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Bernd Hayo and Stefan Voigt

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Coordination: Bernd Hayo • Philipps-University Marburg
School of Business and Economics • Universitätsstraße 24, D-35032 Marburg
Tel: +49-6421-2823091, Fax: +49-6421-2823088, e-mail: hayo@wiwi.uni-marburg.de

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Bernd Hayo* and Stefan Voigt **

*MACIE, Philipps-Universität Marburg

** University of Hamburg

Marburg Centre for Institutional Economics • Coordination: Prof. Dr. Elisabeth Schulte
c/o Research Group Institutional Economics • Barfuessertor 2 • D-35037 Marburg

Phone: +49 (0) 6421-28-23196 • Fax: +49 (0) 6421-28-24858 •
www.uni-marburg.de/fb02/MACIE • macie@wiwi.uni-marburg.de

Philipps Universität Marburg

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Bernd Hayo*

and

Stefan Voigt**

* University of Marburg, School of Economics & Business, Marburg Centre for Institutional Economics (MACIE), Universitaetsstr. 24, 35037 Marburg, Germany, Phone: +49-6421-2823091, Fax: +49-6421-2823088, Email: hayo@wiwi.uni-marburg.de.

**University of Hamburg, Institute of Law & Economics, Johnsalle 35, 20148 Hamburg, Germany, Phone: +49-40-428385782, Fax: +49-40-428386794, Email: stefan.voigt@uni-hamburg.de. and CES-ifo, Munich.

The Puzzling Long-Term Relationship Between *De Jure* and *De Facto* Judicial Independence

Abstract:

We study the long-term and dynamic relationship between *de jure* and *de facto* judicial independence using a large panel dataset covering 50 countries over a period of 50 years. Our analysis shows a negative relationship between these variables, a sharp contrast to the prevailing theoretical view in the literature. However, the magnitude of the relationship is small. The negative association between the two variables is driven by OECD countries, whereas a positive one can be found for non-OECD countries. We discover no evidence of reverse causality running from *de facto* to *de jure* judicial independence.

Keywords: Judicial independence, *de facto*, *de jure*, long-term panel data analysis, cointegration, Granger causality

JEL classification: D 72, D 78, K 42

1 Introduction

The law and economics literature makes an important distinction between *de jure* and *de facto* judicial independence (*de jure*JI and *defactoJI*). It seems straightforward to assume that increases in the former will be followed by increases in the latter. However, findings in the scarce empirical literature are not so straightforward. Using cross-sectional data, Hayo and Voigt (2007) find that *dejureJI* and *defactoJI* are positively related and that *dejureJI* is the single most important predictor for *defactoJI*. However, the magnitude of this relationship is small. Melton and Ginsburg (2014) report that none of the conventional variables used to proxy *dejureJI* are significantly correlated with *defactoJI*. Indeed, a figure in their paper (2014, 189) suggests that *defactoJI* might even cause *dejureJI* to adjust, instead of the other way around as commonly assumed.

In this paper, we tackle this issue from a different perspective. Although it is interesting to compare differences in *dejureJI* and *defactoJI* across countries, the lack of a time dimension makes inferences problematic. Here, we use panel data analyses to study the long-term relationship between *dejureJI* and *defactoJI* over 50 years and across 50 countries. This investigation is possible due to the development of time-based indicators for *dejureJI* and *defactoJI*. Hayo and Voigt (2014, 2016) use and extend the Comparative Constitutions Project (Elkins et al. 2009) and derive a time-varying indicator for *dejureJI* based on factor analysis; Linzer and Staton (2015) come up with a latent variable measurement model combining eight extant indicators to map out *defactoJI* across time. In the latter's context, missing data are a big problem and they deal with it by employing Bayesian methodology.

Combining these two variables in one dataset, we generate the largest possible balanced panel with a length of 50 years (1956–2006), which, coincidentally, contains 50 countries too. Note that the sample is not representative of the world, as more than one third of the countries became OECD members before 1973 and some regions are not adequately covered. See the Appendix for a list of the countries and summary statistics.

2. The Long-Term Relationship Between *De Jure* and *defactoJI*

We commence our analysis by running a static random effects panel data regression, where we use *dejureJI* to explain *defactoJI*.[1] Model 1 in Table 1 shows a significantly negative relationship between the two variables, which is in stark contrast to the theoretical view in the literature. A one standard deviation increase in *dejureJI* is associated with a 0.13 standard deviations decrease of *defactoJI* and the average marginal elasticity is -0.08. Thus, the absolute size of the effect is quite small, suggesting that the linkage between the two variables is weak. This result is in line with Melton and Ginsburg's (2014) conclusion.

