The Economics of State Fragmentation: Assessing the Economic Impact of Secession - Addendum

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4 July 2016

THE ECONOMICS OF STATE FRAGMENTATION: ASSESSING THE ECONOMIC IMPACT OF SECESSION

Addendum

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VIVES Discussion Paper 55. This version:
July 5, 2016

Abstract

Reynaerts and Vanschoonbeek (2016) propose a semi-parametric procedure to estimate the economic impact of secession, finding empirical evidence that declaring independence significantly lowered per capita GDP in newly formed states. To demonstrate that these findings appear to hold irrespective of the estimation procedure employed, this addendum formulates a parametric approach to estimate the independence dividend. Our preferred parametric specifications comprise a dynamic, quasi-myopic model of per capita GDP dynamics that controls for country and year fixed effects, the rich dynamics of GDP, finite anticipation effects and a vector of alternative growth determinants. The results indicate that declaring independence reduces per capita GDP by around 15-20% in the long run. These results are qualitatively confirmed when we use non-regional secession waves to instrument for local incentives to secede.

Keywords: Independence dividend; panel data; dynamic model; generalized method of moments; bootstrap-based bias correction; instrumental variable regression

JEL Classification: C14, C32, H77, O47

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The authors would like to thank Geert Dhaene, Joep Konings, Klaas Staal, David Stadelmann and participants to the VIVES Research Seminar for valuable comments and suggestions on earlier versions of this paper.

1 Introduction

In this addendum, we expound upon Reynaerts and Vanschoonbeek (2016) and follow a parametric approach to identify the independence dividend. After outlining the general economic model, a next section explains the model selection procedure. Subsequently, we present our baseline results and perform some robustness checks. Finally, a last section provides instrumental variable estimates of the independence dividend.[1]

2 Estimation strategy

Based on the literature review discussed in Reynaerts and Vanschoonbeek (2016), we are interested in estimating the following underlying economic model:

$$y_{i,t} = \beta_0 + \beta_1 NIC_{i,t} + \sum_{p=1}^{P} \beta_{1+p} E_t[NIC_{i,t+p}] + \sum_{q=1}^{Q} \alpha_q y_{i,t-q} + \theta X_{i,t} + \lambda Z_{i,t} + \delta_i + \eta_t + \epsilon_{i,t} \quad (1)$$

where $y_{i,t}$ is the log of per capita GDP of country i at time t, $NIC_{i,t}$ is an independence dummy equal to 1 for each NIC in the first 30 years after it gained independence and 0 otherwise, $X_{i,t}$ is a $(1 \times X)$ vector of observed controls, Z_{it} is a $(1 \times Z)$ vector of time-varying unobserved growth determinants, δ_i captures the I country fixed effects, η_t denotes T year fixed effects and the error term, $\epsilon_{i,t}$, collects all random transitory shocks to per capita GDP. The model also includes up to Q lags of the dependent variable to allow for persistency in GDP dynamics. E_t denotes the expectation taken with respect to the information set at time t, reflecting the potential presence of anticipation effects up to P years prior to a declaration of independence.[2] The coefficient of interest is β_1, as this is the coefficient that will capture the economic impact of secession: a statistically significant and positive estimate would indicate the presence of independence gains while a statistically significant negative estimate would signify an independence cost.

As is standard, we deal with the potential problem of time-constant omitted variable bias by removing time-constant unobserved heterogeneity, which may include hard-to-quantify variables such as cultural norms or political institutions, using the standard within estimator. To eliminate common per capita GDP trends that may be correlated with declaring independence, year dummies are included in each specification as well.

To deal with the potential presence of finite anticipation effects, we estimate a quasi-myopic model replacing the expectation of independence p years prior to the actual independence declaration, $E_t[NIC_{i,t+p}]$, with leads of the independence-dummy, $NIC^p_{i,t}$. More specifically, $NIC^p_{i,t}$ is a dummy variable equal to 1 if country i will declare independence at time $t+p$, such that the *ex ante* effect of secession at time $t+p$ is estimated by β_{1+p}.

[1] Further details about data sources and construction is offered in Reynaerts and Vanschoonbeek (2016).

[2] The typical pre-secession dip in per capita GDP observed in NICs may reflect anticipation effects. If so, Malani and Reif (2015) show that their omission generates omitted variables bias.

Denoting the within transformation of y_{it} by $\tilde{y_{i,t}} = y_{i,t} - \frac{1}{T_i}\sum_{t=1}^{T_i} y_{i,t}$, our baseline linear regression specification related to equation (1) can be summarized as follows:

$$\tilde{y}_{i,t} = \beta_1 N\tilde{I}C_{i,t} + \sum_{p=1}^{P} \beta_{1+p} N\tilde{I}C^{p}_{i,t} + \sum_{q=1}^{Q} \alpha_q \tilde{y}_{i,t-q} + \theta \tilde{X}_{i,t} + \eta_t + \tilde{\epsilon}^{Z}_{i,t} \tag{2}$$

where the error term, $\tilde{\epsilon}^{Z}_{i,t} = \lambda \tilde{Z}_{i,t} + \tilde{\epsilon}_{i,t}$, also subsumes variation over time in the z unobserved growth determinants and the contents of $\mathbf{X}$ are detailed below.

