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Re-examining democratization's impact on foreign borrowing costs during the first era of globalization

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This article re-examines the association between democratization and the cost of borrowing abroad in the first era of globalization. Using two representative datasets the literature offers but employing an improved method for panel event study, we find that democratization's impact on the costs of foreign borrowing is uncertain. In one case, the estimated coefficients are similar to the sign and magnitude of the original study but with larger standard errors, rendering the impact statistically insignificant. In the other case, the estimated coefficients hover around zero and are not statistically different from zero.

Keywords: franchise, democratization, autocratization, interest rate spread, panel event study

JEL classification: F02, F21, N10

I

Many studies have targeted the impacts of political institutions on financial markets. We can divide them into two strands. One presumes that market interest rates depend on political institutions, while the other is more skeptical.

The first strand, which claims a link between political institutions and financial markets, suggests two hypotheses according to which dimensions of democracy

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they are confronting. While the early literature targets the deepening of political contestation between rival elites, the recent literature looks into extending political participation to non-elites. Both periods of studies arrive at opposite conclusions.

The early literature emphasizes the weakening of absolute monarchies and the emergence of parliaments as placing limits on arbitrary expropriation by a government (North and Weingast 1989). Such checks on the executive branch commit the government to honor its debts, leading to a decline in interest rates on government borrowing. In this view, democratization benefits holders of sovereign debt and thus reduces the cost of borrowing.

Li (2019) offers an example by assessing the effect of the constitutional change in 1689 on the British government's credibility. She finds that the government's creditworthiness significantly improved under parliamentary supremacy, and government debt yields fell substantially. Acemoglu *et al.* (2019) provide evidence that democracy positively affects GDP per capita. They do not explore the democracy–financial market nexus straightaway. However, a favorable macroeconomic environment may engender a credibility bonus and thus speak against an interest rate premium. Objections come from Sussman and Yafeh (2006), who find that the British political institutional reform of 1689 did not immediately impact government borrowing costs.

Recent redistribution theories of democratization suggest that franchise extension, which transfers political power from capital to labor, increases sovereign risk for three reasons (Boix 2003; Acemoglu and Robinson 2005; Dasgupta and Ziblatt 2022). First, labor has less vested interest in debt payment and thus is more inclined to default. Second, expanding political participation to the mass electorate opens the door to reckless fiscal and monetary policies, thus detrimental to government bond markets. Third, democratization contributes to distributive conflict between elites and workers that could result in political uncertainty and instability and harm sovereign debt. All three factors cause borrowing costs to increase.

A second stream of literature, representing the skeptical view, questions whether financial markets would respond in the short run to institutional reforms because it takes time for new institutions to gain credibility, or any reforms may be followed by turmoil or setbacks. Using 18 emerging market countries from 1870 to 1913, Mauro *et al.* (2006) find that sovereign bond spreads were primarily driven by violent conflict or political instability. Institutional changes, such as links to the British Empire (Ferguson and Schularick 2006), the gold standard (as a Good Housekeeping Seal of Approval, Bordo and Rockoff 1996) and the protection of property rights (North and Weingast 1989), rarely elicited an immediate response from financial markets.

Hence, the impact of political institutions on financial markets remains unsettled. Recent research using panel event study finds that democratization increases a capital-importing country's cost of borrowing abroad during the first era of globalization. The empirical evidence seems consistent with the first strand (that political institutions affect financial markets), especially the redistribution theories of democratization. For example, Dasgupta and Ziblatt (2022) find that franchise extensions significantly increase sovereign bond yields. Tuncer and Weller (2022) show that democracies are associated with higher country risk than autocracies. However, the methodology these studies employ is problematic when there is heterogeneity in treatment effects. They also do not consider that democratization is a process that may involve multiple changes in the polity and even backsliding.

In this article, we use new methods from the panel event study literature, an area in which methodological progress has been made in the last few years, to re-examine the question at hand. The methods are an improvement over those in other studies that ignore the possibility of heterogeneity in treatment effects and reversals in the democratization process. We first employ an identical model specification and the same data but follow different estimation techniques. Second, we conduct further robustness checks. We do not find a statistically strong connection between democratization and the cost for one country to borrow abroad, as suggested by the first strand of the literature. In contrast, using the newly constructed geopolitical risk index by Dario and Iacoviello (2022), we do find that news about instability and wars strongly affects government bond yields, as Mauro *et al.* (2006) suggest.

