Commitment Institutions and Systematic Sovereign Risk

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Abstract

The link between institutional quality and sovereign credit risk is well established, but it is typically evaluated through average returns or alphas. This paper shows that institutional quality is instead priced through exposure to systematic risk. I develop a simple asset-pricing model in which commitment institutions reduce creditor losses under adverse shocks, sorting issuer loadings on a common sovereign-risk factor rather than abnormal returns. Using sovereign bonds traded on the London Stock Exchange between 1870 and 1907, I find that moving from the 25th to the 75th percentile of a composite index of commitment institutions (a 0.41-unit move on the 0–1 scale) is associated with a sovereign-factor loading that is 0.47 lower, which corresponds to roughly a 1.2 percentage-point gap in raw returns from factor exposure. The factor's left-tail realizations cluster during the period's major default waves. Gold-standard adherence—the era's leading monetary regime for signaling credibility—exhibits no comparable association. Instrumental-variables estimates point in the same direction and remain negative under weak-instrument-robust inference. This evidence indicates that commitment institutions shaped the covariance structure of sovereign bond returns rather than generating abnormal returns once common risks are priced.

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1 Introduction

A large body of research links institutional quality to sovereign credit risk. The evidence for the institutional channel in sovereign credit is typically evaluated through average returns, spreads, or risk-adjusted alphas. Asset-pricing evidence, however, indicates that sovereign credit risk is largely systematic (Longstaff et al., 2011; Augustin and Tédongap, 2016; Andrade et al., 2023), so the pricing margin that should reveal an institutional channel is exposure to priced aggregate risks rather than abnormal returns. Institutions, in turn, account for a substantial share of cross-country variation in macroeconomic volatility and crisis incidence (Acemoglu et al., 2003). As a result, standard yield- and alpha-based tests miss the channel through which institutions matter most.

This paper tests that proposition using a large cross-section of sovereign bonds traded on the London Stock Exchange between 1870 and 1907. The setting predates modern credit ratings (Flandreau, 2003), multilateral institutions, and contemporary monetary regimes, and cross-country variation in institutional constraints is large and well documented. Sovereigns borrowed in a relatively unified external market with a common investor base and broadly comparable debt instruments, which permits a comparison of returns across countries that faced similar global pricing conditions but differed in the extent to which political and legal arrangements constrained arbitrary state action against creditors.

The empirical analysis uses a two-step cross-sectional asset-pricing design. The first step estimates issuer-level alphas and factor loadings from a three-factor model. The second step examines whether cross-country variation in institutional quality explains the estimated factor exposures. Endogeneity is addressed with an instrumental-variables design that uses updated log settler mortality as an instrument for institutional quality (Acemoglu et al., 2001, 2012). Inference is robust to weak instruments, and the design includes geography and disease ecology as exogenous controls, addressing the principal econometric concerns raised in subsequent work (Albouy, 2012). The analysis also controls for competing historical explanations one at a time.

The main result is that institutional quality is strongly associated with sovereign-factor exposure but not with abnormal returns after accounting for common risk factors. In the OLS sample of 47 issuers with non-missing values of the institutional composite, an interquartile-

range move on the rescaled composite (0.41 units on the 0–1 scale) is associated with a 0.47 lower sovereign-factor loading, corresponding to a 1.2 percentage-point gap in raw returns attributable to factor exposure. The 26-issuer subsample with non-missing settler mortality, used for the instrumental-variables regression, yields a similar but slightly larger coefficient. Figure 1 shows the cross-sectional pattern: the institutional composite has a clearly steeper negative slope on the sovereign-factor loading than gold-standard adherence. This pattern is consistent with theories in which commitment institutions reduce the set of states in which governments can impair creditors' claims (North and Weingast, 1989) and with frameworks in which limited commitment is one of several frictions through which institutions shape sovereign risk exposures (Aguiar and Amador, 2021).

A second channel operates in tail states: stronger institutions reduce the level of tail-state losses passed to creditors. This level channel is identified primarily in the IV sample, where studies of this period locate the concentration of perceived default risk.

Contemporary accounts are consistent with that interpretation. Investors explicitly conditioned lending on political stability and constitutional constraints, and framed sovereign creditworthiness in terms of the credibility of repayment commitments (Withers, 1916; Corporation of Foreign Bondholders, 1877, 1894; Powell, 1910, p. 15). A simple asset-pricing model formalizes this mechanism.

The paper also examines gold-standard adherence, the leading alternative commitment device during that period. Because gold adoption was itself an endogenous outcome shaped by trade and financial linkages rather than an exogenous institutional commitment (Meissner, 2005), commitment institutions are the preferred measure of institutional quality. Commitment institutions account for sovereign-factor exposure more robustly than gold-standard adherence alone.

The result speaks to how sovereign debt is priced. Other evidence from this market shows that gold-standard adherence did not generate abnormal returns once common risks are accounted for, raising the question of which policy regime shapes sovereign borrowing costs (Alquist and Chabot, 2011). The evidence here points to commitment institutions. Cross-country variation in constraints on sovereign discretion sorts sovereign-factor exposure rather than abnormal returns. This finding recasts the institutional channel emphasized in studies of nineteenth-century sovereign spreads (Flandreau and Zumer, 2004; Mauro et al., 2006) as a systematic-risk margin

rather than as variation in average borrowing costs.

2 Related research

Three strands of research bear on this question: sovereign borrowing costs under the classical gold standard, sovereign debt and default, and institutions and the security of outside claims.

2.1 Sovereign borrowing costs under the classical gold standard

A prominent view in this strand of research is that adherence to gold served as a commitment device that lowered borrowing costs (Bordo and Rockoff, 1996). The London sovereign bond market was large and integrated, and the credibility-and-discipline argument fits into broader accounts of how external anchors interact with domestic policy (Obstfeld and Taylor, 2004; Mauro et al., 2006). Detailed historical work on sovereign spreads in this era documents that fundamentals such as fiscal performance and political stability mattered alongside formal monetary regimes, with substantial cross-country variation in borrowing costs even within the gold-standard window (Flandreau and Zumer, 2004). Open-economy macroeconomics places that intuition in a credibility-versus-flexibility framework. Pegging to an external anchor can mitigate time-inconsistency, but at the cost of monetary autonomy, and its suitability depends on a country's political and monetary institutions (Mishkin, 1999). Du et al. (2020) link cross-country variation in the bond-stock beta of local-currency sovereign debt to inflation credibility, evidence that monetary commitment shapes the systematic-risk properties of sovereign bonds, not just the level of borrowing costs. Gold-standard adoption was, in turn, itself endogenous to trade integration, financial linkages, and development (Meissner, 2005).

Existing evidence on the classical gold standard illustrates the limits of evaluating sovereign credibility through average returns or risk-adjusted alphas alone. Alquist and Chabot (2011) show that adherence to gold did not systematically generate abnormal returns after controlling for common risk factors, indicating that the monetary regime mattered less for sovereign credibility than the gold-standard view implies. This finding raises the question of whether observable commitment devices other than adherence to gold shaped the systematic risk exposures through which investors priced sovereign bonds.

2.2 Sovereign debt and default

In the limited-commitment framework of Eaton and Gersovitz (1981), lenders price sovereign debt based on the borrower's ex post incentive to repay. Recent syntheses emphasize that limited commitment operates alongside frictions such as incomplete state contingency, debt dilution, deadweight default costs, and political-economy distortions (Aguiar and Amador, 2021). A reputational account, in which sovereigns repay to preserve future market access (Tomz and Wright, 2013), is the leading alternative, though Bulow and Rogoff (1989) show that reputation alone cannot sustain sovereign borrowing. The analysis treats commitment institutions and reputation as complementary margins, since both operate through beliefs about future repayment. A broader survey concludes that the evidence is more supportive of accounts emphasizing domestic costs of default than of accounts centered on external sanctions or capital-market exclusion (Panizza et al., 2009). Across two centuries of external sovereign borrowing, default outcomes are continuous rather than binary, with creditor losses dispersed across episodes and ex-post excess returns persistently positive (Meyer et al., 2022).

Empirically, default risk and spreads reflect fundamentals and time-varying risk premia, with global factors quantitatively important (Pan and Singleton, 2008; Longstaff et al., 2011; Ang and Longstaff, 2013). Country fundamentals explain only a fraction of cross-sectional and time-series variation in emerging-market spreads (Eichengreen and Mody, 2000). U.S. expected consumption growth and consumption volatility account for a large share of common time variation in sovereign CDS spreads, and the resulting comovement is priced as compensation for exposure to U.S. macroeconomic risk (Augustin and Tédongap, 2016). Exposure to low-frequency global macroeconomic conditions has been identified as a separate priced channel (Andrade et al., 2023). Late-nineteenth-century sovereign bond returns also load on systematic risk factors that are priced (Alquist, 2010), and foreign securities formed an important component of British investors' feasible portfolios in this integrated capital market (Chabot and Kurz, 2010).

2.3 Institutions and the security of outside claims

Institutional quality has been linked to a range of macroeconomic outcomes, including cross-country differences in growth volatility and crisis incidence (Acemoglu et al., 2003). For sovereign

credit, institutions support credible commitment by constraining arbitrary state action (North and Weingast, 1989; Eichengreen et al., 2021), with within-country evidence from nineteenth-century Brazil showing that political-institutional changes and the delegation of monetary authority to a quasi-independent bank produced persistent reductions in the sovereign risk premium on bonds traded in London (Summerhill, 2008). Legal rules also shape the protection of outside claims and the development of external finance (La Porta et al., 1998; Shleifer and Wolfenzon, 2002; Acemoglu and Johnson, 2005). In international and asset-pricing settings, expropriation risk and weak investor protection raise required returns on outside claims (Stulz, 2005; Albuquerque and Wang, 2008). The most direct antecedent to this paper uses data from the same era to show that firms in countries with weaker property rights institutions faced a higher cost of equity capital (Alquist et al., 2022). Political rights also enter international corporate bond pricing more broadly (Qi et al., 2010). Rajan and Zingales (2003) sharpen the view by emphasizing institutions as mechanisms that filter political pressures, which for sovereign debt suggests that constitutional and legal constraints matter most when repayment becomes politically costly. Such constraints limit the government's ability to shift losses onto foreign creditors in states of distress.

This paper develops these ideas by asking whether institutions sort on sovereign-factor exposure rather than on abnormal returns.

3 Conceptual framework and model

This section presents a simple asset-pricing model that matches the paper's empirical design.

Institutions are interpreted as constraints that limit the sovereign's ability to alter debt contracts ex post. These constraints include constitutional rules, legislative checks, judicial limits, and related political arrangements that restrict discretionary policy action against creditors. A higher value of I_i thus corresponds to stronger constraints on opportunistic nonpayment.

