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Taking Keller seriously: trade and distance in international R&D spillovers

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1. Introduction

Since the seminal work by Coe and Helpman (1995), a number of studies have investigated the role of trade in international R&D spillovers (e.g. Engelbrecht, 1997; Xu and Wang, 1999; Lejour and Nahuis, 2005; Lumenga-Neso *et al.*, 2005; Busse and Groizard, 2008; Coe *et al.*, 2009).

Among the many, Keller (1998) takes a critical stance and questions Coe and Helpman's (1995) empirical findings, by showing that the overall stock of the rest of the world R&D (i.e. the simple sum of the R&D produced abroad over time) performs better than an import weighted sum of it. This seems to suggest that knowledge diffusion is global and trade-independent, contrary to what suggested by Coe and Helpman.

The global vs. local nature of R&D spillovers is addressed again in Keller (2002), where he departs from the former claim and argues that spillovers are localized and their localized nature comes mainly from the fact that trade is geographically concentrated because of transport costs.[1] In Keller (2004) the perspective is broadened further: the localized nature of R&D spillovers comes to an important extent from the partially tacit nature of technology which requires the direct interaction among economic agents. (In this perspective, international trade can still be conducive to spillovers, as far as it facilitates face-to-face interactions which foster the diffusion of tacit knowledge.)

Thus, Keller starts from an hypothesis of "global pool" of technology (1998) and ends up with the idea that spillovers are geographically concentrated because they mainly depend on factors that are not directly related to international trade (2004). Given that both hypotheses are theoretically plausible, it remains an empirical issue whether R&D spillovers are local and to what extent they are trade-related. In this paper we address this issue by means of empirical specifications that nest both hypotheses.

In particular, in Section 2, we put forward a simple test for the hypothesis that R&D spillovers are global and trade-independent. In Section 3, we analyze the distinct roles of distance and trade, and consider the hypothesis that geographical proximity, with its impact on both trade and knowledge spillovers, is the actual determinant of R&D spillovers. Section 4 concludes summing up the main results.

2. Randomizing the randomizer: a simple test of the global pool hypothesis

Keller (1998) starts from Coe and Helpman's (1995) specification:

$$\log F_{it} = \alpha_i + \beta^d \log S^d_{it} + \beta^f \log S^f_{it} + \epsilon_{it} \tag{1}$$

[1] The theoretical model sketched in the Appendix of Keller's (2002) paper is an horizontal innovation model *à la* Rivera-Batiz and Romer (1991), where R&D externalities are rent spillovers (Griliches, 1979) and internal trade in intermediate goods is constrained by transport costs, exponentially increasing with geographical distance.

where the log of the Total Factor Productivity (TFP) of country i at time t (F_{it}) is regressed against a country dummy (α_i), the log of domestic R&D capital stock (S^d_{it}) and the log of foreign R&D stock of country i (S^f_{it}), and the latter is calculated as an import-weighted sum of the domestic R&D stock of the other countries ($S^f_{it} = \sum_j \frac{m_{ijt}}{\sum_j m_{ijt}} S^d_{jt}$).[2]

Keller re-estimates the equation substituting the import-weighted sum of R&D stock with the simple sum of the rest of the world stock of R&D ($S^f_{Kit} = \sum_{j \neq i} S^d_{jt}$) and shows that this gives rise to a higher point estimate of the TFP elasticity with respect to the foreign R&D and a better fitness of the regression. Hence, he concludes that "the composition of imports of a country plays no particular role in estimating a positive and significant impact from foreign R&D on domestic productivity levels" (1998, p.1479).[3]

Somehow at the risk of overstating Keller's position, this can be considered equivalent to an hypothesis of a "global pool" of technology. A simple way to test formally this hypothesis is to start from Keller's original specification (to which we add human capital (H_{it}) among the regressors as in Engelbrecht (1997)):

$$\log F_{it} = \alpha_i + \beta^h \log H_{it} + \beta^d \log S^d_{it} + \beta^f \log S^f_{Kit} + \epsilon_{it} \tag{2}$$

One can then write a more general, nonlinear model that nests Equation (2) as a specific case. In this work, we propose to consider the following nonlinear form:

$$\log F_{it} = \alpha_i + \beta^h \log H_{it} + \beta^d \log S^d_{it} + \beta^f \log \left(S^{fA}_{it} + \iota (S^f_{Kit} - S^{fA}_{it}) \right) + \epsilon_{it} \tag{3}$$

where $S^{fA}_{it} = \sum_{j \in \mathcal{A}_i \setminus \{i\}} S^d_{jt}$ and $\mathcal{A}_i$ is a subset of the set of countries, so that S^{fA}_{it} is the simple sum of the R&D stocks of the foreign countries belonging to a particular subset of the world, which can vary across countries.

