

Shift-Share Identification and Randomization Inference in Sequeira, Nunn and Qian (2020): A Supplementary Note

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Abstract

Sequeira, Nunn and Qian (2020) (hereafter SNQ) instrument historical immigration to US counties with the interaction between railroad access and national immigration inflows over seven decades. This note complements my comment on SNQ (Adema, 2026). This note shows that the instrument is a shift-share instrument whose shocks are the seven decadal inflows, and asks whether the modern shift-share framework of Borusyak, Hull and Jaravel (2022), in which shocks are treated as random and the county-level exposure to them may be endogenous, rescues the design. Controlling for the sum of shares, as that framework requires, lowers the first-stage F-statistic from 12.1 to 8.5 and reduces the 2SLS estimate of the effect on income per capita in 2000 by half, to a value no longer distinguishable from zero. Randomization inference that redraws the seven decadal shocks yields p-values between 0.28 and 0.93 across all specifications, including SNQ’s own. For the specification that controls for the sum of shares, the 90% randomization-inference confidence interval spans effects from -12% to $+7.5\%$ per percentage point of immigrant share. With only seven shocks, the design does not contain enough exogenous variation to estimate the long-run effects of migration with useful precision.

1 Purpose and setting

This note collects the shift-share and randomization-inference analyses that accompanied earlier versions of my comment on SNQ (Adema, 2026, hereafter: the comment). The comment itself focuses on one problem: SNQ’s control for the timing of railroad connection mainly separates ever-connected from never-connected counties, and once the timing of connection is controlled for effectively, the first stage weakens and the 2SLS estimate turns negative. The analyses collected here ask a different question. Even if SNQ’s parametric controls are set aside, can the design be justified within the recent literature on shift-share instruments, and what does inference that respects the structure of the instrument imply? The answer matters beyond SNQ, because the same construction, national migration inflows interacted with a local measure of access, continues to be used in recent work.¹

SNQ estimate the effect of the average immigrant share of a county between 1860 and 1920 on outcomes in 2000, instrumenting it with

$$Z_i = \frac{1}{D_i} \sum_{t=E_i+10}^{1920} \text{Immigrant Flow}_{t-10} \times I_{i,t-10}^{\text{RR Access}}, \quad (1)$$

where $\text{Immigrant Flow}_{t-10}$ is European immigration to the US between census years $t-10$ and t relative to the US population in $t-10$, $I_{i,t-10}^{\text{RR Access}}$ indicates whether county i had railroad access in $t-10$, E_i is the first census decade in which the county appears (its decade of establishment), and D_i is the number of decades the county existed between 1860 and 1920. The first and second stages include state fixed effects, a polynomial in latitude and longitude, two control functions that interact railroad access with national industrialization

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¹For example, Escamilla-Guerrero, Papadia and Zimran (2026) study the effects of immigration on agricultural development in Brazil.

and GDP growth, and the log of one plus the number of years since railroad connection, $\log(\text{years of rail}+1)$. Standard errors allow for spatial correlation within five degrees (Conley, 1999). Because railroad access is an absorbing state, the value of Z_i is fully determined by a county’s decade of establishment and its decade of railroad connection.

Columns 1–6 of Table 1 reproduce the specifications of Table 1 of the comment for log income per capita in 2000, SNQ’s main outcome. Column 1 is SNQ’s baseline. Because 127 of the 2,935 counties never received a railroad connection and are assigned a value of zero, $\log(\text{years of rail}+1)$ mainly distinguishes these counties from all others, rather than early- from late-connected counties. Column 2 replaces it with an indicator for never-connected counties, and Column 3 includes both, which together control for the timing of connection with a specific functional form. Column 4 instead combines the indicator with a linear control for the years since connection, Column 5 includes fixed effects for the decade of railroad connection, and Column 6 adds fixed effects for the decade of county establishment. Once the timing of connection is controlled for in Columns 3–5, the 2SLS estimate turns negative, and in Column 6 the first stage is irrelevant. Column 7 and the randomization-inference p-values are the subject of this note.

2 SNQ’s instrument as a shift-share instrument

Equation 1 can be written as a sum of shocks weighted by county-specific shares:

$$Z_i = \sum_{t=1860}^{1920} \text{Immigrant Flow}_{t-10} \times s_{i,t}, \quad s_{i,t} = \begin{cases} I_{i,t-10}^{\text{RR Access}} / D_i, & \text{if } t - 10 \geq E_i, \\ 0, & \text{otherwise.} \end{cases} \quad (2)$$

The shifts are the seven decadal immigration inflows, and the shares record in which of those decades a county existed and had railroad access. Unlike most shift-share applications, where shocks are indexed by industries or countries of origin, the shock dimension here is time.

The literature offers two routes to identification with such instruments: exogenous shares (Goldsmith-Pinkham, Sorkin and Swift, 2020) or exogenous shifts (Borusyak, Hull and Jaravel, 2022; Adao, Kolesár and Morales, 2019). Exogenous shares are implausible in this setting, as SNQ themselves note (p. 392): counties connected to the railroad in particular periods “may have disproportionately benefitted from being connected to the railway”. Exogenous shifts are more plausible, because the size of the national inflow in a given decade is unlikely to respond to conditions in any one county. The exogenous-shifts framework requires two conditions. First, the shocks should have the same expected value across decades, which is reasonable because the realized inflows fluctuate from decade to decade without a trend; the assumption could be relaxed with additional controls. Second, there must be many shocks that are not strongly correlated. As Borusyak, Hull and Jaravel (2025) write, the approach “requires many shifts $g_1, \dots, g_K$. Otherwise, if K is a small number, the shifts may by chance be correlated with unobservables even if they are truly random”. With seven decadal shocks, this second condition is the binding one.

Under exogenous shifts, the shares may be endogenous as long as the regression controls for each county’s expected exposure to the shocks. Counties whose shares sum to a larger value receive a larger instrument in expectation, whatever the realization of the shocks; if the shares are endogenous, this expected component violates the exclusion restriction. The expected instrument is proportional to the sum of shares,

$$\mu_i = \sum_{t=1860}^{1920} s_{i,t} = \frac{R_i}{D_i}, \quad (3)$$

where R_i is the number of decades between 1860 and 1920 in which the county had railroad access. Like the instrument itself, μ_i is higher for counties connected earlier and lower for older counties (Figure 1). The attraction of the framework is that it tells the researcher which control removes the endogenous part of the instrument and in what functional form, so it is not necessary to choose parametric controls for railroad timing and county age. Table 3 shows that μ_i is strongly correlated with the instrument, the immigrant share, and income per capita in 2000, also conditional on SNQ’s baseline controls.