We thus conclude that political economy theories may attribute too much importance to political institutions. Figure 1 shows China's government bond spread between 1877 and 1913. The figure also shows China's political regime, where 1 indicates true autocracy, and 2 indicates presidential democracy. The data source is Tuncer and Weller (2022). China's borrowing costs declined during this period, going from 4.6 percent in 1877 to 1.6 percent in 1913. During the same period, its political regime hardly changed before 1913, when the political regime changed from true autocracy to presidential democracy. Our findings complement and echo those of Mauro *et al.* (2006, ch. 3), who, using Japan in the Meiji Period (1868–1912), show that the impacts of political institutions on the risk premium are insignificant. Adopting the gold standard in 1897 was the only institutional reform, together with Japan's military victory over Russia in the Russo-Japanese War (1904–5), that led to a lower risk premium.

In the same figure, we also show the geopolitical risk index constructed by Dario and Iacoviello (2022). This index started in 1900 and has a shorter period. Yet, compared to the political regime, it relates more to China's government borrowing costs. The figure also shows China's external debt payments relative to its government revenues for selected years in which data are available. Debt payments (as a percentage of fiscal revenues) declined from 20 percent in 1899 to 11 percent in 1911. Like other emerging markets, China's government bond spreads displayed a clear downward trend after 1900 (Flandreau and Zumer 2004).

Our findings are not surprising from the sovereign debt default literature perspective, as shown by Bulow and Rogoff (1989) and Ritschl (2012). Investors care about the maximum committed payments extractable from a country, which is consistent with the findings of Flandreau and Zumer (2004) that the ratio of interest service to government revenue and uprisings and wars are important determinants of interest spreads. In his comparative study of China (1851–1911), England (1642–1752) and Japan (1868–95) during their transition to a modern fiscal state, He (2013) suggests

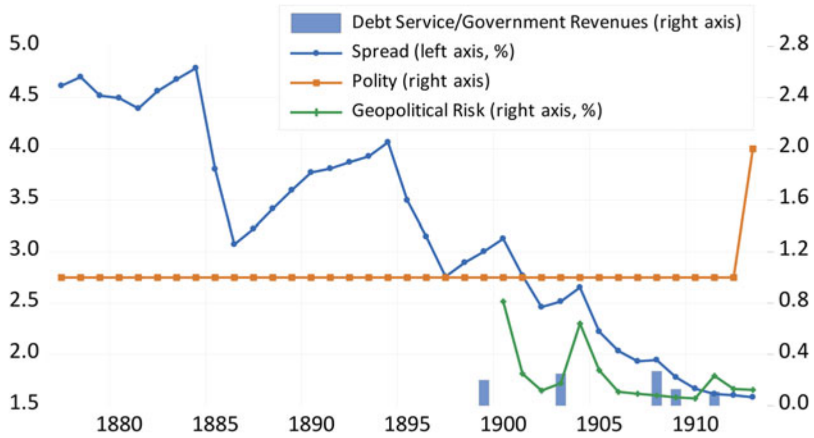

Figure 1. *China's government bond spread, political regime, geopolitical risk and ratio of debt service to government revenues from 1877 to 1913*

Notes: The horizontal axis goes from 1877 to 1913. For political regime, 1 indicates 'true autocracy', and 2 indicates 'presidential democracy'.

Source: China's government bond spread and political regime are taken from Tuncer and Weller (2022), geopolitical risk index is taken from Dario and Iacoviello (2022), debt service is taken from Chen (1995) and government revenues are taken from Zhou (2023).

that the creditworthiness of a government in financial markets' eyes rested on its ability to extract and mobilize indirect taxes from domestic consumption and customs. Thus, the ultimate question is whether political institutions help raise tax revenues committed to debt payments. The answer may be case-dependent. For example, representation (Neo-institutionalists' explanation) may not be a necessary condition for high tax revenues, as the high taxes were imposed on consumers who faced the collective action problem in organizing anti-tax campaigns, and England's Parliament at the time was dominated by landowners (Mathias 1959).

II

We use the dataset of Dasgupta and Ziblatt (2022), which covers 26 major countries in Europe and the Americas between 1800 and 1920. Table A1 in the online Appendix shows the list of sample countries.

Since we are interested in the relationship between democratization and government bond yields under the classical gold standard (the first period of globalization), investigating this issue requires enough heterogeneity among countries. The dataset of Dasgupta and Ziblatt (2022) includes the core European countries that were lending to the rest of the world. These countries also had a rather stable political system during the classical gold standard. The dataset also covers countries that heavily borrowed abroad at that time and experienced important changes in their political system. They were mostly independent (non-colonial) countries of the time in Western

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Europe, Latin America and Eastern Europe. Including Latin America is important for anyone interested in the link between international borrowing and the political system in the late nineteenth century because this is where the question was most urgent.

Dasgupta and Ziblatt (2022) employ franchise as the political-institutional variable. We also use the dataset of Tuncer and Weller (2022), who instead employ institutional arrangements to identify polity regimes. More will be explained below.