If spillovers were truly global and trade-unrelated, so that all countries could absorb knowledge from a common and global pool, then the coefficient ι would not significantly differ from unity, no matter the partition of countries (i.e. the actual $\mathcal{A}_i$ for each country i). In such a case, the nonlinear model (3) would simplify in (2). Hence, given a set of subsets $\mathcal{A}$, one can estimate equation (3) by Nonlinear Least Squares (NLS), and then test the null hypothesis $\mathrm{H}_0 : \iota = 1$ against the alternative $\mathrm{H}_1 : \iota \neq 1$.[4]

To maintain the comparability with Keller (1998) and Coe and Helpman (1995), while extending the dataset, we test the hypothesis on a sample of 24

[2] On the problems entailed by such specification see Lichtenberg and van Pottelsberghe de la Potterie (1998) and Coe *et al.* (2009). Some of the econometric issues are addressed by Kao *et al.* (1999) and Edmond (2001).

[3] Keller (1998) also "randomizes" Coe and Helpman's (1995) measure by creating a weighted sum of foreign R&D stocks with random weights. However, Coe and Hoffmaister (1999) show that such weights are not truly random, but simple averages of the actual data with a random error.

[4] Following Davidson and MacKinnon (1999), we adopt an F-test that is not influenced by the nonlinear specification.

Table 1: Results of the F-test ($H_0 : \iota = 1$)

Experiment-wide significance level (α)	Significance level per comparison (α_0)	Number of F-tests with p-value $< \alpha_0$
0.01	0.0000100503	82
0.05	0.000051292	107
0.10	0.000105355	115

Convergence criterion met in 431 out of 1000 NLS estimates.

OECD countries over the period 1971-2004, using the data on R&D stock, human capital and TFP indexes from Coe *et al.* (2009).

We draw 24 random subsets $\mathcal{A}_i$ (one for each of the countries in the sample) out of the 2^{24} possible ones, with each country having probability 1/2 of belonging to the subset of any other country. The expected number of countries belonging to $\mathcal{A}_i \setminus \{i\}$ for each country i is therefore binomially distributed with expected value $23 \times 1/2 = 11.5$. We repeat this exercise 1000 times and estimate 1000 different models (3), assigning each time a different random subset $\mathcal{A}_i$ out of the 1000 previously drawn for each one of the 24 countries in the sample. With respect to each estimation we test the null $\mathrm{H}_0 : \iota = 1$.

Since we are looking at 1000 independent estimates, in order to reject the null with a significance level α for all the randomizations, we need to set a significance level of $\alpha_0 = 1 - (1 - \alpha)^{\frac{1}{1000}}$ for the single test and reject the null when at least one of the 1000 tests rejects it with α_0.[5] So, for instance, when $\alpha = 0.01$, the value of α_0 is 1.00503×10^{-5}.

Our results, reported in Table 1, show that the global pool hypothesis is strongly rejected by the data.[6] Out of 431 cases in which convergence is achieved, the F-test rejects 82 times (i.e. around 20% of the cases) the null at a significance level of $\alpha_0 = 1.00503 \times 10^{-5}$. We recall that, had the global pool hypothesis been correct, we would have rejected the null in none of the randomized partitions.[7]

Since the data strongly reject the hypothesis that R&D spillovers are global and trade-independent, in the next Section we analyze the impact of trade and geographical proximity on R&D spillovers and assess whether international trade appears positively related with R&D spillovers simply because it correlates with

[5] The probability of Type I error in at least one of the independent tests is:

$$\alpha = 1 - (1 - \alpha_0)^{1000}$$

where α_0 is the probability of Type I error in each test.

[6] NLS is estimated with Gretl 1.9.3 (`http://gretl.sourceforge.net/`), by the Levenberg-Marquardt algorithm with analytical derivatives and 2900 max iterations.