3 Controlling for the sum of shares

Column 7 of Table 1 replaces SNQ’s $\log(\text{years of rail}+1)$ control with μ_i . The Conley-based first-stage F-statistic falls from 12.1 to 8.5, and the 2SLS coefficient falls from 2.62 to 1.33 and is no longer statistically significant; the 90% Anderson-Rubin confidence interval, which remains valid when the instrument is weak (Andrews, Stock and Sun, 2019), ranges from -0.33 to 4.44 . Controlling for expected exposure thus removes about half of SNQ’s estimate, in the same direction as the parametric controls for railroad timing in Columns 3–5, but less strongly.

With only seven shocks, however, conditioning on μ_i need not remove all systematic variation, and it does not. Figures 2 and 3 show that after controlling for μ_i , both the residualized instrument and residualized income in 2000 still differ systematically across decades of railroad connection and decades of establishment, and that the residualized instrument remains on average higher for earlier-connected and older counties. Because earlier-connected and older counties are plausibly richer today for reasons unrelated to migration, this residual variation is likely to bias the estimate upward. Table 2 therefore adds linear controls for the years since railroad connection and the years since county establishment, with and without μ_i . A linear control for railroad timing alone yields a smaller estimate that is not significant at conventional levels, although its Anderson-Rubin confidence interval excludes zero (Column 1); adding μ_i leaves an estimate of 0.57 that is not distinguishable from zero (Column 2). With a linear control for county age the estimate is insignificant with or without μ_i (Columns 3 and 4), and fixed effects for the decade of establishment leave the first stage irrelevant (Column 5). Table 4 shows that the pattern is similar without the industrialization and GDP-growth control functions: controlling for μ_i keeps the first stage strong ($F = 10.9$) but reduces the 2SLS estimate from 4.20 to 1.74, which is significant only at the 10% level. Figure 4 shows the corresponding residual variation in the immigrant share itself.

In sum, the shift-share framework points to the same conclusion as the parametric analysis in the comment. Once the expected exposure to the shocks is removed, the estimate shrinks and loses significance, and the residual variation that remains is still tied to railroad timing and county age.

4 Randomization inference

The small number of shocks also affects inference. Each decadal shock enters the instrument of every county that had railroad access in that decade, so treatment is assigned in clusters defined by the decade of establishment and the decade of railroad connection. Conventional standard errors, including spatial ones, treat distant counties in the same cluster as independent and therefore overstate precision (Moulton, 1990; Adao, Kolesár and Morales, 2019; Abadie et al., 2023). Borusyak, Hull and Jaravel (2022) show that the 2SLS estimator can be re-expressed as a shock-level regression, but valid standard errors from that representation rely on many shocks and on fewer shock-level controls than shocks, which fails with seven shocks. For such cases, Borusyak and Hull (2023) recommend randomization inference: redraw the shocks from a plausible counterfactual distribution, recompute the instrument, and compare the realized test statistic with the distribution of simulated ones. Given the assumed distribution of the shocks, the test is exact in finite samples and remains valid when the instrument is weak. Its validity does depend on that assumed distribution; permuting the shocks across decades, for example, treats them as exchangeable, which would fail if the inflows were strongly serially correlated. I therefore also report results for two alternative shock distributions.

I draw the seven shocks in three ways: permuting the realized shocks across decades (7! possible draws), drawing from the realized shocks with replacement, and drawing from a normal distribution with the mean (0.076) and standard deviation (0.025) of the realized shocks. For each of 999 draws I recompute the instrument and four statistics: the first-stage t-statistic, the reduced-form coefficient and its t-statistic, and the statistic that Borusyak and Hull (2023) recommend for the 2SLS estimate,

$$T(b) = \sum_i (Z_i - \theta\mu_i) (\tilde{y}_i - b\tilde{x}_i), \quad (4)$$

where $\tilde{y}_i$ and $\tilde{x}_i$ are income and the immigrant share residualized on the controls, θ is the mean decadal shock, so that $\theta\mu_i$ is the expected instrument, and b is the hypothesized effect. For $b = 0$, T is proportional

to the covariance between the outcome and the recentered instrument. The two-sided p-value is twice the smaller of the shares of simulated statistics above and below the realized one, and inverting the test over a grid of 1,001 values of b between -15 and 15 yields a confidence interval.

Figure 5 shows the distributions for SNQ’s baseline specification. The realized value of T is unremarkable relative to its randomization distribution: the randomization p-value is 0.791, whereas the conventional p-value based on Conley standard errors is about 0.01. The reduced-form t-statistics under redrawn shocks are skewed toward positive values, and a large share of them exceed 2 (panel d), so conventional inference would frequently detect a positive effect of the instrument even when the shocks are assigned at random. This is what one would expect if the instrument picks up the higher incomes of early-connected counties rather than the realization of the shocks. The first-stage t-statistic is extreme relative to its distribution ($p = 0.002$), so the instrument does predict immigration.

Controlling for μ_i recenters the distributions of T and of the reduced-form coefficient around zero (Figure 6). The first stage remains relevant ($p = 0.028$), and the randomization p-value for T is 0.763. The simulated reduced-form t-statistics, however, still fall below -2 or above 2 in a large share of draws (panel d), which shows how strongly conventional standard errors overstate precision in this design. Drawing shocks with replacement or from a normal distribution gives p-values for T of 0.609 and 0.613 (Figures 7 and 8). Across all specifications in Tables 1 and 2, the randomization p-values for the 2SLS estimate range from 0.28 to 0.93, consistently larger than the conventional p-values. In the 2,808 counties that were ever connected to the railroad, SNQ’s baseline specification yields a 2SLS estimate of -2.32 (Conley standard error 1.64), because $\log(\text{years of rail}+1)$ then measures only the timing of connection; its randomization p-value is 0.404.

Inverting the test for the specification that controls for μ_i gives a 90% confidence interval from -12.6 to 7.2 (Figure 9). A one percentage point higher average immigrant share between 1860 and 1920 is thus consistent with anything from a 12% reduction to a 7.5% increase in income per capita in 2000. This interval is too wide to be informative about the question SNQ pose.

5 Implications for the design

Viewed through the shift-share framework, SNQ’s design has a clear logic: national immigration inflows are plausibly exogenous to any single county, and controlling for expected exposure is meant to remove the endogeneity of railroad access. The framework also makes clear why the design cannot deliver precise answers in this application. Identification rests on only seven decadal shocks, which enter the instrument in a pattern tied to when a county was founded and connected. Controlling for expected exposure reduces the estimate by half and removes its statistical significance, the remaining variation is still correlated with railroad timing and county age, and inference that respects the small number of shocks cannot distinguish sizable negative from sizable positive effects. These features are not specific to SNQ’s data. Designs that rely on a short time series of national shocks for identification are likely to face the same limitation; within the exogenous-shifts framework, they become informative only when the number of independent shocks is large, for example when shocks vary across many countries of origin rather than only over time.