III

Dasgupta and Ziblatt (2022) conducted a panel event study to explore the impact of the passage of an event (franchise reform) over time.[1] For ease of explanation, we follow Clarke and Tapia-Schythe (2021) and consider a staggered assignment design, which means that units can adopt treatments at different times. Once treated, they remain treated after that.

Given the panel data, let $Event_i$ be a variable that records the time t when country (group) i adopts the treatment. The regression for the panel event study is as follows:

$$y_{i,t} = \alpha + \sum_{k=2}^{K} \gamma_k (\mathrm{Lead}^k)_{i,t} + \sum_{j=1}^{J} \beta_j (\mathrm{Lag}^j)_{i,t} + X_{i,t}\Gamma + \theta_i + \delta_t + \varepsilon_{i,t}, \qquad (1)$$

where $y_{i,t}$ is the outcome, $X_{i,t}$ are control variables, θ_i and δ_t are the country and time fixed effects, respectively, and $\varepsilon_{i,t}$ is a stochastic error term. Lead^k and Lag^j are a set of dummy variables that indicate the number of periods that country i is away from the passage of the event. Coefficients γ_k and β_j capture the difference between treated and control countries, compared to the difference in the omitted base period, which is the first lead (one period prior to the treatment). These coefficients are typically interpreted as a measure of the treatment effects at different lengths of exposure to the treatment. To obtain unbiased estimations of the post-event treatment effects, it is assumed that in the absence of treatment, the treated and control countries would have maintained similar differences to those in the base period (the so-called 'parallel trends assumption').

A panel event study has two advantages over the standard regular two-way fixed effect model. First, it allows one to inspect whether the assumption of parallel trends holds in the pre-treatment periods. Second, it allows one to examine the dynamics of the treatment effects, such as whether the effects are transitory or permanent.

In Dasgupta and Ziblatt (2022), the outcome variable is the natural logarithm of the sovereign bond yields. The event is franchise extension, defined as any reform extending the adult population's share of the right to vote by more than five percentage

[1] It is called 'pooled event study design' in the paper.

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points. The event (treatment) is thus binary and staggered. Dasgupta and Ziblatt (2022) set both K and J equal to 20. Instead of the first lag, the omitted base period they use is the period treated. OLS regression estimates do analysis.

The method of Dasgupta and Ziblatt (2022) has two defects. First, the OLS estimates are *contaminated* and do not give a clear causal interpretation when there is heterogeneity in treatment effects. Second, the method does not allow for reversal in the treatment, which can happen in the real world. We will talk about these two problems individually and explain how the latest studies deal with them.

Following Clarke's exposition (2022), we define cohorts as countries (groups) that share a common treatment adoption date. The average treatment effect on the treated for a specific cohort (relative to never being treated) is computed as:

$$CATT_{e,l} = E[Y^1_{i,e+l} - Y^0_{i,e+l} | E_i = e], \tag{2}$$

where e denotes the adoption date, l denotes periods of lag or lead relative to the adoption date, and E_i stores the adoption date for each country. Y^1 (Y^0) is the potential outcome when the country is treated (untreated). Note that the expectation operator E sums over all countries that adopt treatment on date e. Thus, $CATT_{e,l}$ is defined as a basic building block for treatment effects.

Suppose we are interested in the coefficient on the j lags in the panel event study. Sun and Abraham (2021) show that β_j in equation (1) is equal to:

$$\beta_j = \sum_e \omega_{e,j} CATT_{e,j} + \sum_{j' \neq j, j' \geq 0} \sum_e \omega_{e,j'} CATT_{e,j'}, \tag{3}$$

where $\omega_{e,j}$ are weights that sum to one, and $\omega_{e,j'}$ are weights that sum to zero. In other words, the coefficient of equation (1) will not recover a weighted estimate of $CATT_{e,j}$ for this lag length, which is the parameter of relevance. Instead, the coefficient is contaminated by $CATT_{e,j'}$, which are the treatment effects from other lags j'. This is true even when the assumption of parallel trends and the assumption of no anticipatory behavior before treatment hold. The second term in the right-hand side of equation (3) remains unless one imposes treatment effect homogeneity. For a different exhibition of the same issue, see Roth *et al.* (2023, pp. 2225–6).

The issue mentioned above complicates the interpretation of the estimated coefficients. It renders the results of the test of the parallel trends assumption questionable. De Chaisemartin and D'Haultfoeuille (2020, 2024), Sun and Abraham (2021) and Callaway and Sant'Anna (2021) offer solutions to the issues identified by Sun and Abraham (2021), where each study uses a specific building block as its framework. The online Appendix describes the solutions offered by Sun and Abraham (2021), Callaway and Sant'Anna (2021) and Borusyak *et al.* (2024), respectively.