Table 1: Explaining *defactoJI* using *dejureJI*: static long-run regressions

	1 All countries RE	2 All countries GLS	3 OECD countries RE	4 Non-OECD countries RE
Constant	0.56*** (0.06)	0.59*** (0.01)	1.08*** (0.04)	0.31*** (0.03)
DejureJI	-0.27*** (0.01)	-0.03* (0.02)	-2.26*** (0.11)	0.20*** (0.06)
Test variables	$Chi^2(1)$=25***	$Chi^2(1)$=3*	$Chi^2(1)$=471***	$Chi^2(1)$=13***
Test AC(1)	F(1,49)=4894***	n.a.	F(1,17)=809***	F(1,31)=4178***
Observations	2,500	2,500	900	1,600
Countries	50	50	18	32
Years	50	50	50	50

Notes: RE=random effects estimator. GLS=generalised least square estimator with an autocorrelated error of order 1 and allowing for heteroscedastic panels. Test variables=Wald test of all included variables. Test AC(1)=Wooldridge (2002) test for first-order autocorrelation. *, **, and *** indicate significance at the 10%, 5%, and 1% level, respectively.

Are these findings robust? Since we find substantial autocorrelation, we re-estimate the model using GLS with an autocorrelated error of order one and allowing for heteroscedastic

[1] Note that all results reported here hold when estimating fixed effect models.

panels. This model (Model 2 in Table 1) shows that the qualitative result remains, but the quantitative effect is even smaller. In Models 3 and 4, we split up the sample into OECD and non-OECD countries. This sample split proxies for different degrees of institutional development. We observe that the negative coefficient arises from the relationship in the OECD sample. Now the effects are still inelastic but no longer negligible: the reduction in terms of *defactoJI* standard deviations resulting from a one standard deviation hike in *dejureJI* approaches unity and the average elasticity is -0.42. The reverse is found for non-OECD countries. Model 4 shows a significantly positive relationship between *dejureJI* and *defactoJI*. The effects of a one standard deviation increase and the average elasticity are 0.08 and 0.09, respectively. Again, absolute and relative effect sizes are small.

The preceding conclusions assume that both variables are stationary or at least cointegrated. Panel unit roots have very weak power and, thus, a long time series is essential for valid inference. This is one of the reasons why we restrict the sample to countries with 50 years of data. Table 2 studies the time-series behaviour of our dataset. Employing various tests, we find clear evidence of non-stationarity. The cointegration tests suggest that the variables are cointegrated, at least for most of the panels.

Table 2: Testing for unit roots and cointegration

<table>
<tr><td></td><td colspan="3">Unit root tests</td><td colspan="3">Cointegration tests</td></tr>
<tr><td></td><td>Levin-Lin-Chu</td><td>Breitung</td><td>Im-Pesaran-Shin</td><td>Kao</td><td>Westerlund: Some panels</td><td>Westerlund: All panels</td></tr>
<tr><td>DefactoJI</td><td>Adjusted t*=-0.4</td><td>Lambda=11</td><td>W-$\bar{t}$=0.1</td><td rowspan="2">D-F t*=3***</td><td rowspan="2">Variance ratio=6***</td><td rowspan="2">Variance ratio=1</td></tr>
<tr><td>DejureJI</td><td>Adjusted t*=9</td><td>Lambda=-1</td><td>W-$\bar{t}$=2</td></tr>
</table>

Notes: See notes to Table 1. Unit root tests: Levin-Lin-Chu (2002); Breitung (2000); Im-Pesaran-Shin (2003). Cointegration tests: Kao (1999); Westerlund (2005). All tests use demeaning and include a trend.

This finding leads us to compute error-correction terms (EC) based on the results from Table 1. Accounting for the potentially dynamic nature of the relationship of interest, we run EC models using the first difference of the JI variables and employing five lags. The outcome in Table 3 shows that the ECs are highly significant in all specifications. Thus, the relationships estimated in Table 1 appear to be long-term equilibria, deviations from which affect *defactoJI*'s short-term adjustment in a stabilising way.