Tables

Table 1: Re-analysis of SNQ Table 4, Column 1, including the sum-of-shares control

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Panel A: First stage			Average immigrant share, 1860–1920				
Instrument	4.559*** (1.311)	4.512*** (1.290)	3.929** (1.810)	3.953** (1.753)	6.463** (2.544)	3.053 (2.535)	5.587*** (1.914)
Panel B: Reduced Form			Log income per capita, 2000				
Instrument	11.942*** (3.629)	12.254*** (3.626)	-7.767* (4.474)	-8.042* (4.584)	-18.248*** (5.813)	-17.045*** (6.226)	7.414 (5.696)
Panel C: 2SLS			Log income per capita, 2000				
Average Immigrant	2.619** (1.022)	2.716*** (1.046)	-1.977 (1.452)	-2.034 (1.435)	-2.823** (1.348)		1.327 (1.147)
Share 1860-1920							
RI p -value	{0.791}	{0.767}	{0.436}	{0.420}	{0.284}		{0.763}
90% Anderson-Rubin CI	[1.26, 5.28]	[1.35, 5.43]	[-9.09, -0.12]	[-8.28, -0.15]	[-8.1, -1.23]	$[-\infty, \infty]$	[-0.33, 4.44]
log(Years of rail + 1)	0.005 (0.006)		0.348*** (0.081)				
Never connected		-0.013 (0.030)	1.571*** (0.369)	0.266*** (0.073)			
Years of rail				0.003*** (0.001)			
Observations	2935	2935	2935	2935	2935	2935	2935
First-stage F-statistic	12.09	12.23	4.71	5.09	6.46	1.45	8.52
R^2 of IV on controls	0.931	0.930	0.963	0.963	0.980	0.986	0.970
Decade of railroad connection FE					✓	✓	
Decade of county establishment FE						✓	
Sum of shares (μ_i)							✓

Notes: First stage, reduced form and 2SLS estimates based on Column 1 of Table 4 of SNQ. Column 1 presents SNQ's baseline specification; Columns 2–6 are the specifications of Table 1 of the comment; Column 7 replaces the $\log(\text{years of rail}+1)$ control with the sum of shares μ_i . Additional controls are indicated by included coefficients in Panel C or checkmarks in the last rows. Because the first stage in Column 6 is very weak and the Anderson-Rubin confidence interval is unbounded, the 2SLS point estimate and randomization-inference p -value are omitted. Coefficients on additional controls shown in Panel C are omitted for brevity in Panels A and B. Conley standard errors using a five-degree window are shown in parentheses. Anderson-Rubin confidence intervals based on Conley standard errors (square brackets) and randomization-inference p -values based on the statistic T of Section 4, permuting the seven decadal shocks 999 times (curly brackets), are reported for the 2SLS coefficient. The Anderson-Rubin confidence interval is constructed by inverting the Conley-based Anderson-Rubin test over a grid of 1,001 candidate coefficients from -15 to $+15$; the reported bounds are the smallest and largest candidate values not rejected at the 10% level, and are set to $\pm\infty$ when ± 15 is not rejected. Kleibergen-Paap F-statistics based on Conley standard errors and the R^2 of a regression of the instrument on all control variables are reported in the last rows. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table 2: Controlling for railroad timing, county age and the sum of shares

	(1)	(2)	(3)	(4)	(5)
Panel A: First stage					
	Average immigrant share, 1860–1920				
Instrument	4.610*** (1.465)	5.653*** (1.923)	3.707** (1.460)	4.798** (1.977)	1.391 (1.358)
Panel B: Reduced Form					
	Log income per capita, 2000				
Instrument	6.560* (3.790)	3.243 (5.585)	3.669 (3.219)	-1.627 (5.303)	7.035 (4.747)
Panel C: 2SLS					
	Log income per capita, 2000				
Average Immigrant Share 1860-1920	1.423 (0.875)	0.574 (1.016)	0.990 (0.950)	-0.339 (1.096)	
RI p -value	{0.821}	{0.847}	{0.931}	{0.933}	
Anderson-Rubin 90% CI	[0.09, 3.54]	[-1.14, 2.91]	[-0.45, 3.84]	[-2.76, 2.1]	$[-\infty, \infty]$
Observations	2935	2935	2935	2935	2935
First-stage F-statistic	9.91	8.64	6.45	5.89	1.05
R^2 of IV on controls	0.945	0.972	0.942	0.973	0.958
Years of railroad connection	✓	✓			
Years since county establishment			✓	✓	
Sum of shares (μ_i)		✓		✓	
Decade of county establishment					✓
FE					

Notes: First stage, reduced form and 2SLS estimates with alternative controls for the timing of railroad connection and county age. All models include the baseline controls of Table 4 of SNQ, except for the $\log(\text{years of rail}+1)$ control. Column 1 controls linearly for the years of railroad connection in 2000. Column 2 adds the sum of shares μ_i . Column 3 controls linearly for the years since county establishment in 2000, and Column 4 adds μ_i . Column 5 includes fixed effects for the decade of county establishment; because the first stage is very weak and the Anderson-Rubin confidence interval is unbounded, the 2SLS point estimate and randomization-inference p -value are omitted. Conley standard errors using a five-degree window are shown in parentheses. Anderson-Rubin confidence intervals (square brackets) and randomization-inference p -values (curly brackets) are constructed as in Table 1. Kleibergen-Paap F-statistics based on Conley standard errors and the R^2 of a regression of the instrument on all control variables are reported in the last rows. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table 3: Correlation between the sum of shares and the instrument, migration, and income per capita

	Instrument		Average Immigrant Share, 1860-1920		Log income per capita in 2000	
	(1)	(2)	(3)	(4)	(5)	(6)
Expected instrument ($\theta\mu_i$)	0.954*** (0.014)	1.332*** (0.071)	4.481*** (1.167)	4.925** (2.136)	11.001*** (2.365)	20.857*** (7.898)
Observations	2935	2935	2935	2935	2935	2935
Baseline controls		✓		✓		✓

Notes: OLS regressions of the instrument, the average immigrant share between 1860 and 1920 and log income per capita in 2000 on the expected instrument, $\theta\mu_i$ (the sum of shares times the mean decadal shock, in the units of the instrument). Odd columns include no controls; even columns include all baseline controls of Table 4 of SNQ, including the $\log(\text{years of rail}+1)$ control. Conley standard errors using a five-degree window are shown in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