We employ the heterogeneity-robust DID estimators that De Chaisemartin and D'Haultfoeuille propose and apply them to our cases where the treatment is either non-binary or non-staggered and where the lagged treatments may affect the outcome. Our exposition of the method follows the working paper versions of De Chaisemartin and D'Haultfoeuille (2024). Although it takes longer, it fits our purpose of explaining the advantages of that method better.

De Chaisemartin and D'Haultfoeuille (2020, 2024) define an event as the period where a country's treatment changes for the first time. Let F_i denote the period when country i's treatment changes for the first time. De Chaisemartin and D'Haultfoeuille (2020, 2024) use $\delta_{i,l}$ defined below as their building block:

$$\delta_{i,l} = E[Y_{i,F_i+l} - Y_{i,F_i+l}(D_{i,1}, \ldots, D_{i,1})], \tag{4}$$

which is the expected difference between country i's actual outcome in period $F_i + l$ and the counterfactual outcome if the country's treatment had been equal to its period-one value from period one to $F_i + l$.

Consider a country that is untreated in the first period. Let $T_u = \max_{i:D_{i,1}=0} F_i - 1$ (the last period when there is still a group that has been untreated since period one) and $N_l^1 = \sum_{i:D_{i,1}=0, F_i \leq T_u - l} 1$ (the number of countries reaching l periods after their first treatment at or before T_u). The measure of interest that is analogous to β_j in equation 1 and to what De Chaisemartin and D'Haultfoeuille (2020, 2024) propose is:

$$\delta_{+,l} = \sum_{i:D_{i,1}=0, F_i \leq T_u - l} \frac{1}{N_l^1} \delta_{i,l}, \tag{5}$$

which is the average cumulative effect of having changed treatment for the first time l periods ago, across all countries that are initially untreated and that get treated for the first time at least l periods before T_u. They develop an unbiased estimator, $DID_{+,l}$, for $\delta_{+,l}$. Mimicking the estimator, they also propose a placebo estimator, $DID^{pl}_{+,l}$, which can test for whether trends are parallel over several periods. Please refer to De Chaisemartin and D'Haultfoeuille (2020, 2024) for details of $DID_{+,l}$ and $DID^{pl}_{+,l}$.

We next consider a country which is treated in the first period. Parameters of interest and estimators for initially treated countries, say $\delta_{-,l}$ and $DID_{-,l}$, can be defined symmetrically. A DID_l estimator, which is an unbiased estimate for the average effect of having received treatment for $l + 1$ periods across both initially untreated and initially untreated countries, is obtained by taking an average of the $DID_{+,l}$ and $DID_{-,l}$ estimators. The weights are the relative number of observations that $DID_{+,l}$ and $DID_{-,l}$ use.

Defining an event as the period when a country's treatment changes for the first time (rather than the period when a country first gets treated) allows De Chaisemartin and D'Haultfoeuille (2020, 2024) to consider the treatment effect when treatment may be dropped (or in the continuous case when intensity of treatment may decline), as $\delta_{-,l}$ and $DID_{-,l}$ show. The DID_l estimator thus helps to deal with the second methodological concern that Dasgupta and Ziblatt (2022) encounter. It can be directly applied outside of the binary-and-staggered treatments and has a clear meaning: it shows the average results of changing treatments for the first time l periods ago. Since they provide the first DID estimators of instantaneous and dynamic treatment effects that are robust to heterogeneous effects and that can be used in general designs, in the sense that the estimators can be used with non-staggered binary or discrete treatments, in what follows, we use the estimators of De Chaisemartin and D'Haultfoeuille. Their approach produces an event-study graph, with the time to the first treatment change on the x-axis, the DID estimates on the y-axis to the right of zero and placebo estimates on the y-axis to the left of zero.

IV

Replication

Dasgupta and Ziblatt (2022) define franchise extensions as reforms that extend the right to vote by more than five percentage points and analogously for franchise contractions. Figure A1 in the online Appendix shows each country's franchise extension. Most countries have more than one franchise extension. For example, the United Kingdom had four franchise extensions during the sample period. Figure A2 shows franchise contraction. Here, one sees that franchise contraction (reversal) is not rare. For example, during the sample period, Chile experienced three franchise contractions. Figures A1 and A2 show that the franchise variable is not binary or staggered. We cannot use the estimators designed for binary and staggered treatment to determine how franchise affects government bond yields.