In order for this relation between per capita GDP and the independence dummy to be dynamically stable, the sum of the coefficients of the lagged dependent variables ($\sum_{q=1}^{Q} \alpha_q$) measuring the persistence in per capita GDP should be smaller than one. If this would not be the case, a one-time decision to secede would cause explosive changes in per capita GDP trajectories. To check whether explosive behavior can be ruled out, an appropriate F-test can be employed (Acemoglu, Suresh, Restrepo, & Robinson, 2014). This is a relevant issue, as the β_1-coefficient in the regression model summarized in equation (2) only estimates the immediate economic impact of secession or, equivalently, the effect of declaring independence if independence is expected to last for one single period. Computing the effect of a permanent secession, β, involves accounting for the serially correlated nature of per capita GDP which, under dynamic stability, boils down to:

$$\hat{\beta} = \frac{\hat{\beta}_1}{1 - \sum_{q=1}^{Q} \hat{\alpha_q}} \tag{3}$$

Finally, a well-known problem with the model specified in equation (2) lies in the so-called Nickell (1981) bias, which is inherent to the estimation of auto-regressive fixed effects models and arises due to the mechanical correlation of the lagged dependent variables and the error term. The fact that this bias declines in the length of the time series (T), however, suggests that dynamic panel bias should be negligible given the long time-dimension of our dataset.[3] Nevertheless, to account for dynamic panel bias, we also consider two alternative estimators which are unbiased for finite T. First, we use a one-step generalized methods of moment (GMM) estimator for dynamic panel models that was introduced by Arellano and Bond (1991) and further developed by Arellano and Bover (1995). Exploiting the absence of residual auto-correlation, country fixed effects are first removed by applying the forward orthogonal deviations transformation to all regressors while dynamic panel bias is mitigated by subsequently instrumenting the transformed lagged dependent variables using suitable lagged levels of y_{it}.[4] In order to avoid an over-fitting bias (Roodman, 2006) the number of lagged instruments is restricted to a maximum of 5. In addition, we consider the iterative bootstrap-based bias correction procedure for the fixed effects estimator in dynamic panels proposed by Everaert and Pozzi (2007).[5]

[3] In the model lacking additional controls, each country is observed 57 times on average.
[4] The GMM estimator is implemented by Roodman's (2006) *xtabond2*-command in Stata13.1.
[5] The bias correction is implemented by De Vos and Ruyssen's (2015) *xtbcfe*-command in Stata13.1.

3 Model selection

The inclusion of lagged dependent variables and anticipation dummies in equation (2) serves to mitigate endogeneity and omitted variable bias concerns, but requires us to determine of the 'optimal' number of lagged dependent variables, Q^*, as well as the 'optimal' number of anticipation dummies, P^*. To do so, we build on Stock and Watson (2012) who suggest two approaches to choosing the optimal order of an auto-regressive model: through a statistical significance criterion or using an information criterion. This first approach starts with a model including many lags and subsequently sequentially checks whether the coefficient on the last lag (or trailing lag) significantly differs from 0. If not, the last lag is dropped and the procedure is repeated until the 'optimal' number of lags is determined. By a similar procedure, in the second approach, the number of lags to be included is selected according to the model minimizing the AIC criterion. As this second approach penalizes model complexity, in contrast to the first approach, it is expected to yield a more parsimonious model, reducing estimation uncertainty but potentially omitting valuable information contained in more distant lags.

Since we need to determine both Q^* and P^*, we consider the statistical significance and the information criterion approach in two different sequential scenarios. In the first sequential scenario, $Q^* \rightarrow P^*$, we start from the bivariate fixed-effects model to determine the optimal number of lagged dependent variables (Q^*) first and only after this decide on the number of anticipation terms (P^*) to be included. In the second sequential scenario, $P^* \rightarrow Q^*$, the order is reversed and we first decide on P^* and subsequently determine Q^*.

To choose the optimal number of terms in either sequential scenario, we will consider two trails: in the first, which we will refer to as the *long* trail, we start with a 'long' model including 10 sequential terms and subsequently sequentially shorten the model at the last term - the trailing lag or the trailing anticipation term - whenever our model selection criterion dictates so. In the second, so-called *short*, trail we start with our fixed effects bivariate model and subsequently rely on either model selection criterion to determine whether a more extensive model specification is to be preferred.

Finally, to select the optimal specification resulting from this exercise, based on the principle of parsimony, we adopt the most concise candidate model (i.e. the model which minimizes $P^* + Q^*$) lacking first-order auto-correlation in the residuals.[6] Although deciding on the number of lags and anticipation terms to include is always and by necessity somewhat arbitrary, this approach at least has the added benefit of providing some empirical evidence suggesting a well specified estimation model.

The results of this exercise are reported in table 1. As it turns out, including 4 lags of the dependent variable sufficiently accounts for past GDP dynamics while relevant anticipation effects are limited to the first 2 pre-secession years.

[6]The intuition behind this approach is that residual auto-correlation in a dynamic model *"often indicates that the functional form has not been completely specified"* (Wooldridge, 2009, p. 412).

Table 1: **Quasi-myopic, autoregressive models of per capita GDP**

Model selection criterion	*Sequence*	*Trail*	*Result*	*AR(1)*
Statistical significance (t-test)	$Q^* \rightarrow P^*$	*short*	$Q^* = 4; P^* = 2$	.21
Statistical significance (t-test)	$Q^* \rightarrow P^*$	*long*	$Q^* = 8; P^* = 2$	.01
Statistical significance (t-test)	$P^* \rightarrow Q^*$	*short*	$Q^* = 4; P^* = 10$	.21
Statistical significance (t-test)	$P^* \rightarrow Q^*$	*long*	$Q^* = 4; P^* = 10$	.21
Information criterion (AIC)	$Q^* \rightarrow P^*$	*short*	$Q^* = 1; P^* = 2$	.00
Information criterion (AIC)	$Q^* \rightarrow P^*$	*long*	$Q^* = 1; P^* = 9$	.00
Information criterion (AIC)	$P^* \rightarrow Q^*$	*short*	$Q^* = 1; P^* = 10$	.00
Information criterion (AIC)	$P^* \rightarrow Q^*$	*long*	$Q^* = 1; P^* = 10$	.00

Note: All estimations include the independence dummy as well as both country and year dummies. The first three columns specify the model selection procedure. Column 4 indicates the resulting candidate model. AR(1) reports the p-value of the Wooldridge (2009) test for the absence of first-order auto-correlation in the residuals of the candidate model. Standard errors are robust for heteroskedasticity and clustered at the country-level.