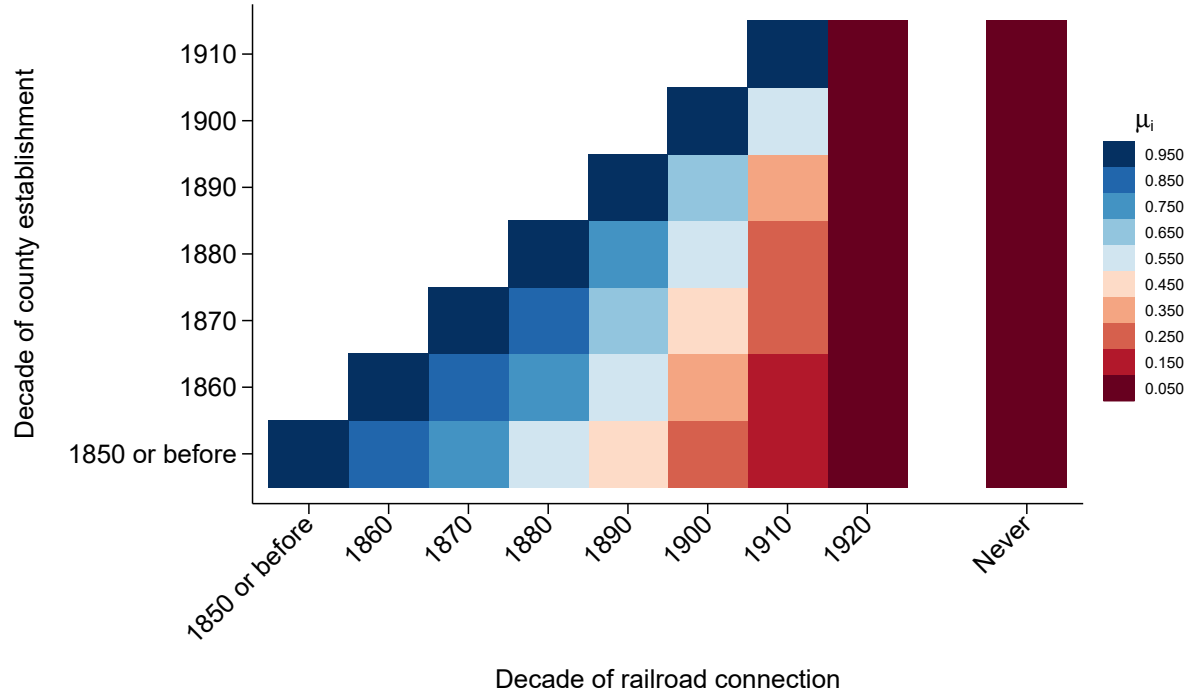
Table 4: Specifications of Table 1 without the industrialization and GDP growth control functions

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Panel A: First stage		Average immigrant share, 1860–1920					
Instrument	1.590*** (0.479)	1.584*** (0.458)	0.635 (0.844)	0.676 (0.775)	0.770 (0.999)	1.655 (1.836)	6.048*** (1.835)
Panel B: Reduced Form		Log income per capita, 2000					
Instrument	6.679*** (1.695)	7.146*** (1.707)	-2.799 (2.130)	-2.029 (2.090)	-2.498 (2.400)	2.395 (3.943)	10.531** (4.997)
Panel C: 2SLS		Log income per capita, 2000					
Average Immigrant Share 1860-1920	4.201** (1.734)	4.512** (1.813)					1.741* (1.026)
log(Years of rail + 1)	0.008 (0.006)		0.426*** (0.156)				
Never connected		-0.031 (0.030)	1.942*** (0.722)	0.296** (0.135)			
Years of rail				0.003** (0.001)			
Observations	2935	2935	2935	2935	2935	2935	2935
First-stage F-statistic	11.02	11.95	0.57	0.76	0.59	0.81	10.86
R^2 of IV on controls	0.507	0.431	0.796	0.770	0.852	0.931	0.967
Decade of railroad connection FE					✓	✓	
Decade of county establishment FE						✓	
Sum of shares (μ_i)							✓

Notes: First stage, reduced form and 2SLS estimates based on Column 1 of Table 4 of SNQ, excluding the two control functions for industrialization and GDP growth; all other baseline controls of SNQ are included. The columns follow the same specifications as Table 1. Because the first stages are very weak in Columns 3–6, the corresponding 2SLS point estimates are omitted. Coefficients on additional controls shown in Panel C are omitted for brevity in Panels A and B. Conley standard errors using a five-degree window are shown in parentheses. Kleibergen-Paap F-statistics based on Conley standard errors and the R^2 of a regression of the instrument on all control variables are reported in the last rows. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

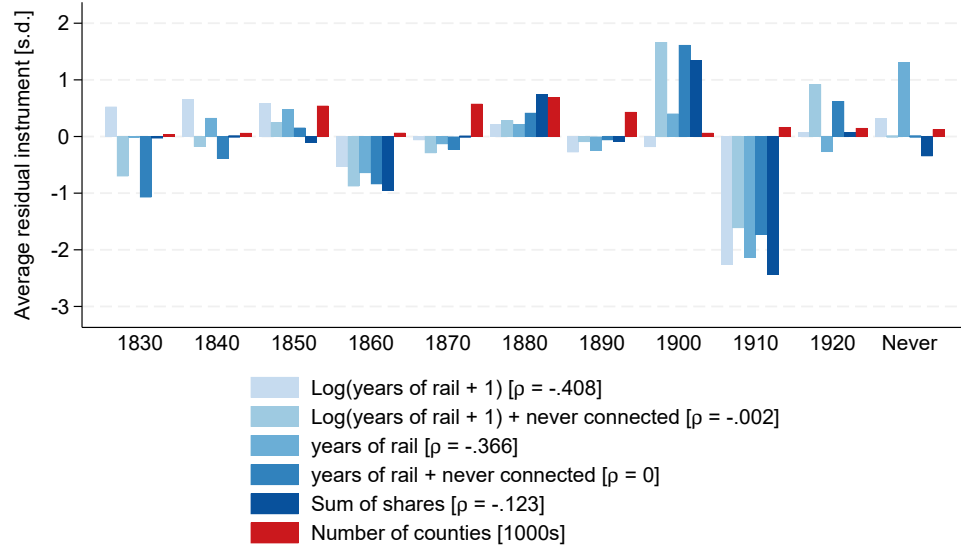
Figures

Figure 1: Sum of shares μ_i by decade of establishment and decade of railroad connection