We combine franchise extensions and contractions to create a franchise variable, as Figure 2 shows. If this variable goes up (down) by one unit, then the country experienced a franchise extension (contraction) in that period. Using this more properly defined franchise variable and the same government bond yields taken from Dasgupta and Ziblatt (2022), we examine whether changes in franchises have any significant impact on the government bond yields.

Figure 3 shows the instantaneous and dynamic effects of a change in franchise. The estimated placebos are small as well as individually and jointly insignificant. The joint null of all placebos produces an F-test with a p-value of 0.645, implying that the assumption of parallel trends holds. Figure 3 shows that changes in franchise cause a sizable jump in government bond yields from the treatment period onward – one that hovers around the conventional statistical significance in the first four treated periods (denoted by 0 to 3 in the x-axis of Figure 3). After that, the impacts of franchise changes remain stable in magnitude, but the standard errors get larger.

To synthesize our results, we average the effects of the franchise variable from the year the franchise variable first changes to ten years after that. The estimated average effects are 0.185, and the standard errors are 0.121. On the one hand, this implies a

Figure 2. *Franchise extensions and contractions*
Notes: Dasgupta and Ziblatt (2022) define franchise extensions (contractions) as reforms that extend (reduce) the right to vote by more than five percentage points. Each subpanel in the figure shows franchise extensions and contractions for each country. The horizontal axis goes from 1800 to 1920. A vertical axis increasing (decreasing) by 1 means there is a franchise extension (contraction).
Source: Authors' calculation.

sizeable long-term increase in the government bond yields following franchise reform. On the other hand, the effects are noisily estimated because the new DID estimator we employ results in a much smaller uncontaminated subset of the data compared to the event study design that Dasgupta and Ziblatt (2022) utilize.

In sum, the franchise variable positively affects government bond yields because the coefficients are of comparable sign and magnitude to those of Dasgupta and Ziblatt (2022). However, when analyzed with the new method that discards contaminated controls, there is much noise in the estimates, making the effects of franchise reforms not statistically significant in the long run.

The franchising rate might not show how democratic a country is because unfair elections in authoritarian countries can still involve many people. These people have a voice on paper, but in reality they do not. Contemporary bondholders might also not know much about the number of adults who could vote during that time (Tuncer and Weller 2022). In such cases, institutional arrangements that could limit the executive branch's power might be more informative. This motivates us to experiment with the dataset of Tuncer and Weller (2022), who constructed a panel dataset covering the period 1870–1913

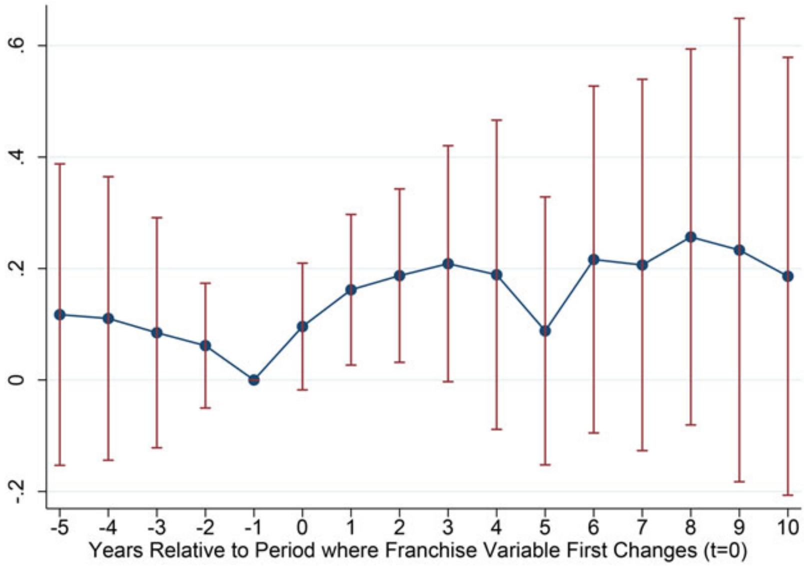

Figure 3. *Effects of franchise variable on government bond yields*
Notes: The horizontal axis denotes the years relative to the period when the franchise variable first changes (either as an extension or a contraction). The vertical axis reports the estimated effects of the franchise variable on the government bond yields and placebo estimates, using the data of Dasgupta and Ziblatt (2022) and DID estimation method of De Chaisemartin and D'Haultfoeuille (2020, 2024). Standard errors are clustered at the country level, while 95 percent confidence intervals, obtained from 200 bootstraps, appear in red.
Source: Authors' calculation.

for 27 independent countries that borrowed in London (see Table A2 for the sample countries).