Consider a small sovereign borrower indexed by i. At date 0, the sovereign issues a one-period bond with face value normalized to one. Let q_i denote the date-0 bond price. At date 1, the bond payoff depends on an aggregate shock z that is common across sovereigns and an idiosyncratic shock η_i. The aggregate shock captures the broad international conditions relevant for London investors, including episodes of tight global liquidity, weak risk appetite, or stress

in sovereign credit markets. The idiosyncratic shock captures country-specific fiscal, political, or external disturbances.

Let the gross payoff to investors be

$$X_i(z, \eta_i; I_i) = 1 - L_i(z, \eta_i; I_i), \qquad 0 \leq L_i(z, \eta_i; I_i) \leq 1,$$

where L_i is the fraction of promised payment lost through default, suspension, restructuring, inflationary erosion, or other impairment of creditor claims. Decompose the loss function as

$$L_i(z, \eta_i; I_i) = L_i^F(z, \eta_i; I_i) + \mathbf{1}\{z < z^*\}\, L_i^T(z, \eta_i; I_i),$$

where L_i^F is the part of creditor losses that admits a linear factor representation in expectation, and L_i^T is a tail-specific component active only when the aggregate state z falls below a threshold z^*.[1] Institutional quality reduces both components,

$$\frac{\partial L_i^F(z, \eta_i; I_i)}{\partial I_i} \leq 0 \qquad \text{and} \qquad \frac{\partial L_i^T(z, \eta_i; I_i)}{\partial I_i} \leq 0,$$

each with strict inequality in a nontrivial set of states.

Each component is also weakly decreasing in the aggregate state,

$$\frac{\partial L_i^F(z, \eta_i; I_i)}{\partial z} \leq 0 \qquad \text{and} \qquad \frac{\partial L_i^T(z, \eta_i; I_i)}{\partial z} \leq 0,$$

capturing the standard pattern that creditor losses are larger when international conditions deteriorate. The empirical regularity that defaults cluster in bad aggregate states, documented across two centuries of sovereign borrowing in Tomz and Wright (2007), is direct evidence of this monotonicity for the tail-specific component.

The tail-specific component is taken to be conditionally independent of f_{SOV} given $D = \mathbf{1}\{z < z^*\}$. The size of tail-specific losses, conditional on being in a tail state, is driven by issuer-level

[1] The two components correspond to two empirical forms of sovereign bond loss. L_i^F is a continuous mark-to-market impairment driven by changes in systematic risk premia. L_i^T is default-related impairment—missed payments, suspensions, restructurings, recovery shortfalls—concentrated in extreme aggregate states. Meyer et al. (2022) document an average haircut of 44% with a standard deviation of 30% across more than 300 external sovereign debt restructurings since 1815, supporting the treatment of L_i^T as a continuous, cross-sectionally dispersed object rather than a binary default indicator.

fundamentals rather than by the realized depth of the tail. Under this stipulation, L_i^T enters the linear factor representation as a tail-state level shift rather than as a slope shift, mapping the two channels into distinct regression objects: L_i^F identifies the sovereign-factor loading $\beta_{i,\mathrm{SOV}}$, and L_i^T identifies the tail-state intercept $\alpha_{i,\mathrm{tail}}$.

The covariance condition governs the factor-driven channel

$$\frac{\partial \mathrm{Cov}\left[L_i^F(z,\eta_i;I_i),\ f_{\mathrm{SOV}}\right]}{\partial I_i} \geq 0,$$

under which stronger institutions reduce the comovement of factor-driven losses with the sovereign factor. Equivalently, an issuer's L_i^F becomes less concentrated in adverse realizations of f_{SOV} as I_i rises. The analogous level condition governs the tail-specific channel

$$\frac{\partial \mathbb{E}\left[L_i^T(z,\eta_i;I_i)\,\middle|\, z<z^*\right]}{\partial I_i} \leq 0,$$

under which stronger institutions reduce the expected tail-specific loss conditional on being in a tail state. Both conditions admit a structural microfoundation in the cross-partial restriction $\partial^2 L_i^c/\partial I_i\,\partial z \geq 0$ for $c \in \{F,T\}$: institutional reductions in losses are concentrated state-by-state in bad aggregate states. The cross-partial nests each integrated condition and adds further state-by-state predictions, steepening the institutional slope on $\beta_{i,\mathrm{SOV}}$ in left-tail states and concentrating $\alpha_{i,\mathrm{tail}}$ in deeper tail realizations. The relative magnitudes of the two channels are an empirical question.

This reduced-form structure has counterparts in quantitative sovereign-default research: standard small-open-economy models generate default in low-z states because outside options are then most attractive (Arellano, 2008), and political-friction extensions tie selective default to the political costs the sovereign internalizes (Bocola et al., 2019).

Competitive British investors price sovereign bonds using the stochastic discount factor $m(z)$. The date-0 price satisfies

$$q_i(I_i) = \mathbb{E}[m(z)X_i(z,\eta_i;I_i)]\,.$$

The corresponding gross return is

$$R_i(z, \eta_i; I_i) = \frac{X_i(z, \eta_i; I_i)}{q_i(I_i)}.$$

If stronger institutions reduce expected creditor losses, then they raise bond prices.

To connect the model to the empirical design, suppose excess returns admit the linear factor representation

$$r^e_{i,t} = \alpha_i + \beta_{i,\text{SOV}} f_{\text{SOV},t} + \beta_{i,\text{MKT}} f_{\text{MKT},t} + \beta_{i,\text{TERM}} f_{\text{TERM},t} + \alpha_{i,\text{tail}} D_t + u_{i,t},$$

where $f_{\text{SOV},t}$ is the sovereign bond market factor, $f_{\text{MKT},t}$ is the London market factor, $f_{\text{TERM},t}$ is the British term premium factor, and $D_t = \mathbf{1}\{z_t < z^*\}$ is a left-tail indicator that activates the tail-specific component of losses. The unconditional intercept α_i summarizes the mean-pricing fit. The tail-state intercept $\alpha_{i,\text{tail}}$ summarizes the tail-specific component L_i^T scaled by the bond price. Idiosyncratic shocks η_i are absorbed by $u_{i,t}$. The empirical analysis first estimates issuer-level alphas and factor loadings from this time-series specification, then relates these estimates to institutional characteristics in the cross-section.

The model delivers two testable comparative statics.

Proposition 1 (Factor loading and institutional quality)**.** *Under the covariance condition on L_i^F, $\partial\beta_{i,\text{SOV}}/\partial I_i \leq 0$: stronger institutions lower the issuer's sovereign-factor loading.*

Proposition 2 (Tail-state intercept and institutional quality)**.** *Under the level condition on L_i^T, $\partial\alpha_{i,\text{tail}}/\partial I_i \geq 0$: stronger institutions raise the issuer's tail-state intercept, equivalently reducing the level of tail-specific losses passed to creditors.*

The propositions are sign predictions delivered by integrated conditions on the joint distribution of losses and aggregate states. The cross-partial restriction is a stronger structural primitive: it implies the integrated conditions but adds state-by-state predictions tested separately in Section 6. A rejection of the cross-partial on either component does not overturn the corresponding proposition; it bounds the structural form that delivers the comparative static.[2]

[2]The proofs and remarks on regularity conditions and on the relationship between the covariance condition, the level condition, and the cross-partial restriction appear in Appendix A.4.

The propositions map directly into the paper's cross-sectional tests. The benchmark specification asks whether institutions predict abnormal performance after controlling for common risks:

$$\alpha_i = a_\alpha + b_\alpha I_i + \Gamma'_\alpha Z_i + \varepsilon_{i,\alpha},$$

where Z_i is a vector of issuer-level controls. The core specification asks whether institutions predict estimated factor loadings:

$$\hat{\beta}_{i,k} = a_k + b_k I_i + \Gamma'_k Z_i + \varepsilon_{i,k}, \qquad k \in \{\text{SOV}, \text{MKT}, \text{TERM}\}.$$

The tail-state specification identifies $\alpha_{i,\text{tail}}$ from the interaction of D_t with I_i:

$$r^e_{i,t} = \alpha_i + \boldsymbol{\beta}'_i \boldsymbol{f}_t + \delta\, D_t + \zeta\, I_i\, D_t + (\text{slope interactions}) + u_{i,t},$$

so that $\partial \alpha_{i,\text{tail}} / \partial I_i$ is identified by ζ. Proposition 1 implies $b_{\text{SOV}} < 0$ and Proposition 2 implies $\zeta > 0$, with the relative magnitudes of the two channels an empirical question. Institutions may also reduce exposure to the London market factor if international stress states spill into sovereign distress. There is no strong reason in this framework for institutional quality to sort the term premium loading. A weak relation between institutions and the unconditional α_i does not imply that institutions were irrelevant for pricing. It may instead reflect that institutions operate primarily through the factor-loading and tail-state level channels rather than through unconditional abnormal performance.

4 Data, factor construction, and institutional measures

4.1 Sovereign bond returns and issuer portfolios

The sample consists of foreign sovereign and colonial bonds traded on the London Stock Exchange between 1870 and 1907. Prices are sampled at four-week intervals from the official quotation lists published in the *Money Market Review* and *The Economist*. I combine prices with coupon, issue, and default information to construct holding-period returns for each bond, and then aggregate bond-level returns to issuer-level portfolios. The cross-sectional tests use issuer-level average excess

returns, alphas, and factor loadings estimated from these portfolio returns.

This setting has two features useful for the empirical design. First, the issuers borrowed in a common external market. The London Stock Exchange supplied a shared trading venue, common quotation practices, and a broadly similar investor base across issuers, and by the late nineteenth century it was the principal global venue for international government debt (Michie, 1999; Neal and Davis, 2006; Mauro et al., 2006). Second, the securities are long-term sovereign or quasi-sovereign claims traded by investors facing the same global pricing conditions. As a result, cross-country variation in returns can be related to borrower characteristics without also comparing distinct market structures. Figure A.1 in the appendix maps the full issuer sample and the final regression sample.