4 Baseline results

The most straightforward situation to derive a reliable parametric estimate of the economic impact of secession would occur when, conditional on past GDP dynamics, anticipation effects as well as country and year fixed effects, the decision to secede and its expectation are entirely unrelated to any other past or present growth determinant while, at the same time, dynamic panel bias can be safely ignored due to the time-span of our dataset. To formalize these ideas, assume that the following conditions are satisfied

Condition 1

$$\mathbb{E}\left[N\tilde{I}C_{i,t}\tilde{\epsilon}^{ZX}_{i,s}\right] = 0 \quad \forall\ t \geq s$$
$$\mathbb{E}\left[N\tilde{I}C^{1}_{i,t}\tilde{\epsilon}^{ZX}_{i,s}\right] = 0 \quad \forall\ t \geq s$$
$$\mathbb{E}\left[N\tilde{I}C^{2}_{i,t}\tilde{\epsilon}^{ZX}_{i,s}\right] = 0 \quad \forall\ t \geq s$$

Condition 2

$$\mathbb{E}\left[\tilde{y}_{i,t-q}\tilde{\epsilon}^{ZX}_{i,s}\right] = 0 \quad \forall\ t,s\ \&\ \forall\ q \in \{1,\ldots,Q^*\}$$

where it is understood that $\tilde{\epsilon_{i,t}}^{ZX} = \theta\tilde{X}_{i,t} + \lambda\tilde{Z}_{i,t} + \tilde{\epsilon}_{i,t}$.

Although, at first sight, these might seem rather strong assumptions, Wooldridge (2009) emphasizes that including lagged dependent variables provides a useful, though crude, strategy to control for more general omitted variable bias.

In this case, the standard within regression of the model outlined in equation (2) lacking any additional controls, in effect leaving the **X**-matrix empty, would provide us with a reliable estimate of the β_1-coefficient. Column (1) in the top panel of table 2 reports the primary results of such an exercise while the full results are relegated to table A1. Similar to all the other results we present in this section, per capita GDP trajectories turn out to be highly persistent (in this case, $\sum_{q=1}^{4}\alpha_q = 0.974$). Nevertheless, an F-test clearly rejects the null hypothesis of dynamic instability. More importantly, the estimated

β_1-coefficient is significantly negative and implies that per capita GDP is reduced by approximately 0.4% in the year of independence. The serially correlated nature in per capita GDP trajectories, however, implies that this adverse effect will accumulate in the ensuing years and converge to a long-run payoff of -0.16. In other words, on average, a declaration of independence significantly reduces per capita GDP levels by an estimated 16% in the long run. In addition, the results also indicate the presence of significantly negative *ex ante* effects associated with a future state break-up, lowering per capita GDP by an estimated 1.6% and 2.4% in the last two years prior to independence. Anticipation effects thus appear to partially explain the pre-secession growth dip typically observed in future NICs.

Nevertheless, the literature review contains several indications that condition 1 may not hold in practice. One complication might arise since NICs, upon gaining independence, tend to liberalize their trade regime to compensate for reduced market size. Secessions also appear to be associated with increased pressures to democratize, accelerating the transition towards more democratic political institutions. Our analysis has also abstracted from political and economic relations among countries. This might be problematic, as Alesina, Spolaore, and Wacziarg (2000) argue that economic integration induces political disintegration. To control for the economic significance of evolutions in these, and other, alternative growth determinants we also consider a specification including several control variables. More specifically, we consider a compact control vector, $\mathbf{X}^{\mathbf{compact}}$, capturing contemporary evolutions in democracy, trade openness, military conflict, human capital, surface area, population differentials and political as well as macroeconomic instability. In addition, we include dummy variables reflecting (historical) membership of the EU, the OECD, the NATO, ASEAN, MERCOSUR, the AU and the group of oil exporting countries.

By taking these covariates out of the error term, it now becomes more likely that condition 1 holds with respect to the decision to secede and its prior anticipation. This, however, comes at the cost of having to impose similar exogeneity assumptions on each additional control (Wooldridge, 2009). We can express this idea formally by rewriting condition 1 such that we now require:

Condition 1a

$$\begin{aligned}
&\mathbb{E}\left[N\tilde{I}C_{i,t}\tilde{\epsilon}^{Z}_{i,s}\right] = 0 && \forall\ t \geq s \\
&\mathbb{E}\left[N\tilde{I}C^{1}_{i,t}\tilde{\epsilon}^{Z}_{i,s}\right] = 0 && \forall\ t \geq s \\
&\mathbb{E}\left[N\tilde{I}C^{2}_{i,t}\tilde{\epsilon}^{Z}_{i,s}\right] = 0 && \forall\ t \geq s \\
&\mathbb{E}\left[\tilde{x}_{i,t}\tilde{\epsilon}^{Z}_{i,s}\right] = 0 && \forall\ t \geq s\ \&\ \forall\ x\ \in\ \mathbf{X}^{\mathbf{compact}}
\end{aligned}$$

where the error term, $\tilde{\epsilon}^{Z}_{i,t} = \lambda \tilde{Z}_{i,t} + \tilde{\epsilon}_{i,t}$, now contains less information.

The results of this compact control specification are reported in column (2) of the top panel of table 2. Table A1 demonstrates that control variables tend to have the expected sign and that most of them are statistically significant. With regard to the coefficients of interest, although a lot of observations are lost due to the inclusion of these alternative growth determinants, the overall pattern is fairly similar. The results do suggest a slightly larger short-run cost of secession, which is now estimated to cut per capita GDP by 0.5% in the first post-independence year, and a more adverse anticipation effect, which is limited to the first pre-independence year and which reduces GDP per capita by 3.9% in the run-up to state break-up. A slightly smaller persistence in per capita GDP implies that the long-run, per capita cost of secession now roughly amounts to 12.4%, somewhat smaller but still broadly similar to our previous estimate.