The dependent variable Tuncer and Weller (2022) employ is government bond spreads, defined as the difference in yield between government bonds and the British consol. Tuncer and Weller (2022) characterize the institutional arrangements from the least to the most democratic as follows: (1) true autocracy, (2) presidential democracy, (3) constitutional monarchy and (4) true democracy. To keep to the framework of a binary and staggered design, they have to define two separate events: one is democratization, which happens when countries move to a more democratic polity, and the other is autocratization. Moreover, only those countries where polity changes and remains the same until the end of the sample period are included.

Since our method allows for discrete and non-staggered treatments, we can define a single event that denotes regime changes in the four institutional arrangements that Tuncer and Weller (2022) construct. In other words, our method allows a country to change its polity more than once. A country can become more democratic at some point but might return to being more autocratic later. This provides a more

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natural way to describe the polity changes we observe in the data. Figure 4 shows the treatment variable thus defined. Here, switches between the four polity regimes are not uncommon.

Figure 5 shows the results using the dataset and the Tuncer and Weller polity regimes (2022) but with a more appropriate method. Again, the estimated placebos are small, individually and jointly insignificant. The joint null of all placebos produces an F-test with a p-value of 0.540, implying that the assumption of parallel trends holds. In Figure 5, the impact of changes in the polity regime hovers around zero, is statistically insignificant and is much smaller than those reported in Figure 3 for changes in franchise. We average the effects of the polity regime from the year it first changed to ten years after that. The estimated average effects are 0.007, implying a tiny long-term increase in the spreads, and the estimated standard errors are 0.284. The estimated effects for the tenth and eleventh treated periods (denoted by 9 and 10 in the x-axis of Figure 5) are even negative, even though the standard errors also get larger, because of the reduced uncontaminated datapoints used for the estimation. In sum, moving towards a more democratic polity regime does not noticeably and positively affect the spread of government bonds. This result differs from Tuncer and Weller's (2022) finding that democratization is associated with sizably and statistically significantly higher spreads.[2]

Robustness

In the replication section, we employ an identical model specification and the same data as in other studies. Only the estimation technique is different. Now, we carry out further robustness checks step by step.

One concern is whether the experiences of emerging economies were different from those of the core countries. Thus, the first robustness check excludes France, Germany and the UK from the sample of Dasgupta and Ziblatt (2022), and we rerun the estimation. Figure A3 shows the results, which are similar to those of Figure 3. When the sample is restricted to emerging economies, the effects of a change in franchise in the first four treated periods become more statistically significant. From the year the franchise variable first changes to ten years after that, the estimated average effects are 0.232, and the estimated standard errors are 0.161. The effects, though sizable, are noisily estimated and statistically not significant. We do not carry out the same exercise as in Tuncer and Weller (2022), whose sample includes only emerging economies.

Mauro *et al.* (2006) find that news about instability and wars is significantly associated with more considerable spread changes, while events concerning reforms of political or economic institutions are not. Mauro *et al.*'s (2006) source for news is

[2] Following Tuncer and Weller (2022), our panel event study controls for years of domestic peace, years on default, years on gold, debt over tax revenue, exports over population and share of exports to Britain.

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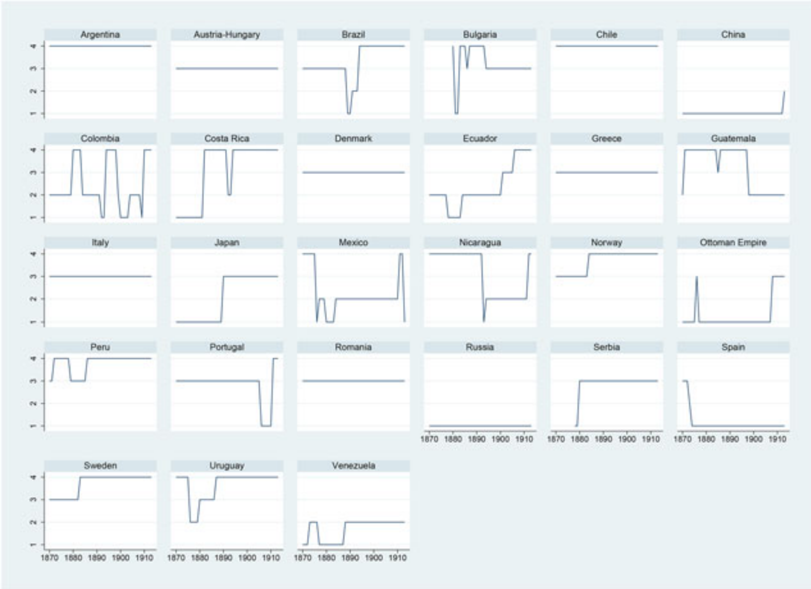

Figure 4. *Polity regimes*

Notes: Tuncer and Weller (2022) characterize the institutional arrangements (polity regimes) from the least to the most democratic as follows: (1) true autocracy, (2) presidential democracy, (3) constitutional monarchy and (4) true democracy. The figure shows franchise extensions and contractions for each country. The horizontal axis in each subpanel goes from 1870 to 1913. The value on the vertical axis denotes the type of polity regime.