The empirical analysis uses two sample definitions. The OLS sample is the maximal set of issuers with non-missing values of the relevant institutional measure (47 for the institutional composite, 45–53 across the other measures).[3] The IV sample is the 26-issuer subset of the OLS sample with non-missing settler mortality and the other historical controls used in the instrumental-variables and competing-explanation regressions. The OLS sample of 47 issuers is at least as large as the cross-sections used in modern asset-pricing studies of sovereign credit (26 in Longstaff et al. (2011), 38 in Augustin and Tédongap (2016), 43 in Andrade et al. (2023)). The OLS sample is used for the main reduced-form factor decompositions. The IV sample is used when settler mortality or the geography-and-disease, legal origin, and precolonial development controls are required. Table 1 reports summary statistics for the full sample (Panel A) and the IV sample (Panel B).[4]

Settler mortality is the binding constraint on the cross-section. The Acemoglu, Johnson, and Robinson series exists only for previously colonized territories, so 27 issuers—15 continental European sovereigns, 9 non-European sovereigns and territories outside its coverage or with missing precolonial controls, and 3 British colonies with missing precolonial or geographic controls—fall outside the regression sample.[5] The exclusion is driven by data coverage rather than by the

[3] 47 issuers for the institutional composite and judicial constraints, 45 for legislative constraints and the strict composite, and 53 for gold-standard adherence.

[4] For comparison, Meyer et al. (2022) report real excess returns over UK gilts of 2.6% for non-serial-defaulter and 4.1% for serial-defaulter external sovereigns over 1870–1913 in a separate long-run dataset.

[5] The 15 continental European sovereigns are Austria-Hungary, Belgium, Bulgaria, Denmark, France, Germany, Greece, Italy, the Netherlands, Norway, Portugal, Romania, Russia, Spain, and Sweden. The 9 non-European sovereigns and territories are China, Hawaii, Japan, Liberia, Morocco, the Orange Free State, the Ottoman Empire, Santo Domingo, and South Africa. The 3 British colonies are Fiji, Saint Lucia, and the Straits Settlements.

sovereign-factor relation.

Appendix Table A.4 compares the kept-26 and dropped-27 issuers on the analysis variables; differences in means are statistically insignificant for all of them. Appendix Table A.2 shows that bivariate sovereign-factor loading regressions in the OLS sample available for each institutional measure deliver the same sign and similar magnitudes as in the IV sample. The IV-sample restriction is required for instrumental-variables and competing-explanation regressions, not for the baseline relation.

4.2 Risk factors and time-series specification

The first-stage asset-pricing regression estimates issuer-level alphas and factor loadings from a three-factor model. For issuer i, excess returns satisfy

$$r^e_{i,t} = \alpha_i + \beta_{i,\mathrm{SOV}} f_{\mathrm{SOV},t} + \beta_{i,\mathrm{MKT}} f_{\mathrm{MKT},t} + \beta_{i,\mathrm{TERM}} f_{\mathrm{TERM},t} + u_{i,t}. \tag{1}$$

The sovereign bond market factor, $f_{\mathrm{SOV},t}$, is the excess return on the value-weighted portfolio of sovereign and colonial bonds, which proxies for common sovereign-credit conditions in the London market. The London market factor, $f_{\mathrm{MKT},t}$, is the excess return on the value-weighted portfolio of British and foreign equities traded in London, which measures broad equity-market conditions and risk appetite in the same trading venue. The term premium factor, $f_{\mathrm{TERM},t}$, is the excess return on British consols over the short-term London bill rate, which reflects interest-rate risk in British government debt. Appendix A.7 describes factor construction and reports robustness checks using leave-one-out and alternative sovereign-factor constructions, which assess whether the estimated sovereign-factor loadings are driven by own-return inclusion, value weighting, or the imperial composition of the baseline sovereign bond market factor.

The time-series estimates produce the main dependent variables for the cross-sectional analysis. The alpha, α_i, measures abnormal performance relative to the three common factors. The factor loadings, $\hat{\beta}_{i,\mathrm{SOV}}$, $\hat{\beta}_{i,\mathrm{MKT}}$, and $\hat{\beta}_{i,\mathrm{TERM}}$, measure the sensitivity of issuer returns to sovereign-risk, London-market, and term-premium shocks. The empirical question is whether institutional quality is associated mainly with alpha or with exposure to these common risks.

4.3 Institutional measures

The primary institutional measures are country-level constraints on sovereign discretion. The baseline measures are V-Dem's judicial-constraints and legislative-constraints indices on the executive, plus two composites built from them: a baseline composite that averages the two indices, falling back to whichever is available when one is missing, and a strict composite that requires both and covers a slightly smaller cross-section. Judicial constraints capture the extent to which the executive respects the constitution, complies with court rulings, and faces an independent judiciary; legislative constraints capture the extent to which legislatures and oversight agencies can question, investigate, and oversee the executive (V-Dem Institute, 2026). Polity's executive-constraints index appears as a supplementary measure in the first-stage figure, and gold-standard adherence—the share of four-week sample periods on gold—is included for comparison.[6]

All institutional measures in the cross-sectional regressions are constructed in two steps. First, the annual V-Dem or Polity series is aggregated to a country-level median over 1870–1907, reflecting historical evidence that markets responded only gradually to institutional changes as investors observed whether constraints were respected over time (Mauro et al., 2006, pp. 57–58). Second, the country-level medians are rescaled across the cross-section to the unit interval as $x_j = (\tilde{x}_j - \min_k \tilde{x}_k)/(\max_k \tilde{x}_k - \min_k \tilde{x}_k)$, so the issuer with the weakest constraints takes value 0 and the strongest takes value 1. OLS coefficients thus measure the gap between the cross-sectional extremes; in the text, I translate each coefficient into the more interpretable interquartile-range move on the rescaled composite, which is 0.41 in the OLS sample and $p_{75} - p_{25} = 0.713 - 0.317 = 0.396$ in the IV sample.

Ten of the 26 issuers are British colonies or possessions during 1870–1907, ranging from self-governing dominions to dependent colonies.[7] V-Dem applies a common scoring template across polities, so its indices record local checks within each imperial arrangement rather than constraints against an independent-state counterfactual. Pooling colonies and sovereigns in a single cross-section lets the institutional coefficient capture the sovereign/colony split and institutional varia-

[6] Codebook references, variables used, and the 1870–1907 aggregation procedure for the V-Dem and Polity series are in Appendix A.6.

[7] The issuers are Australia, British Guiana, Canada, Ceylon, Hong Kong, Jamaica, Mauritius, New Zealand, Newfoundland, and Trinidad. The sample spans self-governing dominions, in which domestic authority rested with elected legislatures and local cabinets, and dependent colonies, in which authority rested with a British-appointed governor and at most a limited local council (Ferguson and Schularick, 2006).

tion within each group (i.e., independent sovereigns and British colonies/possessions).

4.4 Instrumental variable and competing-explanation controls

The instrumental-variables specification uses updated log settler mortality as an instrument for institutional quality, following the logic of the settler-mortality design (Acemoglu et al., 2001). The baseline measure is the alternative "Naval Stations 2" series from Acemoglu et al. (2012), capped at 250 deaths per 1,000 per annum, used in updated work (Johnson, 2025). Because the settler-mortality design has been the subject of extensive debates over data construction and inference (Albouy, 2012; Acemoglu et al., 2012), I report weak-instrument-robust inference alongside the 2SLS estimates and use the IV evidence as disciplined support rather than as the sole basis for the paper's interpretation.

I also construct compact controls for competing explanations. A geography-and-disease endowments index combines absolute latitude, temperature, proximity to the coast, and malaria ecology. A British legal-origin indicator captures inherited legal background, a channel emphasized by law-and-endowments research in finance (Beck et al., 2003). A precolonial-development measure uses the log population density in 1500. A British colony/empire dummy variable indicates an issuer's imperial status, matching the documented empire effect on country risk in this period (Ferguson and Schularick, 2006). These variables are included one at a time because the cross-section is small and many historical controls are strongly correlated.

5 Regression evidence

This section presents the main empirical evidence linking institutional quality to the pricing of sovereign risk.

5.1 Benchmark evidence: average excess returns and alphas

The first tests examine whether institutional quality predicts issuer-level alphas or average excess returns. These regressions serve as benchmarks because the prior discussion of sovereign credibility has been framed in terms of lower borrowing costs. If institutions generate a risk-adjusted

return premium, stronger institutions should show lower alphas after conditioning on the common London factors.

Table 2 reports this benchmark. Panel A shows that average excess returns are negatively related to institutional quality, especially for judicial constraints and the composite measures. Panel B reports the corresponding relations with annualized alphas. The alpha-institution relations are positive and statistically imprecise. Taken together, the two panels reflect the asset-pricing pattern one would expect if institutional quality enters bond pricing through common-factor exposures rather than as an idiosyncratic premium. The unconditional return gap correlated with institutions is largely absorbed by issuer loadings on the common factors, leaving little residual variation in alphas.[8]

5.2 Institutions and sovereign-factor exposure

The central prediction of the simple asset-pricing model is that institutional quality is most strongly related to sovereign-factor exposure. Figure 1 provides visual evidence by comparing the baseline institutional composite with gold-standard adherence. The institutional composite has a clear negative relation with the estimated sovereign-factor loading. Gold-standard adherence also slopes downward, but the pattern is less distinct and does not appear to be the main source of cross-sectional variation in sovereign-factor loadings.[9]

Table 3 reports the corresponding cross-sectional regressions in the OLS sample.[10] The results are strongest for judicial constraints and the institutional composite. The OLS coefficient on the rescaled composite is -1.15 in the OLS sample of 47 issuers, with judicial constraints and the strict composite delivering similar estimates of -1.14 ($N = 47$) and -1.12 ($N = 45$). An interquartile-range move on the composite corresponds to a 0.47 lower sovereign-factor loading. Legislative constraints also enter negatively. Gold-standard adherence is negatively related to

[8] Because the sovereign factor and the alpha regressions are estimated from the same cross-section, the regressions are more precisely a test of whether institutions explain alpha over and above what is already absorbed by a within-sample factor (Lewellen et al., 2010). A weak alpha-institution relation is consistent with institutions affecting the covariance of sovereign returns with priced aggregate shocks rather than abnormal performance.

[9] Appendix Figure A.3 reports scatterplots for judicial constraints, legislative constraints, the institutional composite, and gold-standard adherence.

[10] The HC3 standard errors reported in Table 3 treat $\hat{\beta}_{i,\mathrm{SOV}}$ as a non-stochastic regressor in the cross-section. Lewellen et al. (2010) note that this can understate uncertainty when the loadings are themselves estimated. Appendix Table A.5 reports a complementary date-block bootstrap that propagates the first-stage estimation noise in the betas through to the estimated cross-sectional coefficient. The bootstrap standard errors are smaller than HC3 in every row, suggesting that the HC3 inference does not understate uncertainty for the cross-sectional coefficient.

sovereign-factor exposure on its own, but it does not displace the institutional measures when included alongside judicial constraints or the composite. The judicial constraints and composite specifications have the highest explanatory power among the single-variable regressions.

Appendix Table A.2 reports the same regressions side-by-side in the OLS sample and the IV sample for each institutional measure. Coefficients retain the same sign and similar magnitudes across samples, with the IV-sample magnitudes slightly larger.