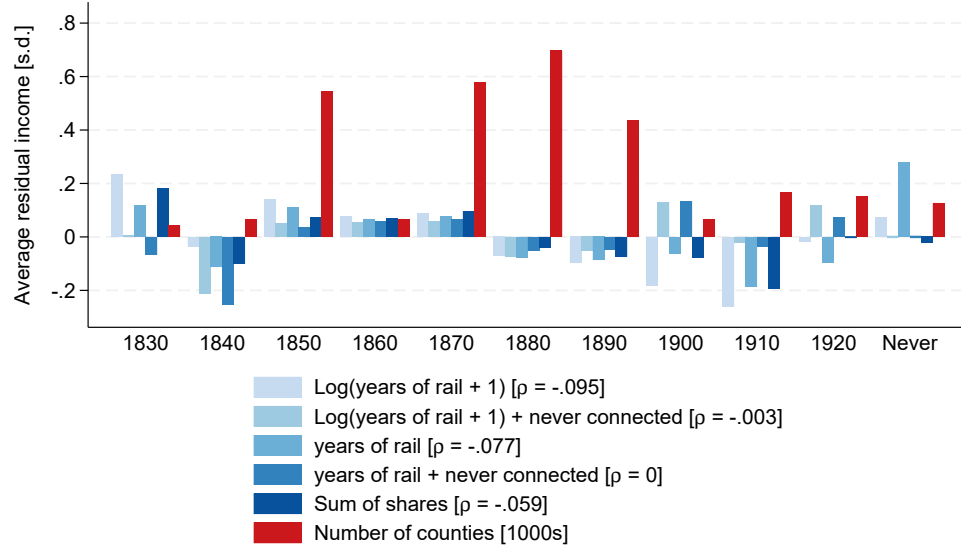


Notes: Sum of shares $\mu_i = R_i/D_i$ (Equation 3) by decade of county establishment and decade of railroad connection. N = 2,935.

Figure 2: Residualized instrument and log income per capita by decade of railroad connection



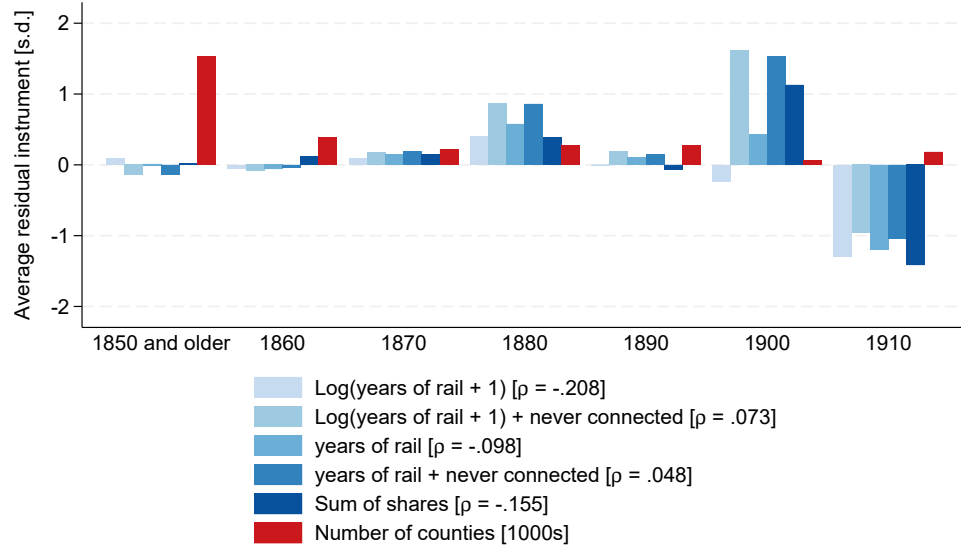
(a) Instrument



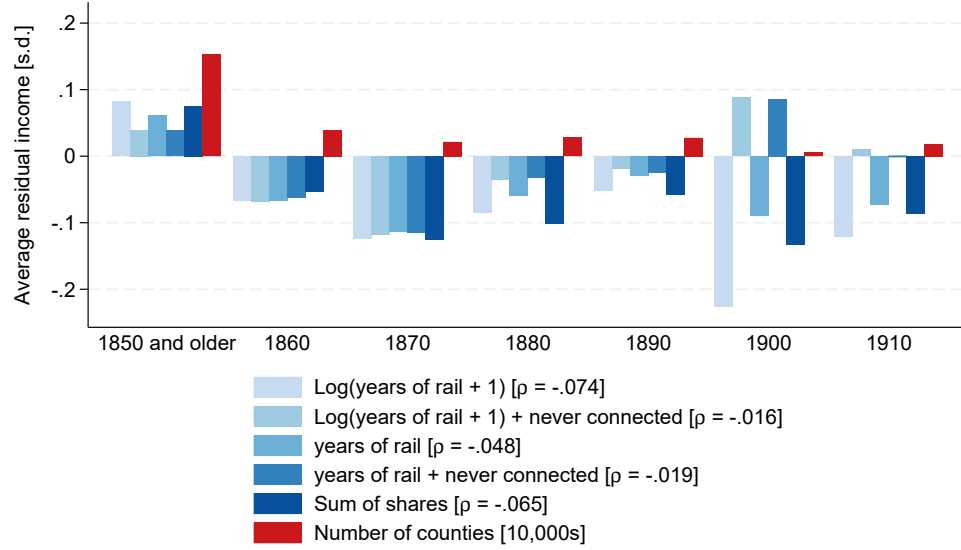
(b) Log income per capita in 2000

Notes: Average residualized and standardized a) instrument and b) log income per capita in 2000 by decade of railroad connection, for different sets of controls. Each specification includes the baseline controls of Table 4 of SNQ (excluding $\log(\text{years of rail}+1)$, unless noted otherwise) and the controls in the legend; “Sum of shares” controls for μ_i . The legend reports the correlation between the residual and the decade of railroad connection among ever-connected counties in square brackets. Red bars show the number of counties per decade.