Finally, to check whether the implosion of multi-ethnic states is more costly in economic terms we also include interactions of the independence dummy with a dummy indicating ex-Soviet states as well as a dummy identifying former members of Yugoslavia. In addition, Qvortrup (2014) suggests that declaring independence by referendum is particularly conducive to peaceful political settlements such that successful independence referendums potentially mitigate at least some of the economic costs of secession. To take these procedural aspects into account as well, we add the interaction between the independence dummy and a dummy indicating the occurrence of a successful independence referendum prior to the official declaration of independence.

The primary results stemming from the inclusion of this more extensive vector of control variables are reported in the last upper column of table 2. As can be seen, controlling for these additional growth determinants does not qualitatively affect the results although, compared to the baseline results, the long-run negative payoff of state fragmentation is now estimated to be somewhat larger, reducing per capita GDP by roughly 16% in the long run. We find some evidence that the implosion of multi-ethnic states is more adverse in economic terms, as the estimated economic cost associated with of the break-up of the Soviet union severely exceeds its sample-average counterpart, although former Yugoslavian countries do not appear to have suffered disproportionally from the implosion of Yugoslavia. Finally, declaring independence by referendum does seem to mitigate these reported secession costs as the coefficient associated with the interaction of the referendum- and the independence-dummy is statistically significant and positive.

Recall that the standard within estimator ignores dynamic panel bias by imposing condition 2. As mentioned, this might be reasonable given the long time-span of our data set. To verify whether this effectively is the case, we repeat the previous exercise using the orthogonal deviations GMM estimator instead, which is employed here to purge dynamic panel bias. The corresponding estimates are reported in columns (1) to (3) in the middle panel of table 2. Reassuringly, overall, the GMM estimates for all the coefficients involved correspond well with the associated within estimates, confirming the intuition that our prior results were not driven by dynamic panel bias. Correspondingly, the estimated payoffs of state fragmentation in both the pre- and the post-independence period turn out to be fairly similar and suggest a significant long-term effect of a 9.8% to 15.3% decrease in per capita GDP due the decision to secede. Interestingly, we find confirmation that ex-Soviet members suffered disproportionally from the breaking up of their mother country and that independence-by-referendum seems preferable from an economic point of view. Importantly, we test the absence of second order auto-correlation in the differenced error terms, which underpins the GMM estimator, and fail to find evidence that this exclusion restriction is not met.

Alternatively, Everaert and Pozzi (2007) develop a simulation-based approach to estimate and remove dynamic panel bias through an iterative bootstrap-procedure. By analogy, the results of applying this estimator instead are reported in columns (1) to (3) in the bottom panel. Once again, bias-corrected estimates closely track their corresponding within estimates. That being said, a slightly larger estimated persistence in per capita GDP trajectories leads to a somewhat more pronounced estimated long-run secession cost, which now hovers around 19% to 32% in per capita terms. Finally, the bootstrap-based bias correction model does indicate that declarations of independence that are preceded by a succesful independence referendum tend to do better while no confirmation is found that former Soviet or Yugoslav states faced above-average costs of state break-up.

Table 2: **Parametric estimates of the economic impact of secession**

		Within estimates	
Independence dummy	-0.004**	-0.005**	-0.007**
	(0.002)	(0.003)	(0.003)
Ex ante effect (t - 1)	-0.024***	-0.039***	-0.038***
	(0.008)	(0.013)	(0.014)
Ex ante effect (t - 2)	-0.016***	-0.007	-0.006
	(0.005)	(0.010)	(0.011)
Independence dummy × Soviet dummy			-0.024**
			(0.011)
Independence dummy × Yugoslav dummy			0.009
			(0.019)
Independence dummy × referendum dummy			0.014***
			(0.003)
Observations [# countries]	11,128 [192]	8,318 [187]	8,318 [187]
Adjusted R-squared	0.980	0.978	0.978
Country + Year FE	yes	yes	yes
Control vector	none	$\mathbf{X}^{\mathbf{compact}}$	$\mathbf{X}^{\mathbf{extensive}}$
$\sum_{q=1}^{4} \alpha_q$ [F-test <1]	.974 [0]	.956 [0]	.956 [0]
Long-run effect [p-value]	-.161 [.037]	-.124 [.038]	-.16 [.013]
		Deviations GMM estimates	
Independence dummy	-0.006**	-0.006**	-0.007**
	(0.003)	(0.003)	(0.003)
Ex ante effect (t - 1)	-0.020**	-0.039***	-0.038***
	(0.009)	(0.013)	(0.013)
Ex ante effect (t - 2)	-0.011**	-0.006	-0.005
	(0.006)	(0.010)	(0.011)
Independence dummy × Soviet dummy			-0.022*
			(0.011)
Independence dummy × Yugoslav dummy			0.009
			(0.018)
Independence dummy × referendum dummy			0.014***
			(0.004)
Observations [# countries]	10,935 [192]	8,131 [187]	8,131 [187]
Country + Year FE	yes	yes	yes
Control vector	none	$\mathbf{X}^{\mathbf{compact}}$	$\mathbf{X}^{\mathbf{extensive}}$
$\sum_{q=1}^{4} \alpha_q$ [F-test <1]	.934 [0]	.951 [0]	.952 [0]
p-value AR2	.213	.265	.298
Long-run effect [p-value]	-.098 [.044]	-.118 [.038]	-.153 [.017]
		Bootstrap-based bias-corrected estimates	
Independence dummy	-0.004*	-0.006**	-0.007***
	(0.002)	(0.002)	(0.002)
Ex ante effect (t - 1)	-0.025**	-0.040***	-0.039***
	(0.010)	(0.010)	(0.011)
Ex ante effect (t - 2)	-0.017***	-0.006	-0.005
	(0.005)	(0.011)	(0.011)
Independence dummy × Soviet dummy			-0.020
			(0.016)
Independence dummy × Yugoslav dummy			0.008
			(0.032)
Independence dummy × Referendum dummy			0.015***
			(0.004)
Observations	11,111 [192]	8,318 [187]	8,318 [187]
Country + Year FE	yes	yes	yes
Control vector	none	$\mathbf{X}^{\mathbf{compact}}$	$\mathbf{X}^{\mathbf{extensive}}$
$\sum_{q=1}^{4} \alpha_q$ [F-test <1]	.988 [.001]	.971 [0]	.972 [0]
Long-run effect [p-value]	-.32 [.117]	-.192 [.019]	-.256 [.004]