The next question is whether institutions sort sovereign-specific exposure or broad factor exposure more generally. Figure 2 reports the estimated coefficient on the baseline institutional composite from separate regressions of each estimated factor loading on institutional quality. The estimated coefficient is largest in magnitude for the sovereign-factor loading, smaller for the London market loading, and close to zero for the term premium loading. Appendix Table A.7 confirms the pattern across institutional measures and gold-standard adherence: the institutional coefficient is concentrated in the sovereign-factor loading, with smaller magnitudes for the London market loading and essentially no relation for the term premium loading. Institutions were therefore not simply proxying for duration, nor were they only sorting exposure to the London market.

5.3 Robustness to factor construction

A potential concern with the regression evidence is that the sovereign bond market factor is constructed from the same bonds whose loadings are the dependent variable. If the institutional composite explains a substantial fraction of return variance, loadings on a value-weighted factor constructed from the same cross-section aligns, by construction, with the composite to some extent (Lewellen et al., 2010). The concern has a narrow form—each issuer's own bonds enter the factor used to price that issuer—and a broader form: the factor's variance is dominated by the same cross-section that is being tested. Table 4 addresses both.

Columns (2)–(3) address the narrow form by dropping each issuer's own bonds from the value-weighted aggregate at the all-issuer and foreign-sovereign-only levels. The estimated institutional coefficient is essentially unchanged at -1.05 ($t = -5.52$) and -0.97 ($t = -5.52$).[11]

Column (5) addresses the broader form. The factor is built from the upper-tercile composite-

[11] Column (4) replaces the bond factor with a leave-one-out version built within the colonial bond panel. The institutional coefficient there is small and imprecise (-0.72, $t = -1.78$). Colonial bonds traded as gilt-anchored credits with distinct fundamental drivers, making the colonial-bond factor a poor proxy for foreign-sovereign systematic risk.

institution issuers in the OLS sample,[12] with leave-one-out applied to those issuers, so the factor's construction is orthogonal to the bottom-two-thirds cross-section being tested. The estimated institutional coefficient remains negative and statistically significant at -1.14 ($t = -3.27$). Appendix Table A.8 reports the same regression with five additional factor constructions, all with negative institutional coefficients.

5.4 Pricing decomposition

The pricing implications also depend on whether those exposure differences are associated with positive factor premia. Table 5 decomposes the fitted raw-return gap between weaker- and stronger-institution issuers—the negative of the one-unit coefficient on the rescaled institutional composite (i.e., the gap between the cross-sectional extremes of the $[0, 1]$ index). The decomposition combines common full-sample factor premia (annualized as 13 times the mean four-week factor return) with the cross-sectional institutional gaps in estimated factor loadings.[13] For an interquartile-range move (0.41 units) on the rescaled institutional composite, the decomposition accounts for 1.33 percentage points of a 1.62 percentage-point raw-return gap in the OLS sample.[14] The sovereign factor contributes 1.21 percentage points, the London market factor 0.68 percentage point, and the term premium factor essentially zero. Alpha offsets about -0.56 percentage point against the factor contributions. The institutional return gap is mainly attributable to differences in exposure to priced common risk, with the sovereign factor as the largest component. The residual alpha gap is statistically imprecise (Table 2, Panel B), and its sign is not central to the institutional interpretation.[15]

[12] The sample includes Australia, Canada, Mauritius, New Zealand, Newfoundland, Sri Lanka, Trinidad, Uruguay, and the United States.

[13] Each of the three factors is itself an excess return on a tradable portfolio, so under no-arbitrage the price of risk is identified directly from the time-series mean (Cochrane, 2005, Ch. 6). The Newey-West t-statistic on the sovereign-factor mean is 2.82 ($p = 0.005$). The time-series mean is thus a more efficient estimator of the price of risk in this setting compared to a cross-sectional Fama-MacBeth approach with $N = 26$ test assets.

[14] Equivalently, between the cross-sectional extremes (a one-unit move on the rescaled composite), the decomposition accounts for 3.2 percentage points of a 3.9 percentage-point raw-return gap.

[15] Appendix Table A.6 reports bootstrap inference for the decomposition. The sovereign-factor contribution has a date-block-bootstrap 95% confidence interval of $[0.69,\ 5.49]$ percentage points and the London-market contribution $[0.12,\ 4.30]$, both excluding zero. The alpha gap has confidence interval $[-7.25,\ 3.15]$, confirming the imprecision noted in the text.

5.5 Gold-standard adherence as a comparison

Gold-standard adherence is related to sovereign-factor exposure but does not replicate the institutional pattern: the institutional composite exhibits a clearer negative relation with sovereign-factor exposure (Figure 1, Table 3), and gold adherence does not displace the institutional measures when both are included in a regression. Because gold adherence is itself an institutional outcome (Meissner, 2005), the contrast is best read not as a horse race but as evidence that broader commitment institutions, rather than a visible monetary regime alone, shaped systematic sovereign risk. Formal gold adherence was an incomplete proxy for the commitment institutions governing sovereign repayment incentives.

5.6 Instrumental-variables evidence

One concern is that institutional quality is correlated with other historical features that also affected sovereign risk, biasing the OLS estimates. Table 6 reports the instrumental-variables regressions using settler mortality as an instrument for institutional quality. Panel A reports the first-stage estimates, which are negative: higher settler mortality is associated with weaker institutional constraints. The first stage is strongest for judicial constraints and the two composite measures, but weaker for legislative constraints; the corresponding IV estimate should be interpreted with that caveat in mind. Because the IV estimates are identified from a small cross-section, I read them alongside first-stage diagnostics and reduced-form evidence rather than as stand-alone causal estimates.

Panel B reports the corresponding just-identified 2SLS estimates for sovereign-factor exposure. The instrumented coefficients are negative across all institutional measures. For the baseline composite, the IV estimate implies that stronger institutional quality is associated with substantially lower sovereign-factor exposure. The IV estimates are larger in magnitude than their OLS counterparts, particularly for legislative constraints, where the first-stage strength is weaker. This pattern fits either attenuation in OLS or amplification with weaker instruments, suggesting that the IV estimates are informative about the direction but less precise about the magnitude. The IV results complement the reduced-form pattern in Table 3.

One concern with the exclusion restriction is that settler mortality is correlated with malaria

ecology and the tropical disease environment, which could affect commodity composition, fiscal cyclicality, and therefore sovereign comovement through a non-institutional channel. Panel C of Table 6 adds the geography-and-disease-endowments index as an exogenous control in both stages of the 2SLS. The instrumented institutional coefficient remains negative and economically large across all four measures: -1.56 (judicial constraints), -1.68 (legislative constraints), -1.63 (composite), and -1.61 (strict composite). The estimated geography coefficient itself is small and statistically indistinguishable from zero in every column. First-stage instrument strength is weakened, reflecting that the geography index absorbs some of the variation in settler mortality, but it remains within the conventional range for the composite measures (robust F between 8 and 13). Anderson–Rubin p-values are 0.06–0.08, near or just above conventional significance and within an order of magnitude of the no-control benchmark. The geography control reduces precision but not the sign or magnitude.

Appendix Figure A.4 plots the corresponding first-stage relations without the geography control. The negative slopes are visible for judicial constraints, the composite, and executive constraints. Legislative constraints are also negatively related to settler mortality, but the relation is weaker. The figure illustrates why the IV evidence is most informative for judicial constraints and the composite measures.

5.7 Competing explanations

The final set of regressions investigates whether the relation between institutional quality and sovereign-factor exposure survives controls for competing historical explanations. Given the high correlation among historical controls, Table 7 adds one control at a time to the regression of the sovereign-factor loading on the baseline institutional composite.[16]

Adding gold-standard adherence, the geography-and-disease endowments index, or precolonial development leaves the negative institutional coefficient in place. It is attenuated and becomes statistically imprecise when either British legal origin or the British colony/empire indicator is included, reflecting the strong overlap among inherited legal background, British imperial status, and institutional constraints in this sample. These controls do not eliminate the main pat-

[16]The geography control should be interpreted as a reduced-form robustness check for the OLS relation, not as a validation of the settler-mortality exclusion restriction in the just-identified IV specification.

tern, but they identify where the evidence is most fragile.

Column (6) of Table 7 shows that the institutional coefficient in the IV sample attenuates by roughly half when a British colony/empire indicator is added. The question is whether this reflects collinearity between institutional measures and imperial status—in which case the institutional channel survives within each group—or substitution by a non-institutional imperial-guarantee channel.

Within-group regressions in Appendix Table A.3 disentangle these alternatives. Panel A reports the OLS sample, where statistical power is higher. Within the 36 independent sovereigns, the institutional slope operates at essentially the pooled magnitude: the composite slope is -1.01 ($t = -4.00$), 88 percent of the pooled -1.15, with judicial and legislative constraints showing similar magnitudes (-0.93 and -0.91, both $|t| > 3.4$). Within the 11 British colonies and imperial borrowers, the composite slope is small and imprecise (-0.16, $t = -1.18$), but the judicial-constraints slope is significant at the 5 percent level (-0.28, $t = -2.14$), indicating that the institutional channel operates within empire on the dimension that retains the most identifying variation.[17] Panel B reports the analogous IV-sample regressions. Within-group point estimates are smaller and uniformly imprecise at $N \in \{8, 10, 16\}$, consistent with the limited within-group identifying variation in the IV sample.

The within-group evidence indicates that the pooled institutional coefficient is not driven solely by the empire/non-empire gap. It operates within independent sovereigns at the pooled magnitude and within the empire subsample on the dimension where identifying variation is strongest. Gold-standard adherence behaves differently in the within-group regressions: its sign flips to weakly positive in both sub-samples, consistent with the pooled negative gold slope tracking the empire/non-empire split rather than within-regime variation.

Taken together, the cross-sectional, IV, and competing-explanation evidence in this section validates the model's prediction in Proposition 1.

[17] Imperial issuers cluster at the high-institution end of the composite (within-empire IQR 0.29 vs. 0.42 for independent sovereigns), which is why the within-empire judicial slope is more precise than the composite slope.

6 Interpretation: institutions and sovereign left-tail risk

6.1 The sovereign factor and aggregate sovereign credit stress

This subsection asks whether the sovereign bond market factor captures states of the world in which sovereign credit risk is elevated, rather than merely reflecting an omitted return component with no direct economic content.

This analysis is deliberately descriptive and not intended to establish that factor returns cause defaults. The aggregate default series is annual (Reinhart and Rogoff, 2009), while the hand-collected bond returns are sampled at four-week intervals. As a result, I treat the annual default series as the binding frequency and use the higher-frequency return data to construct annual measures of factor-level left-tail risk.