Figure 3: Residualized instrument and log income per capita by decade of establishment



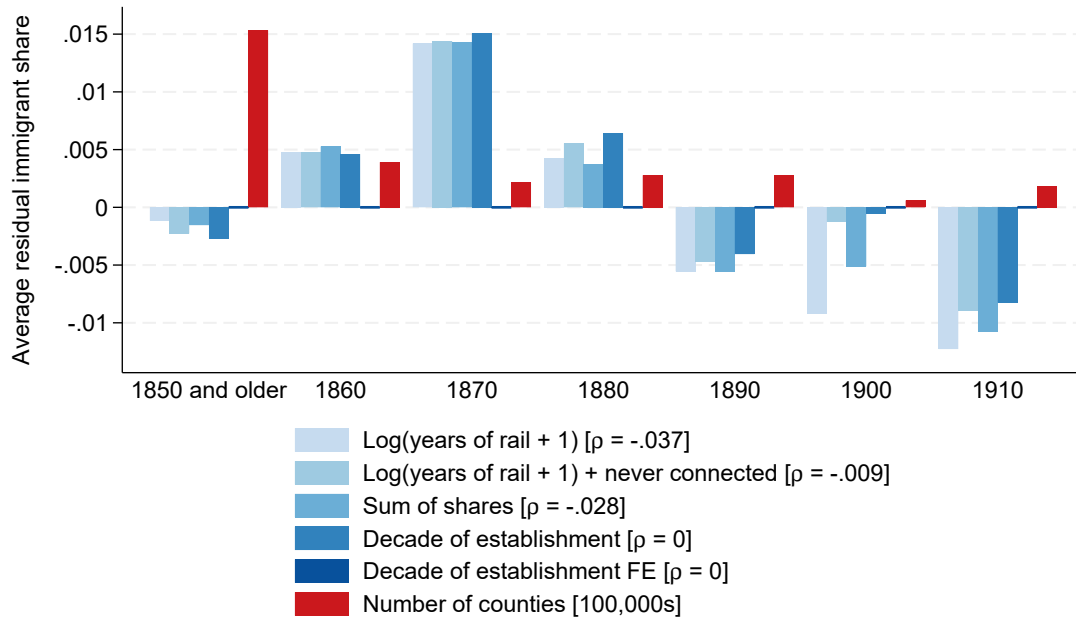
(a) Instrument



(b) Log income per capita in 2000

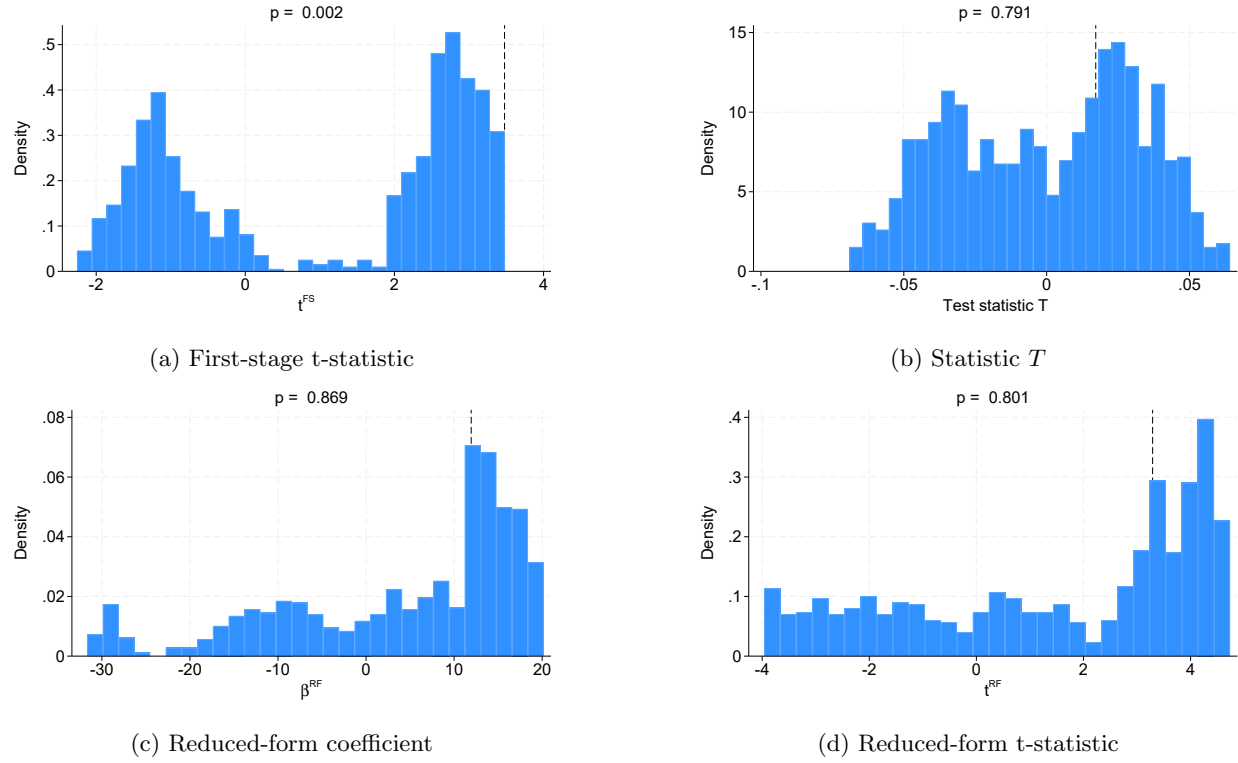
Notes: As Figure 2, by decade of county establishment. The legend reports the correlation between the residual and the decade of establishment in square brackets.

Figure 4: Residualized average immigrant share (1860–1920) by decade of establishment