Source: Authors' calculation.

the *London Times* and *Palmer's Index* when finding news related to their sample countries. They classify the related news articles into several categories, such as wars and violence and investor-friendly reforms and institutional changes. In the second robustness exercise, we employ the Geopolitical Risk (GPR) Index that Dario and Iacoviello (2022) constructed to check whether adverse geopolitical events and associated risks strongly impact government bond yields.

The GPR index is constructed from text-search results in the electronic archives of three newspapers: *Chicago Tribune*, *The New York Times* and *The Washington Post*. The sources do not include *The Times* (London). Much like the 'instability and wars' category of Mauro *et al.* (2006), the GPR index is based on events related to acts and threats of war and terror. GPR indices, which start from 1900, are available for 22 of the 26 sample countries in Dasgupta and Ziblatt (2022). We carry out the estimation for this more limited set of data that spans the years 1900–20.

The GPR index is constructed by calculating the share of articles mentioning the eight geopolitical risk categories and naming the country in question. We classify a country/ year observation as having news events about instability and wars if the monthly GPR index exceeds a 2 percent band in a given year. Though in a different

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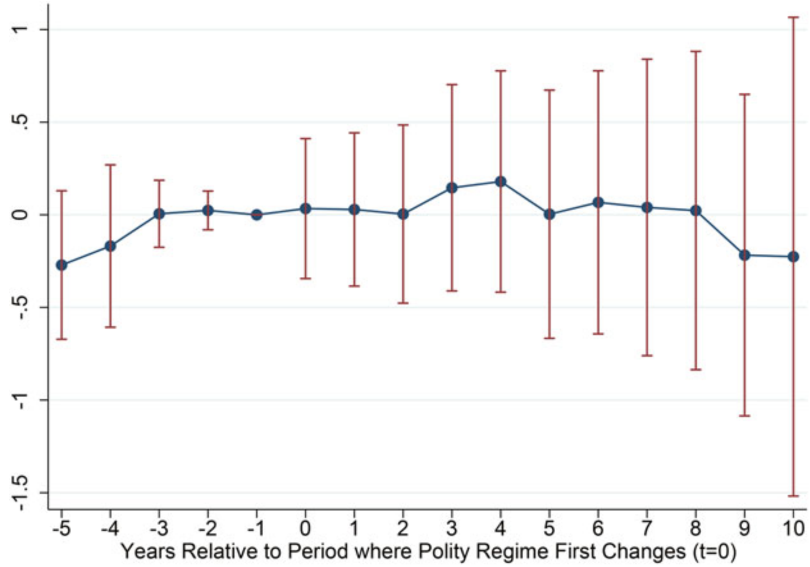

Figure 5. *Effects of polity regime on government bond spreads*
Notes: The horizontal axis denotes the years relative to the period when the polity regime first changes (either towards democratization or autocratization). The vertical axis reports the estimated effects of the polity regime on the government bond spreads and placebo estimates, using the data of Tuncer and Weller (2022) and DID estimation method of De Chaisemartin and D'Haultfoeuille (2020, 2024). Standard errors are clustered at the country level, while 95 percent confidence intervals, obtained from 200 bootstraps, appear in red.
Source: Authors' calculation.

context, Shambaugh (2004) uses the same 2 percent band to define periods of exchange rate stability. As Figure 6 shows, the band works well. It can identify the news regarding the outbreak and armistice of World War I for the belligerent countries. Figure 6 also highlights William McKinley's re-election in the US, the introduction of the US Gold Standard Act in 1900 and the 1904 Russian Revolution.

With a shorter sample, we estimate the effect of news about instability and wars on government bond yields for only the first seven years after specific news. Figure 7 shows that the effects become statistically significant two years after the news and hover around 0.2 between the fourth and the seventh years. The estimated average effects are 0.151, and the standard errors are 0.050. Using panel regression with fixed effects, Mauro *et al.* (2006, table 5.2) obtained an estimate between 0.232 and 0.359. Like Mauro *et al.* (2006), we also find that news about instability and wars strongly affects government bond yields.