The main figure uses an annual left-tail count of sovereign-factor returns. Let $r_{y,m}^{SOV}$ denote the four-week return on the sovereign bond market factor in year y and period $m = 1, \ldots, M_y$, and let $q_{.10}^{SOV}$ denote the unconditional 10th percentile of four-week sovereign-factor returns. The left-tail count is

$$\mathrm{LT10}_y^{\mathrm{SOV}} = \sum_{m=1}^{M_y} \mathbf{1}\{r_{y,m}^{SOV} < q_{.10}^{SOV}\}.$$

The complementary measures are the worst four-week sovereign-factor return in the year, cumulative downside returns, and within-year and rolling 13-period maximum drawdowns.[18] Figure 3 plots this count against the annual share of sovereigns in default. Years with elevated default shares tend to coincide with more frequent poor sovereign-factor returns, especially during the 1870s default wave (Eichengreen et al., 2021, Ch. 6). The relation is episodic rather than steady year-to-year: the left tail of the sovereign bond market factor is concentrated during periods of aggregate sovereign credit stress.

Table 8 compares this relation across the sovereign bond market, London market, and British term premium factors. The annual default share is most strongly related to the downside measures of the sovereign bond market factor. The correlations are 0.41 and 0.40 for the 10th- and 5th-percentile left-tail counts, 0.37 for cumulative downside returns, 0.34 for the rolling 13-period

[18] The cumulative return index is $W_{y,m} = \prod_{j=1}^{m}(1 + r_{y,j}^{SOV})$ with $W_{y,0} = 1$, and the annual maximum drawdown is $MDD_y = \max_{1 \le m \le M_y}\left(1 - W_{y,m} / \max_{0 \le k \le m} W_{y,k}\right)$, the largest within-year peak-to-trough loss expressed as a positive number.

maximum drawdown, and 0.27 for the within-year maximum drawdown. The corresponding correlations for the London market and British term premium factors are smaller, mixed in sign, or concentrated in measures that lack the same sovereign-specific interpretation. The correlation with the annual sovereign-factor return itself is only 0.15. The default share tracks the left tail of the sovereign-factor distribution, not the average annual factor return.

Appendix Table A.9 reports a compact set of timing and episode checks. The lagged correlations are similar or somewhat larger than the contemporaneous correlations, which is consistent with the Reinhart–Rogoff series measuring the share of countries in default rather than the onset of new defaults. At the same time, the relation weakens substantially when the 1875–1878 default wave is removed.

Taken together, this evidence is consistent with the sovereign bond market factor capturing periods of aggregate sovereign credit stress, but it does not establish a stable annual mapping from default shares to factor drawdowns.

6.2 Two channels in tail states

The model also makes a second tail-state prediction. Proposition 2 states that stronger institutions should raise the tail-state intercept $\alpha_{i,\text{tail}}$ via the tail-specific component L_i^T. Under the cross-partial restriction, institutional effects should sharpen in left-tail states for both channels: steepening $\hat{\beta}_{i,\text{SOV}}$ via L_i^F and concentrating $\hat{\alpha}_{i,\text{tail}}$ in deeper tail realizations via L_i^T. This subsection nests both channels in a single panel regression and reports their relative magnitudes.

The augmented panel regression is

$$r^e_{i,t} = \mu_i + \beta\, f_{\text{SOV},t} + \beta_{\text{tail}}\, f_{\text{SOV},t}\, D_t + \gamma\, f_{\text{SOV},t}\, I_i + \gamma_{\text{tail}}\, f_{\text{SOV},t}\, I_i\, D_t + \delta\, D_t + \zeta\, I_i\, D_t + u_{i,t},$$

with issuer fixed effects μ_i, $D_t = \mathbf{1}\{f_{\text{SOV},t} < q_{0.10}(f_{\text{SOV}})\}$, and I_i the rescaled baseline institutional composite. Standard errors are two-way clustered by issuer and by date (Cameron et al., 2011). The slope coefficients γ and γ_{tail} identify how institutions sort $\hat{\beta}_{i,\text{SOV}}$ unconditionally and in tail states, with the cross-partial restriction on L_i^F predicting $\gamma_{\text{tail}} \leq 0$. The level coefficients δ and ζ identify the tail-state intercept and its institutional slope, with Proposition 2 predicting $\zeta > 0$.

Table 9 reports the result. The first two rows estimate the augmented specification in the OLS

sample (47 issuers) and the IV sample (26 issuers); the third row implements the cleanest-test variant discussed below. The unconditional slope channel survives strongly in both samples: $\gamma = -1.92$ ($t = -5.80$) in the OLS sample and $\gamma = -2.24$ ($t = -4.74$) in the IV sample, consistent with Table 3 and confirming that the cross-sectional relation between institutions and $\hat{\beta}_{i,\mathrm{SOV}}$ is negative on average. The triple-interaction γ_{tail} carries the opposite sign from the cross-partial prediction on L_i^F in both samples: $+1.00$ ($t = +1.99$) in the OLS sample and $+1.90$ ($t = +2.36$) in the IV sample. The institutional gap in $\hat{\beta}_{i,\mathrm{SOV}}$ compresses in left-tail states rather than steepening, rejecting the cross-partial restriction on L_i^F in both samples.

The cleanest test of the level channel addresses the concern that f_{SOV} is value-weighted across the same cross-section being tested. I replace f_{SOV} with the high-institution-subsample factor of Table 4 and restrict attention to issuers excluded from the upper-tercile composite, a sample for which the factor's construction is exogenous. This sample restriction yields $\delta = -0.081$ ($t = -6.68$) and $\zeta = +0.181$ ($t = 5.02$) in the third row of Table 9: a low-institution issuer loses 8.1 percentage points per tail period at the level, while a high-institution issuer's tail-state level return is $\delta + \zeta = +0.100$. The IV-sample analog with the original f_{SOV} (second row) delivers the same pattern at smaller magnitude ($\delta = -0.029$, $\zeta = +0.037$; high-institution tail-state returns essentially flat at $+0.008$).

In the OLS sample with the original f_{SOV} (first row of Table 9), the level magnitudes attenuate by more than half ($\delta = -0.012$, $\zeta = +0.012$) and are not statistically distinguishable from zero. The attenuation reflects the well-documented core/peripheral distinction in 1870–1907: cross-sectional dispersion in default risk was concentrated in Latin America, the Mediterranean, and the Ottoman Empire, while core European sovereigns and British colonies paid spreads close to British consols throughout the period (Bordo and Rockoff, 1996; Flandreau and Zumer, 2004; Mauro et al., 2006; Ferguson and Schularick, 2006). Of the 21 OLS-only issuers, 15 had no external debt restructuring with default in 1870–1907; the remaining 6—Santo Domingo, Greece, Liberia, Portugal, Spain, and Turkey—experienced 10 restructurings in the window with a mean present-value haircut of 58% (Meyer et al., 2022). Within the OLS-only defaulter subgroup the mean institutional composite (0.37) is essentially identical to the IV defaulter group's (0.36), so the attenuation does not reflect a shift in within-defaulter institutional composition. It reflects the addition of 15 zero-loss observations, leaving the level channel empirically identified from the issuers in which perceived default

risk was concentrated (Meissner, 2005).[19]

The cross-partial restriction on L_i^F is rejected in both samples; stronger institutions do not steepen the sovereign-factor loading in tail states. The cross-partial restriction on L_i^T is supported where the level channel has identifying power: stronger institutions reduce tail-state level losses, consistent with Proposition 2.[20] One caveat is that the test conditions on left-tail realizations of f_{SOV} rather than directly on the primitive shock z, and the two align only episodically, mostly in the 1875–1878 default wave. With that qualification, institutional protection in tail states operates through the tail-specific loss component—default-related impairment, recovery shortfalls, and similar discontinuous losses—rather than through state-dependent comovement of factor-driven losses with the sovereign factor.

7 Conclusion

The evidence from late-nineteenth-century London shows that stronger institutions are associated with substantially lower exposure to the common sovereign risk factor, but not with lower abnormal returns once common risks are accounted for. Institutions mattered by affecting the covariance between sovereign bond returns and priced aggregate risks, not by producing alpha. The sovereign factor is best viewed as a traded common component of sovereign bond returns whose left-tail realizations are concentrated in periods of aggregate sovereign credit stress, and issuers with weaker institutional constraints were more exposed to those periods. The comparison with gold-standard adherence refines the interpretation: gold was an observable regime and a plausible commitment device, but it does not display the same concentrated relation with sovereign-factor exposure. The evidence points to deeper constitutional and legal constraints, rather than formal regime adherence alone, as the institutional features most relevant for pricing.

These findings have three broader implications. First, although the analysis is not a formal

[19] Appendix Table A.10 lists the 10 OLS-only restructurings from the Meyer et al. (2022) archive with default dates in the sample window. The 15 OLS-only issuers without an in-window event split between 7 sovereigns absent from the haircut archive (Belgium, Denmark, Fiji, France, Netherlands, Norway, Sweden) and 8 whose first default falls after 1907 (Austria-Hungary, Bulgaria, China, Germany, Italy, Japan, Romania, Russia).

[20] Translating the IV-sample coefficients into expected-return contributions per year, the loading channel accounts for roughly $+12$ percentage points of the average expected-return gap between weak- and strong-institution issuers, while the tail-state level channel accounts for about -5 percentage points. The two channels pull in opposite directions in tail states. In the OLS sample with the original f_{SOV}, the level channel attenuates, and the loading channel carries the gap on its own.

out-of-sample test of modern factor structure, the evidence is consistent with the proposition that sovereign credit risk has a systematic component extending well beyond the modern period of credit ratings, multilateral institutions, and contemporary monetary regimes. Second, average returns and risk-adjusted alphas, which dominate existing work on sovereign credibility, are likely to miss the channel through which institutions enter sovereign pricing. Third, the comparison with gold suggests that visible policy regimes are an incomplete proxy for the commitment institutions that govern sovereign repayment incentives in adverse states.

Commitment institutions thus help explain cross-country heterogeneity in exposure to a traded common component of sovereign bond returns whose left-tail realizations are concentrated in periods of aggregate sovereign credit stress.