[7] The convergence criterion is not met in more than half of the repetitions mainly because the estimate of ι associated with some random partitions of the countries would take negative values (that render negative the argument of the log). On this basis one could rule out that the null hypothesis could be retained in the 569 non-converging cases and this would make our results even stronger. This notwithstanding, to lean on the cautionary side, we focus exclusively on the cases in which the convergence criterion is met.

geographical proximity, which is in fact the actual driver of the spillovers.[8]

3. Is trade proxying for geographical proximity?

To consider both distance and trade in the diffusion of knowledge and to test for the relevance of trade once accounting for geographical proximity, we introduce international trade in a slightly different version of the specification proposed in Keller (2002), where only distance appears as a determinant of knowledge spillovers:[9]

$$\log F_{it} = \alpha_i + \beta^h \log H_{it} + \beta^d \log S^d_{it} + \beta^f \log \left(\sum_{j \neq i} S^d_{jt} \mathrm{e}^{-\delta D_{ij}} \right) + \epsilon_{it} \qquad (4)$$

where D_{ij} is the geodesic distance between the capital cities of country i and country j, normalized so that the minimum smallest bilateral distance in the sample (that between Belgium and the Netherlands – 173.03 kilometers) is equal to one.[10] This can be seen as a nested version of a more general model that, following the specifications of Coe and Helpman (1995), also allows for a distance-unrelated role played by international trade in affecting R&D spillovers:

$$\begin{aligned} \log F_{it} = \alpha_i + \beta^h \log H_{it} + \beta^d \log S^d_{it} + \beta^f \log \left(\sum_{j \neq i} S^d_{jt} \mathrm{e}^{-\delta D_{ij}} \right) + \\ + \beta^{fm} m_{it} \log \left(\sum_{j \neq i} S^d_{jt} \mathrm{e}^{-\delta D_{ij}} \right) + \beta^m m_{it} + \epsilon_{it} \end{aligned} \qquad (5)$$

[8] Some hints of the superior performance of Coe and Helpman's (1995) trade-related measure of R&D over Keller's (1998) one emerge from the dynamic OLS (DOLS) estimates done using the enlarged sample of Coe *et al.* (2009), where leads and lags of first differenced independent variables are added to the original equation so as to obtain coefficient estimates with better limiting distribution properties (see Kao *et al.*, 1999). Results are reported in the Appendix A.

[9] In fact, Keller's (2002) original specification is at the industry level. Moreover, its exact aggregate version is slightly different from Equation 4 and would look like

$$\log F_{it} = \alpha_i + \tau_t + \beta \log \left(S^d_{it} + \sum_{j \neq i} S^d_{jt} \mathrm{e}^{-\delta D_{ij}} \right) + \epsilon_{it}$$

The differences between the latter and Equation 4 are that: i) we introduce human capital; ii) we separate domestic from foreign R&D stock because, following Coe and Helpman (1995), we allow for a systematic difference between the TFP elasticities of domestic vs. foreign R&D stock; iii) we do not include time dummies. We estimate the different specifications also with time dummies, but they always show a worse fit. The results are reported and discussed in Appendix B.

[10] As also noted by Keller (2002), this normalization amounts to a change in the measurement unit of distance and it does not affect elasticities. However, it *does* affect the size of δ. Therefore, because of the different minimum distance in the sample (that in Keller (2002) is the distance between Germany and the Netherlands, which is 3.34 times the distance between the Netherlands and Belgium), our estimates of δ cannot be directly compared with his. To do so, one would need to divide (multiply) his (our) value by 3.34.

where m_{it} is the share of imports on GDP in country i.

In Equation (5), the sum of the elasticities of productivity with respect to the foreign R&D stocks is not constant as in Equation (4), but it is an increasing function of the share of imports in the domestic economy:[11]

$$\sum_{j \neq i} \frac{\partial \log F_{it}}{\partial \log S^d_{jt}} = \beta^f + \beta^{fm} m_{it} \tag{6}$$

In turn, the marginal effect of the import share on the log TFP is an increasing function of the distance-weighted foreign R&D stock:

$$\frac{\partial \log F_{it}}{\partial m_{it}} = \beta^m + \beta^{fm} \log \left(\sum_{j \neq i} S^d_{jt} \mathrm{e}^{-\delta D_{ij}} \right) \tag{7}$$