Table 3: Explaining Δ*defactoJI* using Δ*dejureJI*: EC model

	1 All countries RE	2 All countries GLS	3 OECD countries RE	4 Non-OECD countries RE
Constant	0.001*** (0.0001)	0.001*** (0.0001)	0.001* (0.0003)	0.001*** (0.0003)
Δ*defactoJI*$_{t-1}$	1.088*** (0.021)	0.912*** (0.024)	1.027*** (0.036)	1.101*** (0.027)
Δ*defactoJI*$_{t-2}$	-0.295*** (0.032)	-0.036 (0.033)	-0.150*** (0.054)	-0.330*** (0.040)
Δ*defactoJI*$_{t-3}$	0.045 (0.032)	0.052 (0.033)	-0.042 (0.054)	0.082** (0.041)
Δ*defactoJI*$_{t-4}$	-0.009 (0.032)	-0.056 (0.032)	0.100* (0.053)	-0.052 (0.040)
Δ*defactoJI*$_{t-5}$	-0.016 (0.021)	-0.006 (0.024)	-0.076** (0.039)	0.015 (0.027)
Δ*dejureJI*$_{t-1}$	-0.001 (0.011)	-0.006 (0.006)	0.097*** (0.029)	-0.006 (0.012)
Δ*dejureJI*$_{t-2}$	-0.007 (0.011)	0.003 (0.006)	-0.006 (0.029)	-0.001 (0.012)
Δ*dejureJI*$_{t-3}$	-0.004 (0.011)	-0.001 (0.006)	0.016 (0.029)	-0.002 (0.012)
Δ*dejureJI*$_{t-4}$	-0.001 (0.011)	-0.004 (0.006)	-0.003 (0.029)	0.002 (0.012)
Δ*dejureJI*$_{t-5}$	-0.001*** (0.012)	-0.002*** (0.006)	0.026 (0.028)	-0.004 (0.014)
ECM$_{t-1}$	-0.003*** (0.001)	-0.001*** (0.0002)	-0.006*** (0.0003)	-0.007*** (0.001)
Test variables	$Chi^2(11)$= 6658***	$Chi^2(11)$= 6279***	$Chi^2(11)$= 2765***	$Chi^2(11)$= 4105***
Test AC(1)	F(1,49)=57***	n.a.	F(1,17)=11***	F(1,31)=42***
Granger causality	$Chi^2(5)$=0.5	$Chi^2(5)$=2.2	$Chi^2(5)$=13**	$Chi^2(5)$=0.4
Weak exogeneity	$Chi^2(1)$=0.1	$Chi^2(1)$=0.04	$Chi^2(1)$=1.2	$Chi^2(1)$=1.4
Observations	2,200	2,200	792	1,408
Countries	50	50	18	32
Years	44	44	44	44

Notes: See notes to Table 1. Granger (1969) causality: joint test involving 5 lags of dejureJI. Weak exogeneity: test of weak exogeneity of dejureJI with regard to de fact JI (Johansen 1992).

Do we find short-term Granger-causality from *dejureJI* to *defactoJI*? Not generally, but in the case of OECD countries, we detect a significant test outcome. The absolute effect of the short-term dynamic reaction is small, though. A 1 percentage increase in *dejureJI* leads to an increase of 0.13 percentage points in *defactoJI*. Thus, for OECD countries, there is a positive relationship between the two variables in the short run and a negative one in the long run.

Finally, we analyse whether the causal relationship might run from *defactoJI* to *dejureJ*, as suggested by the figure in Melton and Ginsburg (2014, 189). We use a VAR-type setup (Johansen 1992) to test for weak exogeneity of *defactoJI* with regard to *dejureJI*. None of the tests reject the null of weak exogeneity (see Table 3), which supports the view that *defactoJI* adjusts to *dejureJI* and not the other way around. Moreover, we find no evidence of Granger-causality running from *defactoJI* to *dejureJI*.

3. Conclusion

Using two recently published new indicators for *defactoJI* and *dejureJI*, we study their long-term relationship as well as their short-term dynamics. In contrast to the theoretical view in the extant literature, we find that the relationship between the two variables is negative and weak in terms of magnitude, in line with findings by Melton and Ginsburg (2014). In our sample, the negative relationship holds only for OECD countries, whereas we discover a positive relationship outside of the OECD. Employing a different methodology, Gutmann and Voigt (2018) report a similar finding for EU countries.

We find evidence of cointegration between the two variables, which, according to the Engle-Granger representation theorem (Engle and Granger 1987), can be interpreted as the existence of long-term equilibria. In the case of the OECD countries, we observe different dynamics: the long-term relationship is negative, but short-term changes in *dejureJI* positively affect changes in *defactoJI*. Finally, we discover no evidence of reverse causality, i.e., that *dejureJI* is influenced by *defactoJI*.

Appendix

A) Sample countries (*=OECD countries in 1972)

Albania, Argentina, Australia*, Austria*, Belgium*, Bolivia, Brazil, Bulgaria, Canada*, Sri Lanka, Chile, China, Taiwan, Colombia, Cuba, Denmark*, Dominican Republic, Ecuador, El Salvador, Ethiopia, France*, Germany*, Greece*, Guatemala, Haiti, Honduras, Iceland*, India, Indonesia, Ireland*, Italy*, Japan*, Lebanon, Luxembourg*, Mexico, Netherlands*, Nicaragua, Norway*, Panama, Paraguay, Peru, Poland, Portugal*, Spain*, Syria, Thailand, Egypt, United States*, Uruguay, Venezuela.

B) Variable descriptions (annual data, 1956–2006)

Variable	Source	Obs.	Mean	Standard Deviation	Min	Max
DejureJI	Normalised indicator from Hayo and Voigt (2016)	2,500	0.14	0.16	0	1
DefactoJI	Indicator from Linzer and Staton (2015)	2,500	0.52	0.32	0	1
DejureJI OECD	OECD: Normalised indicator from Hayo and Voigt (2016)	900	0.10	0.09	0	0.4
DefactoJI OECD	OECD: Indicator from Linzer and Staton (2015)	900	0.85	0.22	0	1
DejureJI Non-OECD	Non-OECD: Normalised indicator from Hayo and Voigt (2016)	1,600	0.16	0.18	0	1
DefactoJI Non-OECD	Non-OECD: Indicator from Linzer and Staton (2015)	1,600	0.34	0.20	0	0.9

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