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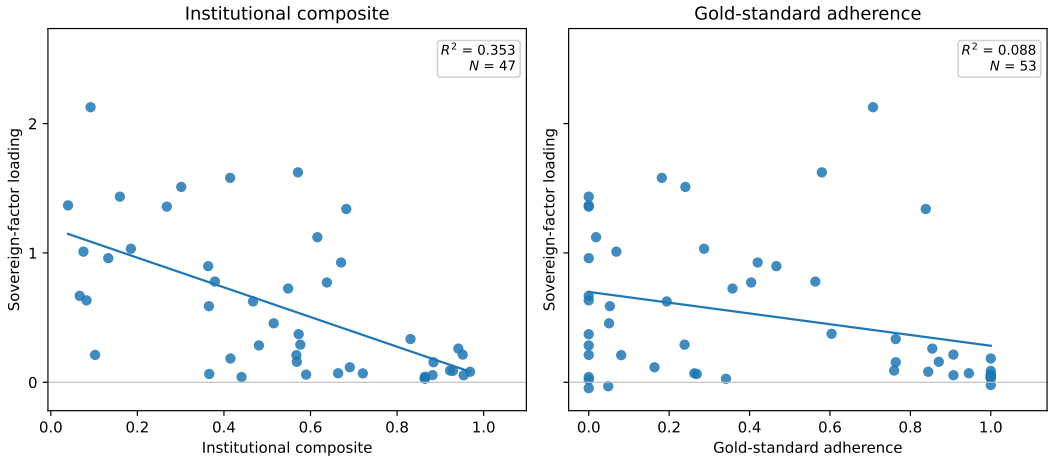
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8 Figures and tables

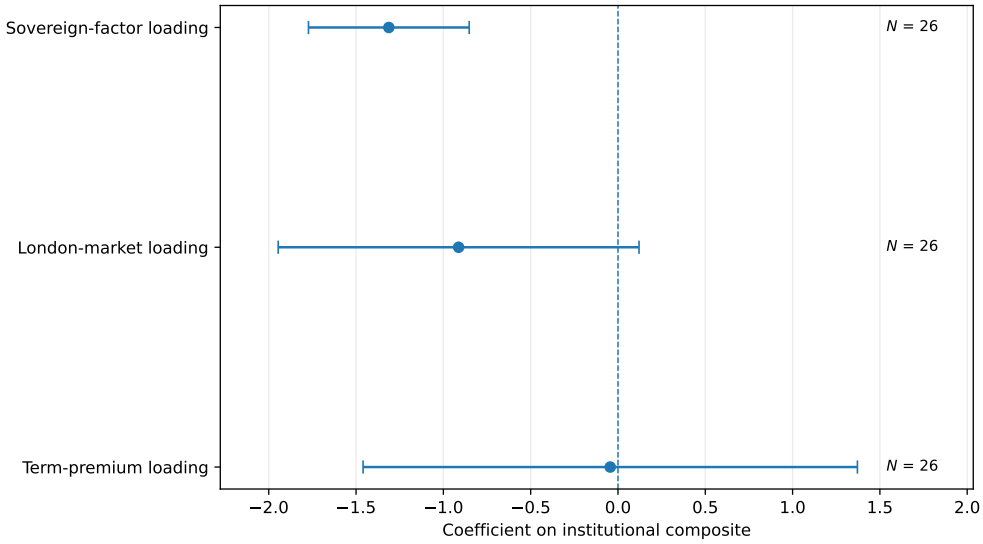
8.1 Figures

Figure 1: Institutions, gold-standard adherence, and sovereign-factor exposure



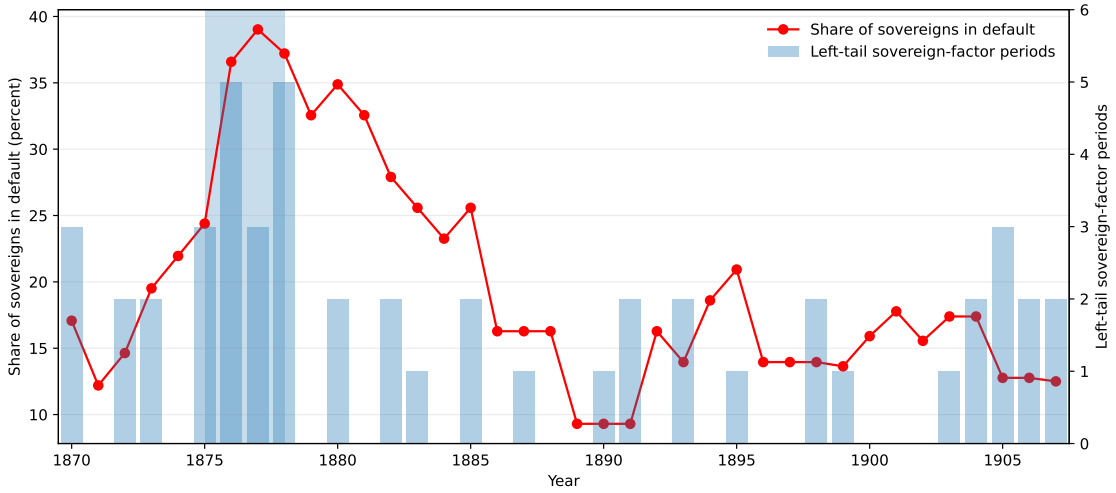
Notes: The figure plots the estimated sovereign-factor loading against the rescaled baseline institutional composite (left panel, $N = 47$) and gold-standard adherence (right panel, $N = 53$) for issuers in the OLS sample. Each panel uses the maximal cross-section with non-missing data on the corresponding x-variable. The institutional composite is constructed from V-Dem measures of judicial and legislative constraints on the executive, aggregated to country-level medians over 1870–1907 and then min–max rescaled across issuers to the unit interval, so that the issuer with the weakest constraints in the sample takes value 0 and the issuer with the strongest takes value 1. Gold-standard adherence is the share of four-week sample periods in which the issuer was on gold. The dashed line is the OLS fit.

Figure 2: Institutional quality and factor loadings



Notes: The figure reports the estimated coefficient on the rescaled baseline institutional composite from separate cross-sectional regressions of each estimated factor loading on the institutional composite, in the IV sample of 26 issuers. The composite is the country-level V-Dem median over 1870–1907, min–max rescaled across issuers to the unit interval. Coefficients therefore measure the difference between hypothetical issuers at the cross-sectional extremes, and an interquartile-range move (0.40 units in the IV sample) scales each estimate by 0.40. Horizontal bars are 95 percent confidence intervals based on heteroskedasticity-robust standard errors. The corresponding regressions appear in Appendix Table A.7.

Figure 3: Aggregate sovereign defaults and sovereign-factor left-tail risk



Notes: The figure compares the annual share of sovereigns in external default from Reinhart and Rogoff (2009) (red line, left axis) with the number of four-week sovereign-factor returns in year y that fall below the unconditional 10th percentile of the four-week sovereign-factor return distribution (blue bars, right axis). Each year contains 13 four-week periods, so the right-axis count could, in principle, reach 13. The maximum realized in the sample is six, observed during the 1875–1878 default wave (shaded region) associated with the suspensions of payment by Egypt, Turkey, Peru, Uruguay, Costa Rica, Honduras, Liberia, Paraguay, Bolivia, and Spain. The series covers 1870–1907.

8.2 Tables

Table 1: Summary statistics

Panel A. Full sample	mean	sd	p25	median	p75	*N*
Average excess return	3.246	2.904	1.420	2.912	4.791	53
Annualized alpha	1.372	3.547	0.577	1.484	2.793	53
Sovereign-factor loading	0.518	0.549	0.069	0.285	0.898	53
First-stage SE of $\hat{\beta}_{i,\mathrm{SOV}}$	0.180	0.156	0.060	0.120	0.277	53
London-market loading	0.299	0.544	0.034	0.086	0.329	53
Term-premium loading	0.063	0.425	-0.042	0.060	0.145	53
Gold-standard adherence	0.435	0.391	0.048	0.341	0.844	53
Judicial constraints	0.579	0.301	0.389	0.590	0.873	47
Legislative constraints	0.482	0.314	0.205	0.499	0.750	45
Institutional composite	0.537	0.287	0.364	0.568	0.776	47
Institutional composite, strict	0.526	0.287	0.364	0.567	0.691	45
Executive constraints	0.494	0.382	0.167	0.333	1.000	41
Settler mortality	3.746	0.845	2.874	4.217	4.535	29
Panel B. IV sample	mean	sd	p25	median	p75	*N*
Average excess return	3.140	2.832	1.131	2.346	5.237	26
Annualized alpha	0.684	3.566	0.473	1.437	1.877	26
Sovereign-factor loading	0.630	0.597	0.074	0.480	1.257	26
First-stage SE of $\hat{\beta}_{i,\mathrm{SOV}}$	0.218	0.170	0.061	0.205	0.336	26
London-market loading	0.438	0.606	0.033	0.181	0.737	26
Term-premium loading	0.051	0.576	-0.080	0.046	0.130	26
Gold-standard adherence	0.392	0.408	0.013	0.217	0.819	26
Judicial constraints	0.576	0.311	0.376	0.622	0.877	26
Legislative constraints	0.438	0.313	0.132	0.438	0.706	24
Institutional composite	0.521	0.285	0.317	0.569	0.713	26
Institutional composite, strict	0.498	0.284	0.293	0.517	0.685	24
Executive constraints	0.465	0.383	0.167	0.333	0.833	19
Settler mortality	3.795	0.844	3.418	4.231	4.535	26

Notes: This table reports cross-sectional summary statistics for the issuer-level variables used in the paper. Panel A reports the full issuer sample available in the current dataset. Panel B reports the IV sample used in the cross-sectional regressions. Average excess returns and alphas are annualized and reported in percentage points. The three factor-loading variables are estimated from time-series regressions of issuer excess returns on the sovereign bond market factor, the London market factor, and the British term premium factor. Gold-standard adherence is the share of four-week sample periods in which the issuer was on gold. The institutional variables are cross-sectionally min-max rescaled country-level medians over 1870–1907 of V-Dem indices (judicial constraints, legislative constraints, and the composite) and Polity executive constraints, with the weakest-constraint issuer at 0 and the strongest at 1; see Section 4 for details. N is the number of non-missing observations for each variable. The sample period is 1870–1907.