If R&D spillovers were mainly trade-unrelated knowledge spillovers limited by geographical distance and/or rent-spillovers originated by trade in intermediates constrained by transport costs exponentially increasing with distance (as in the theoretical model by Keller (2002)), Equation (5) would simplify to (4). On the contrary, if international trade were an additional source of R&D spillovers (e.g. in case of trade-related knowledge spillovers), the marginal impact of trade on productivity – Equation (7) – and the effect of trade on the TFP elasticity of foreign R&D stocks – β^{fm} in Equation (6) – would be positive.[12]

Equation (5) does not impose any cross restriction on the estimated elasticity of TFP with respect to the import share and the foreign R&D capital stock because these latter and their interaction are all included in the estimation. However, as shown by Coe *et al.* (2009), panel unit root tests reject the null of unit root for m in all groups and Edmond (2001) shows that, in all the linear specifications where m is included as an independent regressor, Pedroni's (2004) test retains the null of no cointegration. Therefore, in order to reduce the risk of spurious regressions due to the improper inclusion of a stationary m, we estimate also the following specification:

$$\begin{aligned} \log F_{it} = {} & \alpha_i + \beta^h \log H_{it} + \beta^d \log S^d_{it} + \\ & + (\beta^f + \beta^{fm} m_{it}) \log \left(\sum_{j \neq i} S^d_{jt} \mathrm{e}^{-\delta D_{ij}} \right) + \epsilon_{it} \end{aligned} \tag{8}$$

where m does not appear as an independent regressor.

[11] The elasticity of the country i's TFP with respect to the R&D stock of country j is:

$$\frac{\partial \log F_{it}}{\partial \log S^d_{jt}} = \frac{\partial \log F_{it}}{\partial S^d_{jt}} \left(\frac{\mathrm{d} \log S^d_{jt}}{\mathrm{d} S^d_{jt}} \right)^{-1} = (\beta^f + \beta^{fm} m_{it}) \frac{S^d_{jt} \mathrm{e}^{-\delta D_{ij}}}{\sum_{c \neq i} S^d_{ct} \mathrm{e}^{-\delta D_{ic}}}$$

[12] We focus our discussion on the estimates of these combined parameters rather than on the individual estimated parameters because the individual coefficients of interacting variables do not typically have a direct interpretation.

Table 2: Estimation results (Pooled data 1971-2004 for 24 countries: 816 observations)

	I - eq. 4	II - eq. 5	III - eq. 8
β^h	0.527***	0.515***	0.490***
	(0.0516)	(0.0466)	(0.0483)
β^d	0.038***	0.029***	0.044***
	(0.0084)	(0.0073)	(0.0078)
β^f	0.168***	0.057**	0.128***
	(0.0153)	(0.0147)	(0.0167)
δ	0.046***	0.069***	0.064***
	(0.0116)	(0.0104)	(0.0085)
β^{fm}		0.358***	0.036***
		(0.0525)	(0.0079)
β^m		-4.370***	
		(0.7805)	
$\beta^f + \beta^{fm}\bar{m}$ [a]		0.165***	0.139***
		(0.0139)	(0.0153)
$\beta^m + \beta^{fm} \log \bar{S}^f$ [b]		0.401***	
		(0.0588)	
$\beta^{fm} \log \bar{S}^f$			0.487***
			(0.0962)
AIC	-1624.8	-1766.1	-1681.9
BIC	-1493.1	-1624.9	-1545.5

[a] $\bar{m} \approx 0.3$ is the median import share in the sample.

[b] $\bar{S}^f = \left(\sum_{j \neq i} \bar{S}^d_{jt} \mathrm{e}^{-\delta \overline{D}_{ij}}\right)$, where the bar stands for the median value in the sample.

Unreported country dummies. Bootstrapped standard errors robust to serial and cross sectional dependence (or heteroskedasticity-asymptotically robust standard errors – variant HC1 – when larger than bootstrapped ones) in parentheses. Coefficient significance based on bootstrapped two-tailed confidence intervals. Significance levels: * 10%; ** 5%; *** 1%.

On the one hand, the latter specification necessarily imposes some cross restrictions on the estimated elasticity of TFP (see Lichtenberg and van Pottelsberghe de la Potterie, 1998; Coe and Hoffmaister, 1999). On the other hand, when δ is taken as given, it is similar to the specifications estimated and discussed in Coe *et al.* (2009) and panel cointegration cannot be rejected by the data.