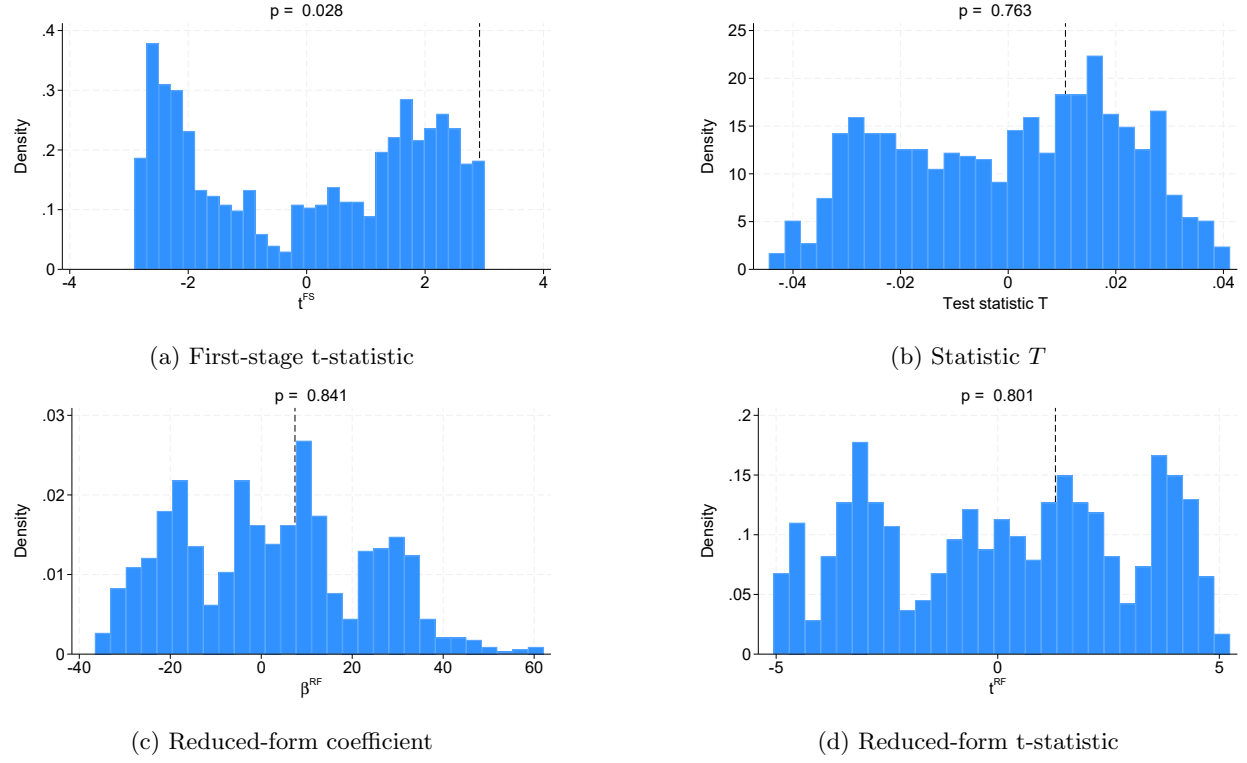
Notes: Average residualized immigrant share between 1860 and 1920 by decade of county establishment, for different sets of controls. Each specification includes the baseline controls of Table 4 of SNQ (excluding *log(years of rail+1)*, unless noted otherwise) and the controls in the legend. The legend reports the correlation between the residual and the decade of establishment in square brackets. Red bars show the number of counties per decade.

Figure 5: Randomization-inference distributions, SNQ's baseline specification, permuted shocks



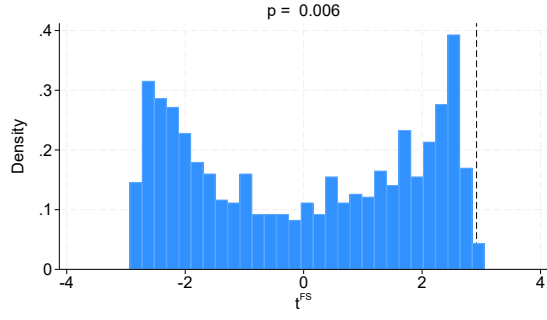
Notes: Distributions of the test statistics over 999 random permutations of the seven decadal shocks across decades, for the specification in Column 1 of Table 1 (SNQ's baseline controls, including $\log(\text{years of rail}+1)$). Dashed lines mark the statistic for the realized shocks; p-values are shown above each panel. The statistic T is defined in Section 4 with $b = 0$. t-statistics use Conley standard errors with a five-degree window.

Figure 6: Randomization-inference distributions, controlling for μ_i , permuted shocks

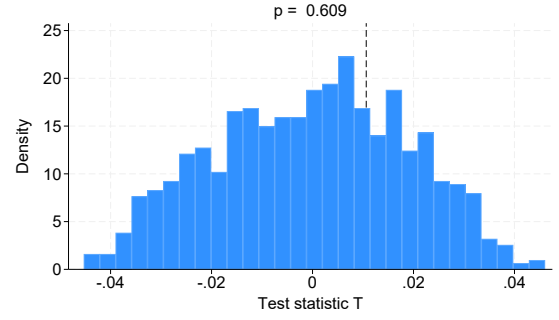


Notes: As Figure 5, for the specification in Column 7 of Table 1, which controls for μ_i instead of $\log(\text{years of rail}+1)$.

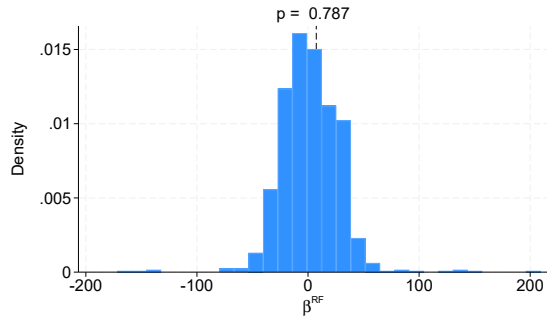
Figure 7: Randomization-inference distributions, controlling for μ_i , shocks drawn with replacement



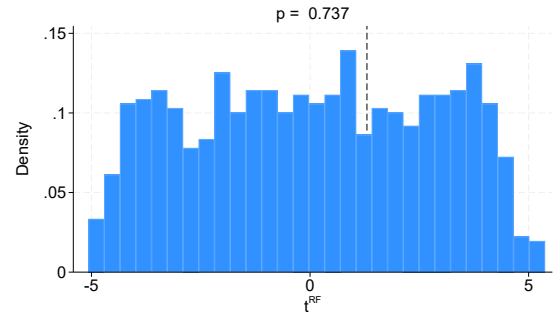
(a) First-stage t-statistic



(b) Statistic T



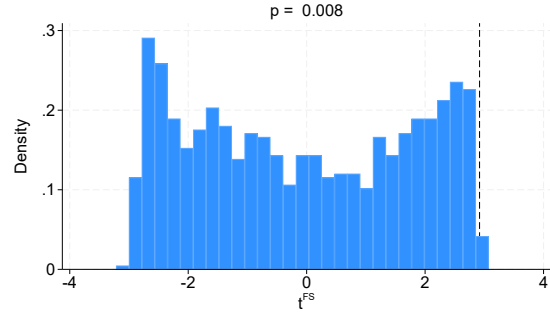
(c) Reduced-form coefficient



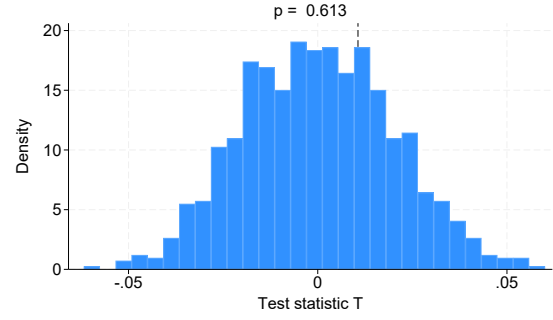
(d) Reduced-form t-statistic

Notes: As Figure 6, with the seven shocks drawn from the realized shocks with replacement.