Table 2: Benchmark regressions: average excess returns and alphas

Panel A. Average excess return	(1)	(2)	(3)	(4)	(5)
Judicial constraints	-4.414***				
	(1.179)				
Legislative constraints		-2.308**			
		(1.113)			
Institutional composite			-3.931***		
			(1.061)		
Institutional composite, strict				-3.791***	
				(1.085)	
Gold-standard adherence					-3.082***
					(0.931)
adj.-R^2	0.173	0.033	0.118	0.105	0.156
N	47	45	47	45	53
Panel B. Annualized alpha	(1)	(2)	(3)	(4)	(5)
Judicial constraints	0.948				
	(1.817)				
Legislative constraints		1.446			
		(1.431)			
Institutional composite			1.350		
			(1.758)		
Institutional composite, strict				1.409	
				(1.779)	
Gold-standard adherence					-0.677
					(1.257)
adj.-R^2	-0.016	-0.009	-0.011	-0.012	-0.014
N	47	45	47	45	53

Notes: This table reports benchmark cross-sectional regressions in the OLS sample (OLS sample available for each institutional measure). Panel A uses the average annualized excess return as the dependent variable. Panel B uses the annualized three-factor alpha. Each column reports a separate regression of the indicated dependent variable on one institutional measure or on gold-standard adherence. Columns (1)–(5) correspond to judicial constraints, legislative constraints, the baseline institutional composite, the strict institutional composite, and gold-standard adherence, respectively. Judicial constraints, legislative constraints, and the composite institutional measures enter as country-level medians for 1870–1907, then min–max rescaled across issuers to the unit interval, so the issuer with the weakest constraints in the sample takes value 0 and the strongest takes value 1. The reported OLS coefficients therefore measure the difference between the two cross-sectional extremes; an interquartile-range (0.5-unit) move halves each estimate. Heteroskedasticity-robust standard errors are reported in parentheses, using the HC3 leverage adjustment of MacKinnon and White (1985). $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

Table 3: Institutions, gold-standard adherence, and sovereign-factor exposure

Sovereign-factor loading	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Judicial constraints	-1.139***					-1.306***		
	(0.238)					(0.383)		
Legislative constraints		-0.826***					-0.720*	
		(0.247)					(0.411)	
Institutional composite			-1.147***					-1.272***
			(0.222)					(0.424)
Institutional composite, strict				-1.115***				
				(0.225)				
Gold-standard adherence					-0.416**	0.194	-0.152	0.139
					(0.164)	(0.275)	(0.315)	(0.321)
adj.-R^2	0.371	0.200	0.338	0.319	0.070	0.367	0.189	0.329
N	47	45	47	45	53	47	45	47

Notes: This table reports cross-sectional regressions of the estimated sovereign-factor loading, $\hat{\beta}_{i,\mathrm{SOV}}$, on institutional quality and gold-standard adherence in the OLS sample (OLS sample available for each institutional measure). Columns (1)–(4) regress $\hat{\beta}_{i,\mathrm{SOV}}$ on one institutional measure at a time. Column (5) regresses $\hat{\beta}_{i,\mathrm{SOV}}$ on gold-standard adherence alone. Columns (6)–(8) include both gold-standard adherence and the indicated institutional measure. Judicial constraints, legislative constraints, and the composite institutional measures enter as country-level medians for 1870–1907, then min–max rescaled across issuers to the unit interval, so the issuer with the weakest constraints in the sample takes value 0 and the strongest takes value 1. The reported OLS coefficients therefore measure the difference between the two cross-sectional extremes; the cross-sectional IQR of the institutional composite in the OLS sample is 0.41 units. Heteroskedasticity-robust standard errors are reported in parentheses, using the HC3 leverage adjustment of MacKinnon and White (1985). $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

Table 4: Sovereign-factor exposure under alternative factor constructions

Sovereign-factor loading	Baseline VW	LOO sovereign	LOO foreign	LOO colonial	HI subsample
Institutional composite	-1.147***	-1.045***	-0.967***	-0.718*	-1.135***
	(0.222)	(0.189)	(0.175)	(0.403)	(0.347)
adj.-R^2	0.338	0.367	0.366	0.097	0.235
N	47	47	47	47	47

Notes: This table reports cross-sectional regressions of issuer-level loadings on the column-specific bond-market factor on the baseline institutional composite in the OLS sample of 47 issuers with non-missing baseline institutional composite. Each loading is estimated from a three-factor time-series regression that also includes the London market factor and the British term premium factor. The five columns vary the construction of the bond-market factor: Baseline VW uses the value-weighted sovereign-and-colonial bond factor (the model's main specification); LOO sovereign drops each issuer's own bonds from the factor used to price that issuer; LOO foreign restricts the factor to the foreign-sovereign bond panel and applies leave-one-out within that panel (the full foreign-only factor is used for colonial issuers); LOO colonial does the analogous exercise on the colonial bond panel; HI subsample uses a factor built from the upper-tercile composite-institution issuers (Australia, Canada, Mauritius, New Zealand, Newfoundland, Sri Lanka, Trinidad, Uruguay, USA), with leave-one-out applied to those issuers. The HI-subsample construction is the most stringent reflexivity test: the factor's construction does not depend on the bottom two-thirds of the institutional cross-section being tested. The institutional composite enters as a country-level median for 1870–1907, then min–max rescaled across issuers to the unit interval. Heteroskedasticity-robust standard errors are reported in parentheses, using the HC3 leverage adjustment of MacKinnon and White (1985). Coefficient magnitudes are not directly comparable across columns because the alternative bond-market factors have different return scales. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

Table 5: Decomposition of the institutional raw-return gap

	London market	Term premium	Sovereign	Alpha	Fitted total	Raw return gap
Contribution (pp)	1.64	0.00	2.94	-1.35	3.23	3.93

Notes: This table decomposes the institutional raw-return gap into factor-exposure and alpha components in the OLS sample of 47 issuers with non-missing baseline institutional composite. The raw-return gap is the fitted low-minus-high difference from a one-unit increase in the rescaled institutional composite, equivalent to comparing the issuer with the weakest constraints in the sample to the issuer with the strongest. An interquartile-range move (0.41 units in the OLS sample) scales each entry by 0.41. Factor premia are common full-sample annualized factor means, computed as 13 times the four-week factor mean. Exposure gaps are the corresponding low-minus-high differences in estimated factor loadings. The contribution of factor k is $\hat{\lambda}_k \Delta\hat{\beta}_k$. The fitted total need not equal the raw-return gap exactly because the common-premium decomposition uses full-sample factor premia rather than issuer-specific factor means.

Table 6: Settler mortality and institutional quality

Panel A. First stage	(1)	(2)	(3)	(4)
Settler mortality	-0.233***	-0.189***	-0.204***	-0.219***
	(0.052)	(0.069)	(0.052)	(0.051)
Partial R^2	0.400	0.273	0.364	0.446
Robust first-stage F	20.265	7.436	15.532	18.563
MOP effective F	22.050	8.472	17.087	20.704
N	26	24	26	24
Panel B. Second stage	(1)	(2)	(3)	(4)
Instrumented institution	-1.598***	-2.093***	-1.832***	-1.808***
	(0.343)	(0.801)	(0.473)	(0.454)
AR p-value	0.000	0.000	0.000	0.000
AR 95% conf. set	[-2.444, -0.899]	[-7.065, -0.931]	[-3.262, -0.953]	[-3.105, -0.961]
2SLS/OLS ratio	1.224	2.714	1.396	1.450
N	26	24	26	24
Panel C. Second stage with geography-and-disease-endowments control	(1)	(2)	(3)	(4)
Instrumented institution	-1.555***	-1.683*	-1.632**	-1.606**
	(0.530)	(0.948)	(0.659)	(0.698)
Geography/endowments	0.025	0.068	0.053	0.045
	(0.226)	(0.488)	(0.318)	(0.344)
Robust first-stage F	8.100	10.548	11.071	13.037
AR p-value	0.057	0.082	0.057	0.082
N	24	22	24	22

Notes: This table reports instrumental-variables evidence using updated log settler mortality as an instrument for institutional quality in the IV sample. Columns (1)–(4) correspond to judicial constraints, legislative constraints, the baseline institutional composite, and the strict institutional composite, respectively. Panel A reports first-stage regressions of each institutional measure on settler mortality. The partial R^2, robust first-stage F, and MOP effective F summarize the excluded instrument. Panel B reports just-identified 2SLS regressions of the estimated sovereign-factor loading, $\hat{\beta}_{i,\mathrm{SOV}}$, on the corresponding instrumented institutional measure. Panel C repeats the second stage with the geography-and-disease-endowments index added as an exogenous control, with the geography index also included in the first stage; this is the specification that most directly tests whether the IV reduced form operates through a non-institutional channel. The geography index is missing for two issuers (Newfoundland and Hong Kong), reducing N for the corresponding columns. AR denotes the Anderson–Rubin weak-instrument-robust test and confidence set. The 2SLS/OLS ratio compares the IV coefficient with the corresponding OLS coefficient estimated using the same outcome and institutional measure. Judicial constraints, legislative constraints, and the composite institutional measures enter as country-level medians for 1870–1907, then min–max rescaled across issuers to the unit interval, so the issuer with the weakest constraints in the sample takes value 0 and the strongest takes value 1. The reported OLS coefficients therefore measure the difference between the two cross-sectional extremes; an interquartile-range (0.5-unit) move halves each estimate. Heteroskedasticity-robust standard errors are reported in parentheses, using the HC3 leverage adjustment of MacKinnon and White (1985). $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

Table 7: Sovereign-factor exposure and competing explanations

Sovereign-factor loading	(1)	(2)	(3)	(4)	(5)	(6)
Institutional composite	-1.312***	-1.394***	-1.223***	-0.519	-0.982**	-0.593
	(0.235)	(0.478)	(0.385)	(0.412)	(0.458)	(0.423)
Gold-standard adherence		0.081				
		(0.412)				
Geography/endowments			-0.077			
			(0.256)			
British legal origin				-0.716***		
				(0.242)		
Precolonial development					0.143	
					(0.188)	
British colony/empire						-0.710***
						(0.232)
adj.-R^2	0.366	0.341	0.346	0.545	0.347	0.589
N	26	26	24	25	21	26

Notes: This table reports cross-sectional regressions of the estimated sovereign-factor loading, $\hat{\beta}_{i,\mathrm{SOV}}$, on the baseline institutional composite in the IV sample, adding one competing explanation at a time. Column (1) includes the institutional composite alone. Columns (2)–(6) add gold-standard adherence, the geography-and-disease endowments index, British legal origin, precolonial development, and the British colony/empire indicator, respectively. The institutional composite enters as a country-level median for 1870–1907, then min–max rescaled across issuers to the unit interval; coefficients therefore measure the gap between the two cross-sectional extremes, and an interquartile-range (0.5-unit) move halves each estimate. Sample size varies across columns because some controls are missing for subsets of issuers. Heteroskedasticity-robust standard errors are reported in parentheses, using the HC3 leverage adjustment of MacKinnon and White (1985). $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

Table 8: Aggregate sovereign defaults and sovereign-factor left-tail risk

	SOV	MKT	TERM
Maximum drawdown	0.274	-0.022	-0.200
Worst four-week loss	0.325	0.054	0.246
Cumulative downside return	0.370	-0.066	-0.000
Left-tail periods: 10th percentile	0.406	0.099	0.125
Left-tail periods: 5th percentile	0.404	0.053	0.176
13-period rolling maximum drawdown	0.342	0.048	-0.351
Annual factor return	0.146	0.078	0.264

Notes: This table reports correlations between the annual share of sovereigns in default and annual factor-level left-tail risk measures from 1870 to 1907. The default-share series is from Reinhart and Rogoff. SOV is the sovereign bond market factor, MKT is the London market factor, and TERM is the British term premium factor. Maximum drawdown is the largest within-year peak-to-trough loss in the factor return index. Worst four-week loss is the negative of the minimum four-week factor return in the year. Cumulative downside return is the sum of negative four-week factor returns, expressed as a positive value. Left-tail-period counts measure the number of four-week factor returns below the unconditional 10th or 5th percentile of the corresponding factor-return distribution. The 13-period rolling maximum drawdown is computed over a trailing 13-period window and assigned to the calendar year by taking the maximum rolling drawdown observed within that year. Annual factor return is the compounded annual factor return.