The results of our estimations are reported in Table 2.[13] As inference based on bootstrapped standard errors tends to be more reliable, for all the specifications we compute bootstrapped standard errors and report the coefficient significance based on bootstrapped confidence intervals. To be conservative, we report asymptotically robust standard (HC1) errors when they are larger than the bootstrapped ones. In particular, we employ the panel moving blocks bootstrap, proposed and discussed by Gonçalves (2011) in the context of linear panels with

[13] The data on countries' distance are borrowed from CEPII.

individual fixed effects, with a block size equal to 2.[14]

Results confirm Keller's (2002) finding that R&D spillovers decline with the geographic distance between sender and recipient countries and provide further evidence that productivity spillovers of R&D are geographically localized. Given the estimates of δ, the implied "half-life distance of technology", that is the distance at which only half of the foreign country's R&D stock is domestically available,[15] ranges roughly from 10 to 15 times the distance between Belgium and the Netherlands (i.e. 1,750 to 2,600 kilometers). In models II and III (respectively Equations (4) and (5)), the estimated decay rate increases, pointing to a possible underestimation in model I because of the omission of trade among the regressors, which is positively correlated with TFP and negatively correlated with distance.[16]

What is important is that, even when distance is accounted for, trade remains an additional source of R&D spillovers. In particular, the marginal impact of the countries' import share on productivity calculated at the median value of the distance-weighted foreign R&D stock – $\beta^m + \beta^{fm} \log \bar{S}^f$ in model II (Equation (5)) and $\beta^{fm} \log \bar{S}^f$ in model III (Equation (8)) – is positive and significant. The elasticity of the distance-weighted R&D stock positively depends on the import-share (β^{fm} is positive and significant in both models II and III).

Hence, geographic distance is not the only "actual" driver of R&D spillovers and there are important trade-related knowledge spillovers which can significantly increase the impact of foreign R&D on domestic productivity.

4. Conclusions

Notwithstanding a rich literature on the international R&D spillovers, whether the effects on productivity of R&D are global or localized and trade-related or trade-unrelated still remains an open empirical issue as evidence has been provided in support of both hypotheses. A case in point is the work of Wolfgang Keller, who presents contrasting findings in different works (i.e. Keller, 1998, 2002, 2004) adopting different (and non nested) empirical specifications.

By using the enlarged sample of Coe *et al.* (2009) and adopting two empirical specifications that nest the various models proposed by Keller, we test the hypotheses that spillovers are global (localized) and trade-unrelated (trade-related).

[14] In the panel moving blocks bootstrap (MBB), a standard MBB is applied to the vector containing the individual observations at each point in time. Like the standard MBB, the panel MBB is robust to serial dependence of unknown form as long as it satisfies a mixing type condition. Furthermore, because it does not resample the individual observations directly, this bootstrap is also robust to any arbitrary form of cross sectional dependence. See Gonçalves (2011) for details.

[15] Because of the exponential specification, this distance is assumed to be constant, and it is equal to $\ln 2/\delta$.

[16] Our estimates are however in general lower than those reported in Keller (2002), where the half-life of technology ranges from 162 to about 1,200 kilometers in his preferred specification.

In particular, we carry out a simple test based on a randomization of the original model in Keller (1998) to test the hypothesis of a "global pool" of technology, which we strongly reject. On this basis we investigate further the determinants of these localized spillovers.

By nesting the model proposed in Keller (2002) to analyze the impact of geographical distance on R&D spillovers in a more general model that is able to account for both trade and distance, we test the hypothesis that international trade appears positively related with R&D spillovers simply because it correlates with geographical proximity, which is instead the "actual" driver of the spillovers. This would be true if R&D spillovers were mainly trade-unrelated knowledge spillovers and trade-related spillovers were in fact rent-spillovers originating from trade in intermediates limited by geographic distance that increases transport costs.

We reject also this hypothesis and show that R&D spillovers depend on both geographical distance and international trade. Even when distance is accounted for, trade remains an important source of productivity spillovers and considerably increases the elasticity of TFP with respect to the distance-weighted foreign R&D stock.

