0.22	-
Hansen J (p-value)	0.22	0.15	0.20	0.30	0.23	0.18	0.30
Diff. Sargan (p-value)	0.99	0.99	0.99	0.99	0.98	0.99	-

***, ** and * stand for statistically significant at 1%, 5% and 10%, respectively. Models 1 to 6 are estimated with GMM-SYS, using the XTABOND routine in STATA while Model 6 is estimated with IV. m1 and m2 are Arellano-Bond tests for first and second order serial correlation, respectively;
Hansen is a test of overidentifying restrictions; Difference Sargan is a test on the validity of additional moment conditions used in the level equations.
The instruments used were l, m and k, all dated T-2 for the equations in difference and $\triangle l, \triangle m$ and $\triangle k$ dated T-1 for the equation in levels; h and Tariffs, dated T-1 for the equation in differences, and $\triangle h$, $\triangle$Tariffs for the equation in levels.
In turn l, m and k, all dated T-2; h and *EB* dated T-1, plus *Tariffs, PR, and HHI* were used as instruments in Model 7.

Table A5: Country Elasticities	
Austria	0.12^{***}
Belgium	0.10^{***}
Denmark	0.10^{**}
Spain	0.08^{***}
France	0.08^{***}
Italy	0.06^{**}
UK	0.09^{***}
Netherlands	0.10^{***}
Sweden	0.13^{***}
Finland	0.16^{***}
Germany	0.09^{***}

*** and ** stand for statistically significant at 1%, 5% , respectively

Table A6: Industry Elasticities

FBT	-0.02	MACH	0.07^{*}	PT	0.09^{***}
TEXT	0.08^{**}	ELE-EQ	0.05	AHF	0.03
WOOD	0.01	TRANSP-EQ	0.06^{*}	MIN	0.06
PULP	0.11^{**}	MAN-REC	0.05	CONST	0.09^{*}
CHEM	0.09^{**}	TS	0.12^{**}	WR	0.11^{**}
NON-MET	0.03	FIN	0.13^{***}	HR	0.03
MET	0.04	EGW	0.13^{***}	REBA	0.09^{**}

***, ** and * stand for statistically significant at 1%, 5% and 10%, respectively