We adopt the method that Liu *et al.* (2024) propose as a third robustness test. This method, which is an imputation approach, takes the treated observations as missing, uses untreated observations to build models and employs the estimated models to

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Figure 6. *News events on instability and wars*
Notes: The GPR index, taken from Dario and Iacoviello (2022), is constructed by calculating the share of articles mentioning the eight geopolitical risk categories and naming the country in question. We classify a country/year observation as having news events about instability and wars if the monthly GPR index exceeds a 2 percent band in a given year. Each subpanel in the figure shows the news events about instability and wars for each country. The horizontal axis goes from 1900 to 1920. If the vertical axis is equal to one, then there was a news event.
Source: Authors' calculation.

impute counterfactuals of treated observations. Like the estimator of De Chaisemartin and D'Haultfoeuille (2020, 2024), the method applies to cases concerning treatment effect heterogeneity. It may be statistically more efficient than De Chaisemartin and D'Haultfoeuille (2020, 2024) because the latter method drops many observations. Like De Chaisemartin and D'Haultfoeuille (2020, 2024), the method of Liu *et al.* (2024) allows for treatment reversals, but it applies only to the cases of dichotomous treatments.

We utilize a restricted sample of Figure 2 for which the franchise variable lies between zero and one because Liu *et al.*'s (2024) method applies only to binary treatments. The estimated effect of a change in the franchise, shown in Figure 4A of the online Appendix, remains statistically insignificant and even more so with a more efficient estimator. The placebo effect is statistically indistinguishable from zero (p-value = 0.404). The estimated average effect is 0.18, with a p-value of 0.08.

If weak checks on the executive branch either lessen the government's commitment to honor its debts or reduce distributive conflicts, investors would place more weight on democracy backsliding. In the next robustness check, we assume that a

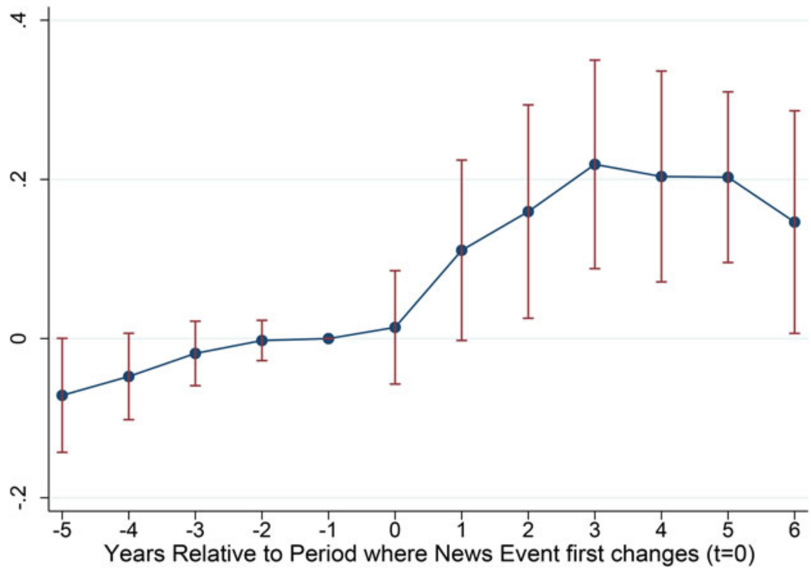

Figure 7. *Effects of news events on government bond yields*
Notes: The horizontal axis denotes the years relative to period when there are news events about instability and wars. The vertical axis reports the estimated effects of news about instability and wars on the government bond yields and placebo estimates, using the data of Dasgupta and Ziblatt (2022) and Dario and Iacoviello (2022) and DID estimation method of De Chaisemartin and D'Haultfoeuille (2020, 2024). Standard errors are clustered at the country level, while 95 percent confidence intervals, obtained from 200 bootstraps, appear in red.
Source: Authors' calculation.

franchise contraction is twice as important as a franchise extension and reconstruct our variable accordingly. While it does not directly address the threshold effects, it does allow some non-linear effects of franchise changes. The results, reported in Figure A5 of the online Appendix, are qualitatively similar to Figure 3.

To allow for diminishing effects, we let a franchise extension (or contraction) be lowered by half for the first three concessive periods following a franchise change. Flandreau and Zumer (2004) use a similar way to capture the memory effect of past defaults. We restrict diminishing effects to the first three years to preserve a reasonable number of control observations. The results, reported in Figure A6 of the online Appendix, are qualitatively similar to Figure 3.

V

Several studies suggest a strong and negative correlation between democratization and interest rate spread during the first age of globalization (1870–1914). Using the same datasets but improving the estimation methods, our research finds that the effects of franchise changes on government bond yields are of comparable sign and magnitude to the findings of Dasgupta and Ziblatt (2022), but the effects are estimated with large uncertainty. Lastly, news about instability and wars strongly affects government bond yields, as Mauro *et al.* (2006) suggest. Our findings imply that international investors care more about the political stability than political reforms of a borrowing country.

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Supplementary material

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