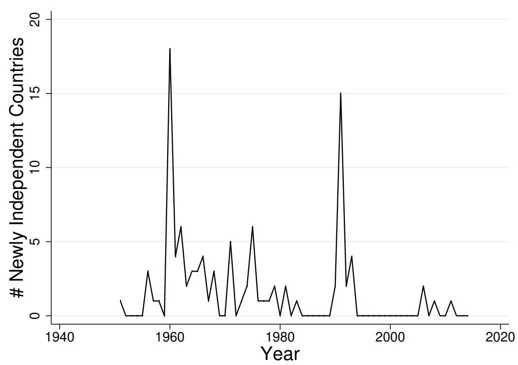
Note: The table presents parametric estimates of the effect of declaring independence on log per capita GDP. The top panel uses the within estimator; the middle panel uses the orthogonal deviations GMM estimator, which treats all lagged dependent variables as endogenous and all other variables as predetermined. The bottom panel uses the bootstrap-based bias-correction estimator implemented with i.i.d resampling, analytical heterogeneous initiation and bootstrapped standard errors based on 50 iterations. F-test reports the p-value of the persistence in GDP being smaller than 1; AR(2) reports the p-value of the Arellano and Bond (1991) test for second-order auto-correlation in first differences. The estimated long-run economic impact of secession, as well as the p-value for this being different from 0, is reported in the bottom row. Standard errors of the long-run effect are computed through the delta-method, see Papke and Wooldridge (2005). The full results are reported in table A1.
*** p<0.01, ** p<0.05, * p<0.1.

5 Robustness check: instrumental variable estimates

One major drawback of the previous estimators lies in the fact that they critically hinge on the strict exogeneity assumption imposed on the independence dummy and its anticipation effects, see conditions 1 and 1a. Although we relaxed this assumption by including several alternative growth predictors as controls, it is still possible for the decision to secede to be related to any number of past or present unobserved growth determinants, z. For example, increasing inter-regional income differentials in the mother country may both depress future growth and stimulate demands for autonomy.[7] To more extensively deal with the concern of unobserved local economic conditions co-determining both incentives to secede as well as growth potential, we conclude this parametric route by developing an instrumental variable (IV) approach.

To do so, recall that secessions historically have tended to occur in waves (Spencer, 1998; Fazal & Griffiths, 2008; Dahl, Gates, Hegre, & Strand, 2013; Qvortrup, 2014). This observation is visualized in figure 1, which plots world wide state entry for each year between 1950 and 2013: the 1960-peak coincides with the African decolonization process; the elevated level of state entry between 1970 and 1980 reflects several South American countries gaining independence; the peak around 1990 captures the dissolution of both the Soviet Union and Yugoslavia. Following similar arguments as those related to the spatio-temporal clustering of global democratization waves (Acemoglu et al., 2014), we argue that, while it is still unclear why state fragmentation has a tendency to occur in waves, these waves appear unrelated to contemporary economic trends. Rather, they seem to reflect political demands for self-determination and democracy. To the extent that this is the case, global waves of secession provide an attractive source of exogenous variation in local incentives to secede.

Figure 1: **Global waves of secession**



Note: The figure plots net state entry in the world in the period 1950-2013.

[7]See Alesina and Rodrik (1994) and Persson and Tabellini (1994) for evidence on income inequality being a drag on economic growth. For a theoretical explanation as to why inter-regional income inequalities may trigger secessionist tendencies, see Bolton and Roland (1997).

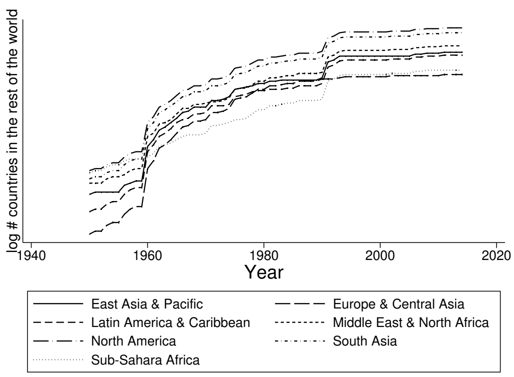
Therefore, to instrument the independence dummy and its anticipation effects, we construct an index of global, non-regional secession waves for each of the seven regions distinguished by the World Bank (2015). More specifically, this index collects the natural log of the total number of countries in the rest of the world, excluding countries in the own region. Formally, denoting the relevant regions by $r \in \{1, \ldots, 7\}$, we define our index of non-regional secession waves, $I_{i,t}$, for country i in region r at time t as

$$I_{i,t} = ln\left(\sum_{j \notin r} 1_{j,t}\right) \tag{4}$$

where $1_{j,t}$ is an indicator function equal to 1 if country j is independent at time t.

Figure 2 plots the values of this index for the period 1950-2013. Irrespective of geographical location, there are clear temporary upward trends in the values of the index reflecting that global waves of secession were tightly clustered in time during this period. For example, the total number of countries in the rest of the world drastically increased for each region around 1960, except for Sub-Sahara Africa where a lot of these NICs were located. Similarly, the break-up of the Soviet Union and Yugoslavia caused an increase in the value of the index for each region except Europe & Central Asia.

Figure 2: **Non-regional waves of secession**



Note: The figure plots non-regional state entry for the seven regions distinguished by the World Bank (2015).