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Appendixes

A. Dynamic OLS estimates

To fully exploit the cointegrating properties of the series, one can calculate the dynamic OLS (DOLS) estimates where leads and lags of first differenced independent variables are added to the original equation to obtain coefficient estimates with better limiting distribution properties (see Kao *et al.*, 1999).

In the following table, we report the DOLS estimates (including unreported country-dummies) and compute Newey-West standard errors (in parenthesis) for the original specifications proposed by Coe and Helpman (1995) (Model I in this appendix) and Keller (1998) (Model II in this appendix). On the basis of the Akaike Information Criterion, we include one lead and two lags of first differenced variables for model I (720 observations) and a lag of order two of first differenced variables in model II (744 observations).

	$\log H$	$\log S^d$	$\log S^f_{CH}$	$\log S^f_K$
I	0.710571^{***}	0.060832	0.126415^{**}	
	(0.146073)	(0.043420)	(0.051484)	
II	0.607976^{***}	0.066295		0.107212^{*}
	(0.180654)	(0.046932)		(0.0612691)

The elasticity of TFP with respect to $\log S^f_{CH}$ is higher than with respect to $\log S^f_K$ and its standard error lower. This hints to a superior performance of Coe and Helpman's (1995) measure over Keller's (1998).

B. Estimates with time dummies

As a robustness check, we re-estimate all the models discussed in Section 3 with the inclusion of time dummies. Results are reported in Table 3.

Information criteria are all worse than those of the respective models without time dummies. The qualitative results are all confirmed. The point estimates with and without time dummies are fairly similar for models III and I. The main difference is in model II the higher point estimates of β^f and β^{fm}, which produce a strongly higher estimate of the elasticity of the distance-weighted foreign R&D stock (valued at the median import share: $\beta^f + \beta^{fm}\bar{m}$). Such estimate is not in line with the estimates of this elasticity in all the other models, although also associated with a high standard error.

Because of the cointegration issues, possibly magnified by the inclusion of time dummies in model II;[17] the lower information criteria with respect to the models without time dummies; and the fact that the estimated elasticity in all the other models are fairly similar, this result should not be given too much emphasis.

[17] Levin *et al.*'s (2002) and Im *et al.*'s (2003) panel unit root tests both retain the null of a unit root in all groups for the residuals of the regressions of $\log F$ and $\log S^d$ on time dummies; while they both reject the null at the 1% significance level for the residuals of the regression of m on time dummies.

Table 3: Estimation results with unit and time dummies (Pooled data 1971-2004 for 24 countries: 816 observations)

	I	II	III
β^h	0.542^{***}	0.643^{***}	0.515^{***}
	(0.0586)	(0.0717)	(0.0602)
β^d	0.038^{***}	0.021^{***}	0.046^{***}
	(0.0081)	(0.0079)	(0.0086)
β^f	0.130^{**}	0.348^{***}	0.148^{***}
	(0.0787)	(0.1360)	(0.0861)
δ	0.053^{***}	0.047^{***}	0.063^{***}
	(0.0174)	(0.0066)	(0.0145)
β^{fm}		0.539^{***}	0.044^{***}
		(0.0655)	(0.0091)
β^m		-6.875^{***}	
		(0.9019)	
$\beta^f + \beta^{fm}\bar{m}$ [a]		0.509^{***}	0.161^{***}
		(0.1385)	(0.0859)
$\beta^m + \beta^{fm} \log \bar{S}^f$ [b]		0.487^{***}	
		(0.0661)	
$\beta^{fm} \log \bar{S}^f$			0.585^{***}
			(0.0897)
AIC	-1570.500	-1754.303	-1636.154
BIC	-1283.531	-1457.925	-1344.480

[a] $\bar{m} \approx 0.3$ is the median import share in the sample.

[b] $\bar{S}^f = \left(\sum_{j \neq i} \bar{S}^d_{jt} \mathrm{e}^{-\delta \overline{D}_{ij}}\right)$, where the bar stands for the median value in the sample.

Unreported unit and time dummies. Bootstrapped standard errors robust to serial and cross sectional dependence (or heteroskedasticity-asymptotically robust standard errors – variant HC1 – when larger than bootstrapped ones) in parentheses. Coefficient significance based on bootstrapped two-tailed confidence intervals. Significance levels: * 10%; ** 5%; *** 1%.

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