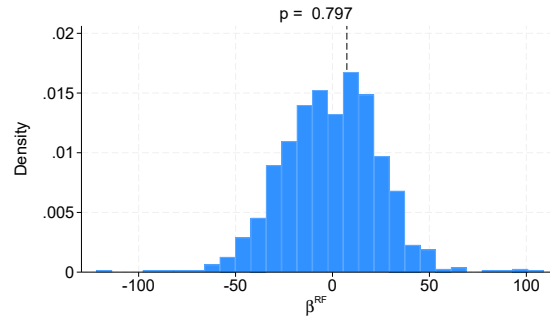
Figure 8: Randomization-inference distributions, controlling for μ_i , normally distributed shocks



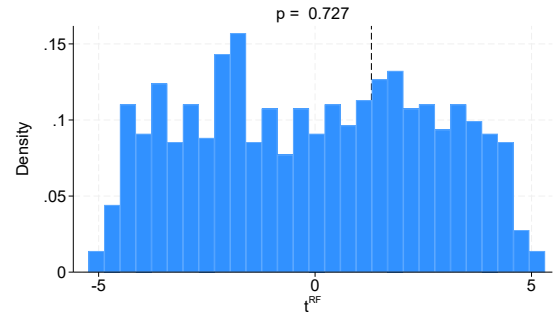
(a) First-stage t-statistic



(b) Statistic T



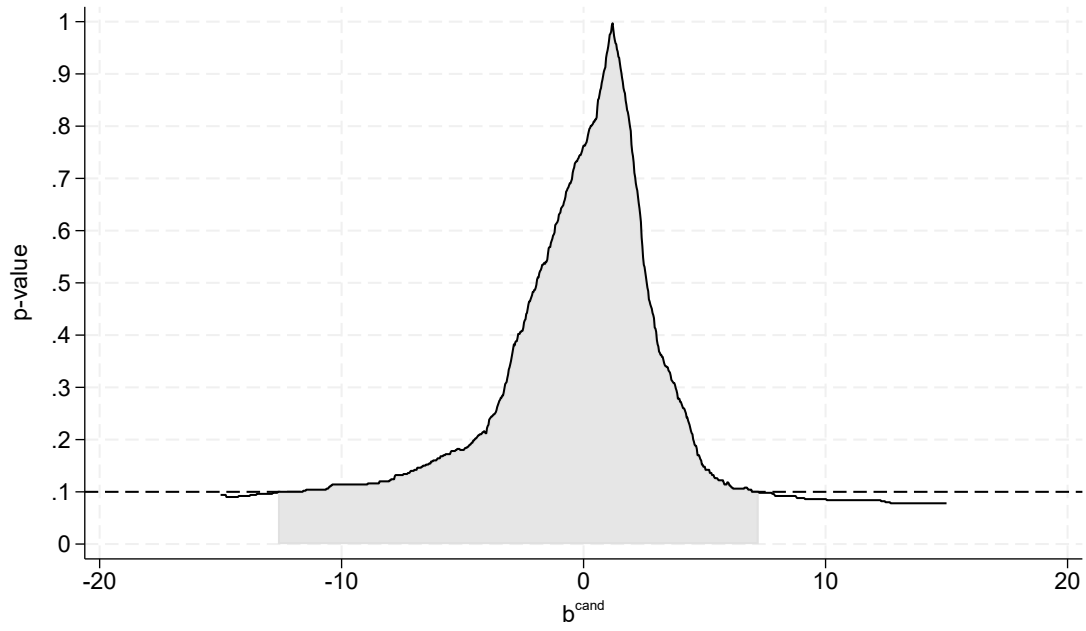
(c) Reduced-form coefficient



(d) Reduced-form t-statistic

Notes: As Figure 6, with the seven shocks drawn from a normal distribution with the mean (0.076) and standard deviation (0.025) of the realized shocks.

Figure 9: Randomization-inference confidence interval, controlling for μ_i



Notes: Randomization-inference p-values based on the statistic T for candidate effects b on a grid of 1,001 values between -15 and 15 , for the specification in Column 7 of Table 1. The instrument is recalculated for 999 random permutations of the seven decadal shocks. The shaded area marks candidate values not rejected at the 10% level; the implied 90% confidence interval ranges from -12.6 to 7.2 .

References

- Abadie, Alberto, Susan Athey, Guido W Imbens, and Jeffrey M Wooldridge.** 2023. “When should you adjust standard errors for clustering?” *The Quarterly Journal of Economics*, 138(1): 1–35.
- Adao, Rodrigo, Michal Kolesár, and Eduardo Morales.** 2019. “Shift-share designs: Theory and inference.” *Quarterly Journal of Economics*, 134(4): 1949–2010.
- Adema, Joop.** 2026. “Immigrants and the Making of America: A Comment.” Conditionally accepted at the Review of Economic Studies.
- Andrews, Isaiah, James H. Stock, and Liyang Sun.** 2019. “Weak instruments in instrumental variables regression: Theory and practice.” *Annual Review of Economics*, 11: 727–753.
- Borusyak, Kirill, and Peter Hull.** 2023. “Nonrandom exposure to exogenous shocks.” *Econometrica*, 91(6): 2155–2185.
- Borusyak, Kirill, Peter Hull, and Xavier Jaravel.** 2022. “Quasi-experimental shift-share research designs.” *Review of Economic Studies*, 89(1): 181–213.
- Borusyak, Kirill, Peter Hull, and Xavier Jaravel.** 2025. “A practical guide to shift-share instruments.” *Journal of Economic Perspectives*, 39(1): 181–204.
- Conley, Timothy G.** 1999. “GMM estimation with cross-sectional dependence.” *Journal of Econometrics*, 92(1): 1–45.
- Escamilla-Guerrero, David, Andrea Papadia, and Ariell Zimran.** 2026. “The effects of immigration on agricultural development.” *Journal of Development Economics*, 183: 103861.
- Goldsmith-Pinkham, Paul, Isaac Sorkin, and Henry Swift.** 2020. “Bartik instruments: What, when, why, and how.” *American Economic Review*, 110(8): 2586–2624.
- Moulton, Brent R.** 1990. “An illustration of a pitfall in estimating the effects of aggregate variables on micro units.” *The Review of Economics and Statistics*, 72(2): 334–338.
- Sequeira, Sandra, Nathan Nunn, and Nancy Qian.** 2020. “Immigrants and the making of America.” *Review of Economic Studies*, 87(1): 382–419.