We contend that this index provides us with both strong and valid instruments for the independence dummy and its anticipation terms in the first-stage regressions. First, given the time-clustered nature of state entry at the global level, this index should capture global political opportunities of independence and reflect, among other factors, the global political willingness to accept the right to self-determination. This, in turn, is expected to co-determine local aspirations of independence, implying that the index contains useful information to gauge local incentives to secede. Our first-stage results, reported below, provide further empirical evidence in this regard. On the other hand, we maintain that there is no *a priori* theoretical reason why the (evolution in the) total number of countries in the rest of the world should influence local per capita GDP trajectories, except through their influence on local incentives to secede. To maximally ensure that the latter condition holds, we exclude regional trends in the computation of the index since these could proxy for the presence of local trade disruptions, military conflict and/or macroeconomic instability. In addition, we show below that explicitly including these covariates as controls does not qualitatively affect the results. Finally we only rely on contemporary and future values of the index because it seems unlikely that unknown future evolutions in non-regional state formation would alter current economic behavior, as opposed to known past trends. In order to enable a Hansen-type test for the overall validity of the instrumentsinstruments, we include the contemporary values of this index along with its nearest three leads.. This yields the following fixed effects two-stage least squares model:

$$
\begin{aligned}
\tilde{y_{i,t}} &= \beta_1 N\tilde{I}C_{i,t} + \beta_2 N\tilde{I}C^1_{i,t} + \beta_3 N\tilde{I}C^2_{i,t} + \sum_{q=1}^{4} \alpha_q \tilde{y}_{i,t-q} + \theta \tilde{X}_{i,t} + \eta_t + \tilde{\epsilon}^Z_{i,t} \\
N\tilde{I}C_{i,t} &= \sum_{s=0}^{3} \sigma_s \tilde{I}_{i,t+s} + \sum_{q=1}^{4} \phi_q \tilde{y}_{i,t-q} + \xi_t + \tilde{v}_{i,t} \\
N\tilde{I}C^1_{i,t} &= \sum_{s=0}^{3} \varphi_s \tilde{I}_{i,t+s} + \sum_{q=1}^{4} \rho_q \tilde{y}_{i,t-q} + \zeta_t + \tilde{u}_{i,t} \\
N\tilde{I}C^2_{i,t} &= \sum_{s=0}^{3} \nu_s \tilde{I}_{i,t+s} + \sum_{q=1}^{4} \varpi_q \tilde{y}_{i,t-q} + \chi_t + \tilde{\gamma}_{i,t}
\end{aligned} \tag{5}
$$

where the relevant exclusion restrictions now boil down to

Condition 3

$$
\begin{aligned}
&\mathbb{E}\left[N\tilde{I}C_{i,t}\left(\tilde{I}_{i,t}, \tilde{I}_{i,t+1}, \tilde{I}_{i,t+2}, \tilde{I}_{i,t+3}\right)'\right] && \neq 0 \\
&\mathbb{E}\left[N\tilde{I}C^1_{i,t}\left(\tilde{I}_{i,t}, \tilde{I}_{i,t+1}, \tilde{I}_{i,t+2}, \tilde{I}_{i,t+3}\right)'\right] && \neq 0 \\
&\mathbb{E}\left[N\tilde{I}C^2_{i,t}\left(\tilde{I}_{i,t}, \tilde{I}_{i,t+1}, \tilde{I}_{i,t+2}, \tilde{I}_{i,t+3}\right)'\right] && \neq 0 \\
&\mathbb{E}\left[\tilde{\epsilon}^Z_{i,s}\left(\tilde{I}_{i,t}, \tilde{I}_{i,t+1}, \tilde{I}_{i,t+2}, \tilde{I}_{i,t+3}\right)'\right] && = 0 \ \forall \ t \geq s
\end{aligned}
$$

The baseline results of our IV-approach are reported in column (1) of table 3. In line with subsequent estimations, the strong first-stage results (F-statistic = 227) confirm the expected correlation between contemporary incentives to secede in the own country and contemporary as well as future realizations of independence in the rest of the world. Reassuringly, a Hansen J-test fails to reject the null hypothesis that our over-identification restrictions hold (p-value = 0.78), finding no evidence of global waves of secessions co-determining local per capita GDP trajectories outside of their impact on local incentives to secede. Similar to our previous results, there is a fairly high amount of persistence in per capita GDP (in this case, $\sum_{q=1}^{4} \alpha_q = 0.973$) but an F-test rejects dynamic instability.

With regard to the coefficients of interest, first note that our results now suggest the absence of any relevant anticipation effects, as the coefficient associated with the lead of the independence dummy is now statistically insignificant. The absence of meaningful pre-independence trends in per capita GDP trajectories of NICs nearing their declaration of independence is reassuring, in the sense that this points towards the instrument exogeneity requirement being met: *prior* to their actual declaration of independence, global waves of secession seem not systematically negatively related to economic conditions in future NICs. Only to the extent that they coincide with the period following a NIC's *actual* declaration of independence do they statistically significantly and negatively affect per capita GDP trajectories in NICs. Remarkably, the estimated adverse long-run impact of state fragmentation is now roughly twice as high in comparison to the simple pooled OLS estimates reported earlier. More specifically, our benchmark IV-estimates imply a short-run per capita secession cost of around 1.4% and an implied long-run cost of 54%, which are in the range of our semi-parametric estimates.

Column (2), then, controls for the possibility that global, non-regional waves of secession do violate the exclusion restriction of instrument validity by proxying for local military conflict, macroeconomic instability and/or trade disruptions. Reassuringly, as can be seen, our baseline results are fairly stable with respect to the inclusion of these control variables and the estimated long-run per capita cost of independence now increases to 75%. The third column, finally, confirms once again our prior finding that independence-by-referendum seems preferable from an economic point of view.

One possible explanation for these rather gloomy estimates of the economic impact of secession is that they stem from the bluntness of the instruments at our disposal, which approximate *country-level* incentives to secede with *regional-level* information on non-regional secession waves. Doing so, our first-stage results might lead to an over-accentuation of the economic trends of highly secessionist countries located in historically secession-rich regions, such as Sub Sahara Africa. Compared to highly secessionist countries located in secession-poorer regions, these countries may face more severe independence costs precisely because of the accumulation of adverse effects of state fragmentation in their associated region.[8] Keeping this in mind, exploiting variation in contemporary and future global, non-regional waves of secessions qualitatively confirms our prior conclusions suggesting that past instances of state fragmentation persistently adversely impacted per capita GDP trajectories in NICs, and potentially severely so.

[8] Venables (2010), for instance, claims that excessive state fragmentation in Sub Sahara Africa hampers growth by, among other factors, diminishing agglomeration economies.

Table 3: **Economic impact of secession (IV estimates)**

	IV regression results		
	(1)	(2)	(3)
Independence dummy	-0.014***	-0.035**	-0.041**
	(0.004)	(0.015)	(0.018)
Ex ante effect (t - 1)	-0.033	-0.057	-0.075
	(0.023)	(0.125)	(0.149)
Ex ante effect (t - 2)	0.009	-0.018	-0.038
	(0.017)	(0.077)	(0.092)
Independence dummy × Soviet dummy			-0.019
			(0.102)
Independence dummy × Yugoslav dummy			0.003
			(0.087)
Independence dummy × referendum dummy			0.038***
			(0.014)
Observations	10,768 [192]	8,184 [187]	8,184 [187]
Adjusted R-squared	0.979	0.976	0.958
Country + Year FE	yes	yes	yes
Control vector	none	$\mathbf{X}^{\mathbf{compact}}$	$\mathbf{X}^{\mathbf{extensive}}$
$\sum_{q=1}^{4} \alpha_q$ [F-test <1]	.973 [.000]	.954 [.000]	.954 [.000]
Hansen-J test [p-value]	.078 [.78]	1.444 [.23]	1.063 [.303]
F-test [first stage]	227.58	23.13	13.92
Long-run effect [p-value]	-.537 [0]	-.753 [.023]	-.895 [.026]

Note: This table presents fixed effects two-stage least squares estimates of the effect of gaining independence on per capita GDP, where both the decision to secede and its anticipation are instrumented by the four nearest (future) non-regional secession waves. Standard errors, robust against heteroskedasticity and serial correlation at the country level, in parantheses. F-test reports the p-value of the persistence in GDP being smaller than 1; F-test [first stage] reports first stage F-statistic of the excluded instruments; Hansen-J reports the p-value of a Hansen over-identification test. The estimated long-run economic impact of secession, as well as the p-value for this being different from 0, are reported in the bottom row. Standard errors of the long-run effect are computed through the delta-method, see Papke and Wooldridge (2005).
*** p<0.01, ** p<0.05, * p<0.1.

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Table A1: **Baseline parametric estimates (full results)**

	Within estimates		
	(1)	(2)	(3)
Independence dummy	-0.004**	-0.005**	-0.007**
	(0.002)	(0.003)	(0.003)
Independence dummy × Soviet dummy			-0.024**
			(0.011)
Independence dummy × Yugoslav dummy			0.009
			(0.019)
Independence dummy × referendum dummy			0.014***
			(0.003)
Ex ante effect (t - 1)	-0.024***	-0.039***	-0.038***
	(0.008)	(0.013)	(0.014)
Ex ante effect (t - 2)	-0.016***	-0.007	-0.006
	(0.005)	(0.010)	(0.011)
log Per capita GDP (t-1)	1.091***	1.136***	1.135***
	(0.034)	(0.032)	(0.032)
log Per capita GDP (t-2)	-0.030	-0.095**	-0.095**
	(0.048)	(0.039)	(0.039)
log Per capita GDP (t-3)	-0.010	-0.012	-0.013
	(0.046)	(0.025)	(0.025)
log Per capita GDP (t-4)	-0.077***	-0.072***	-0.071***
	(0.023)	(0.018)	(0.018)
log Trade openness		0.005	0.005
		(0.004)	(0.004)
log Surface area		-0.070	-0.064
		(0.062)	(0.062)
log Educational attainment		-0.007	-0.007*
		(0.004)	(0.004)
Life Expectancy		0.001***	0.001***
		(0.000)	(0.000)
log Battle deaths		-3.440	-3.475
		(2.660)	(2.681)
log Population density		-0.029***	-0.030***
		(0.009)	(0.009)
Democracy		0.000	0.000
		(0.000)	(0.000)
Macroeconomic uncertainty		-0.008***	-0.008***
		(0.002)	(0.002)
Political instability		-0.038***	-0.038***
		(0.014)	(0.014)
Oil exporting countries		0.001	0.001
		(0.005)	(0.004)
WTO-membership		0.004	0.005
		(0.003)	(0.003)
EU-membership		-0.005	-0.005
		(0.004)	(0.004)
AU-membership		-0.002	-0.001
		(0.005)	(0.005)
ASEAN-membership		0.029***	0.029***
		(0.008)	(0.008)
OECD-membership		-0.008	-0.008
		(0.005)	(0.005)
MERCOSUR-membership		-0.003	-0.003
		(0.004)	(0.004)
NATO-membership		-0.003	-0.003
		(0.004)	(0.004)
Observations [# countries]	11,128 [192]	8,318 [187]	8,318 [187]
Adjusted R-squared	0.980	0.978	0.978
Country + Year FE	yes	yes	yes
Control vector	none	$\mathbf{X}^{\mathbf{compact}}$	$\mathbf{X}^{\mathbf{extensive}}$
$\sum_{q=1}^{4} \alpha_q$ [F-test <1]	.974 [0]	.956 [0]	.956 [0]
Long-run effect [p-value]	-.161 [.037]	-.124 [.038]	-.16 [.013]

Note: See relevant notes table 2.
*** p<0.01, ** p<0.05, * p<0.1.

Baseline parametric estimates (full results) *continued*

	Deviations GMM estimates		
	(1)	(2)	(3)
Independence dummy	-0.006**	-0.006**	-0.003
	(0.003)	(0.003)	(0.002)
Independence dummy × Soviet dummy			-0.009
			(0.009)
Independence dummy × Yugoslav dummy			0.011
			(0.021)
Independence dummy × referendum dummy			0.009
			(0.006)
Ex ante effect (t - 1)	-0.020**	-0.039***	-0.034**
	(0.009)	(0.013)	(0.014)
Ex ante effect (t - 2)	-0.011**	-0.006	-0.002
	(0.006)	(0.010)	(0.010)
log Per capita GDP (t-1)	1.095***	1.163***	1.259***
	(0.045)	(0.067)	(0.069)
log Per capita GDP (t-2)	-0.116**	-0.151*	-0.193**
	(0.055)	(0.084)	(0.085)
log Per capita GDP (t-3)	0.005	-0.009	-0.012
	(0.048)	(0.026)	(0.027)
log Per capita GDP (t-4)	-0.050**	-0.053***	-0.064***
	(0.021)	(0.020)	(0.019)
log Trade openness		0.006	0.006
		(0.004)	(0.004)
log Surface area		-0.080	-0.070
		(0.066)	(0.066)
log Educational attainment		-0.007	-0.008*
		(0.005)	(0.005)
Life Expectancy		0.001***	0.001***
		(0.000)	(0.000)
log Battle deaths		-3.326	-3.359
		(2.625)	(2.648)
log Population density		-0.033***	-0.033***
		(0.012)	(0.012)
Democracy		0.000	0.000
		(0.000)	(0.000)
Macroeconomic uncertainty		-0.009***	-0.009***
		(0.003)	(0.003)
Political instability		-0.037***	-0.037***
		(0.014)	(0.014)
Oil exporting countries		0.001	0.001
		(0.005)	(0.005)
WTO-membership		0.005*	0.005
		(0.003)	(0.003)
EU-membership		-0.004	-0.004
		(0.004)	(0.004)
AU-membership		-0.003	-0.002
		(0.005)	(0.005)
ASEAN-membership		0.031***	0.031***
		(0.010)	(0.010)
OECD-membership		-0.009	-0.009
		(0.006)	(0.006)
MERCOSUR-membership		-0.004	-0.004
		(0.004)	(0.004)
NATO-membership		-0.003	-0.003
		(0.004)	(0.004)
Observations	10,935 [192]	8,131 [187]	8,131 [187]
Country FE	yes	yes	yes
Control vector	none	$\mathbf{X}^{\mathbf{compact}}$	$\mathbf{X}^{\mathbf{extensive}}$
$\sum_{q=1}^{4} \alpha_q$ [F-test <1]	.934 [0]	.951 [0]	.952 [0]
p-value AR2	.213	.265	.298
Long-run effect [p-value]	-.098 [.044]	-.118 [.038]	-.153 [.017]

Note: See relevant notes table 2.
*** p<0.01, ** p<0.05, * p<0.1.

Baseline parametric estimates (full results) *continued*

	Bootstrap-based bias corrected estimates (1)	(2)	(3)
Independence dummy	-0.004*	-0.006**	-0.007***
	(0.002)	(0.002)	(0.002)
Independence dummy × Soviet dummy			-0.020
			(0.016)
Independence dummy × Yugoslav dummy			0.008
			(0.032)
Independence dummy × Referendum dummy			0.015***
			(0.004)
Ex ante effect (t - 1)	-0.025**	-0.040***	-0.039***
	(0.010)	(0.010)	(0.011)
Ex ante effect (t - 2)	-0.017***	-0.006	-0.005
	(0.005)	(0.011)	(0.011)
log Per capita GDP (t-1)	1.106***	1.149***	1.149***
	(0.031)	(0.035)	(0.035)
log Per capita GDP (t-2)	-0.031	-0.096**	-0.096**
	(0.044)	(0.042)	(0.041)
log Per capita GDP (t-3)	-0.009	-0.012	-0.012
	(0.049)	(0.028)	(0.028)
log Per capita GDP (t-4)	-0.077***	-0.070***	-0.069***
	(0.025)	(0.017)	(0.016)
log Trade Openness		0.005	0.005
		(0.004)	(0.004)
log Surface area		-0.040	-0.033
		(0.064)	(0.065)
log Educational attainment		-0.007	-0.007
		(0.005)	(0.005)
Life Expectancy		0.001***	0.001***
		(0.000)	(0.000)
log Battle deaths		-3.297	-3.339
		(4.782)	(4.764)
log Population density		-0.025***	-0.026***
		(0.007)	(0.007)
Democracy		0.000	0.000
		(0.000)	(0.000)
Macroeconomic uncertainty		-0.008***	-0.008***
		(0.002)	(0.002)
Political instability		-0.040**	-0.040***
		(0.017)	(0.014)
Oil exporting states		0.001	0.002
		(0.005)	(0.005)
WTO-membership		0.004	0.004
		(0.003)	(0.003)
EU-membership		-0.005	-0.005
		(0.004)	(0.004)
AU-membership		-0.003	-0.002
		(0.005)	(0.004)
ASEAN-membership		0.029***	0.029***
		(0.010)	(0.010)
OECD-membership		-0.007	-0.006
		(0.006)	(0.006)
MERCOSUR-membership		-0.001	-0.001
		(0.004)	(0.004)
NATO-membership		-0.005	-0.005
		(0.004)	(0.004)
Observations	11,111 [192]	8,318 [187]	8,318 [187]
Country + Year FE	yes	yes	yes
Control vector	none	$\mathbf{X}^{\mathbf{compact}}$	$\mathbf{X}^{\mathbf{extensive}}$
$\sum_{q=1}^{4} \alpha_q$ [F-test <1]	.988 [.001]	.971 [0]	.972 [0]
Long-run effect [p-value]	-.32 [.117]	-.192 [.019]	-.256 [.004]

Note: See relevant notes table 2.
*** p<0.01, ** p<0.05, * p<0.1.