Table 9: State-contingent sovereign-factor exposure

Sample	f_{SOV}	$f_{\mathrm{SOV}} \cdot D$	$f_{\mathrm{SOV}} \cdot I$	$f_{\mathrm{SOV}} \cdot I \cdot D$	D	$I \cdot D$	N obs	N issuers	N dates
OLS sample	1.960***	-0.862**	-1.922***	1.001**	-0.012	0.012	15,701	47	495
	(0.254)	(0.410)	(0.331)	(0.502)	(0.011)	(0.013)			
IV sample	2.273***	-1.531***	-2.235***	1.897**	-0.029**	0.037**	8,839	26	495
	(0.362)	(0.589)	(0.472)	(0.803)	(0.014)	(0.018)			
HI factor, non-HI issuers	3.081***	-5.273***	-3.140**	11.896***	-0.081***	0.181***	5,647	17	484
	(0.843)	(1.056)	(1.598)	(1.997)	(0.012)	(0.036)			

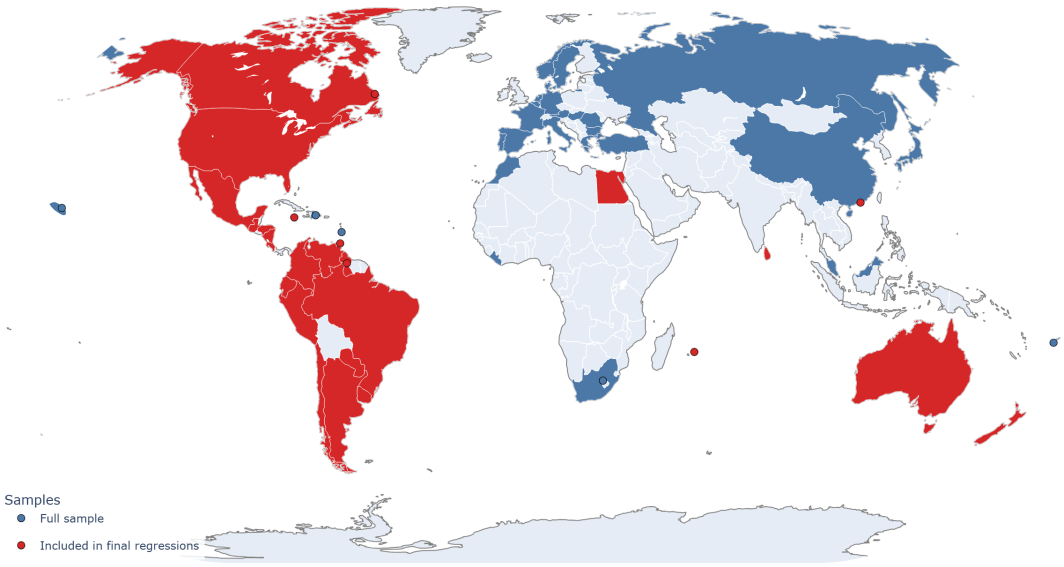
Notes: Panel regressions of issuer four-week excess returns on the value-weighted sovereign-bond factor f_{SOV} and its interactions with the institutional composite I_i and a left-tail-state indicator D_t, with issuer fixed effects. All three specifications include D_t and $I_i \cdot D_t$ as level main effects so that ζ identifies the institutional gradient on the tail-state intercept rather than absorbing a tail-state level shift. The first row uses the OLS sample of issuers with non-missing institutional composite and panel returns. The second row uses the 26-issuer IV sample with the same value-weighted f_{SOV}. The third row replaces f_{SOV} with the high-institution-subsample sovereign factor (built from the upper-tercile composite-institution issuers in the IV sample, with leave-one-out applied to those issuers) and restricts attention to the 17 non-HI issuers, a sample for which the factor's construction is exogenous to the cross-section being tested. D_t equals one when f_{SOV} falls below its tenth-percentile realization in the 1870–1907 sample. I_i is the rescaled baseline institutional composite, the same measure used in Table 3. Standard errors, in parentheses, are clustered by both issuer and date (Cameron et al., 2011). The triple-interaction $f \cdot I \cdot D$ tests the cross-partial restriction applied to the factor-driven loss component L^F (model predicts non-positive); the main effect $I \cdot D$ tests the level condition on the tail-specific component L^T (model predicts non-negative). $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

A Appendices

A.1 Foreign sovereign and colonial bond data

Figure A.1 maps the geographic distribution of issuers in the sample. The dataset comprises 213 foreign sovereign bonds issued by 39 countries and 110 colonial bonds issued by 14 British colonies, possessions, and members of the British Empire.

Figure A.1: Geographic distribution of issuers



Notes: Blue markers indicate issuers in the full sample. Red markers indicate the 26 issuers in the final cross-sectional regressions. National boundaries shown are modern. Issuer locations are mapped to the modern country area corresponding to each historical issuer; sovereigns and territories whose borders or status differed in 1870–1907 (e.g., the Austro-Hungarian and Ottoman Empires) are shown at the closest modern equivalent.

The full sample includes the following issuers: Argentina, Australia, Austria-Hungary, Belgium, Brazil, British Guiana, Bulgaria, Canada, Ceylon, Chile, China, Colombia, Costa Rica, Denmark, Ecuador, Egypt, Fiji, France, Germany, Greece, Guatemala, Hawaii, Honduras, Hong Kong, Italy, Jamaica, Japan, Liberia, Mauritius, Mexico, Morocco, Netherlands, New Zealand, Newfoundland, Nicaragua, Norway, Orange Free State, the Ottoman Empire, Paraguay, Peru, Portugal, Romania, Russia, Saint Lucia, Santo Domingo, South Africa, Spain, the Straits Settlements, Sweden, Trinidad, the United States, Uruguay, and Venezuela.

The 14-issuer British imperial subsample consists of Australia, British Guiana, Canada, Ceylon,

Fiji, Hong Kong, Jamaica, Mauritius, New Zealand, Newfoundland, Saint Lucia, South Africa, the Straits Settlements, and Trinidad.

The IV sample of 26 issuers used in the cross-sectional regressions comprises Argentina, Australia, Brazil, British Guiana, Canada, Ceylon, Chile, Colombia, Costa Rica, Ecuador, Egypt, Guatemala, Honduras, Hong Kong, Jamaica, Mauritius, Mexico, New Zealand, Newfoundland, Nicaragua, Paraguay, Peru, Trinidad, the United States, Uruguay, and Venezuela.

A.2 Sovereign borrowing costs and gold-standard adherence

Appendix Table A.1 shows that gold-standard adherence was not associated with risk-adjusted returns in this setting. Although off-gold bonds earned higher raw returns than on-gold bonds, that spread disappears once returns are projected on common London risk factors. The off-minus-on portfolio does not earn a statistically significant alpha under either weighting scheme. This result is substantively consistent with earlier return-based evidence, even though the estimates are not numerically identical because this sample includes additional countries, revised bond classifications, and corrected gold-adherence dates (Alquist and Chabot, 2011). This alpha result serves as a useful benchmark for analyzing factor exposures, where deeper institutional measures are more informative than gold adherence alone.

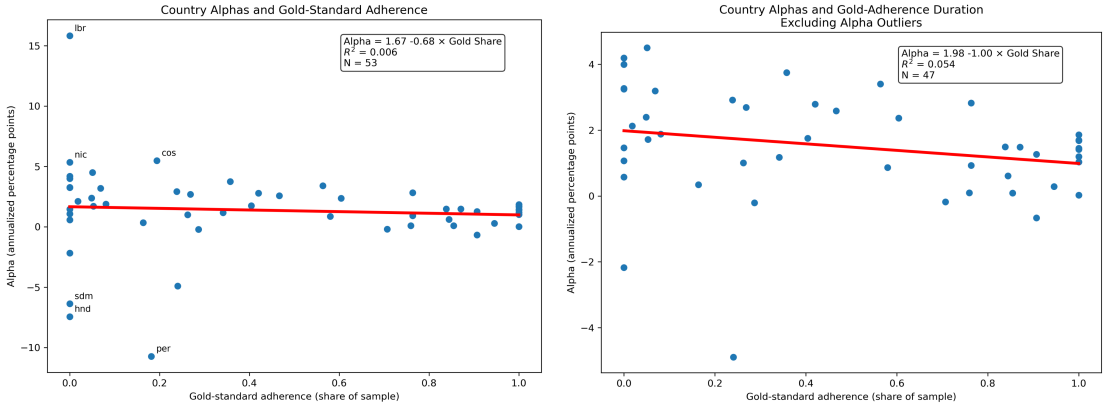
Appendix Figure A.2 plots each country's annualized three-factor alpha against the share of sample periods during which the country adhered to gold. In the full sample (left panel), the fitted relation is negative but weak, with little explanatory power for the cross-country variation in alphas. After excluding the three highest- and three lowest-alpha observations (right panel), the slope becomes somewhat steeper, but the relation remains modest and continues to explain little of the variation in returns.

Table A.1: Excess return regressions: off- and on-gold portfolios

	Value-weighted			Equally weighted		
	Off-gold	On-gold	Off–On	Off-gold	On-gold	Off–On
Mean excess return	4.232	2.488	1.743	4.378	2.489	1.888
α	1.312	0.795	0.517	1.170	1.246***	-0.075
	(1.564)	(1.443)	(0.421)	(1.268)	(2.787)	(-0.072)
β_{SOV}	1.053***	0.591***	0.462***	0.964***	0.388***	0.576***
	(10.783)	(5.833)	(2.668)	(8.160)	(4.917)	(3.863)
β_{MKT}	0.102**	0.086**	0.016	0.347***	0.116**	0.231***
	(2.247)	(2.041)	(0.211)	(5.521)	(2.532)	(2.983)
β_{TERM}	0.106**	-0.046	0.152*	0.023	0.068	-0.045
	(2.343)	(-0.841)	(1.737)	(0.313)	(1.327)	(-0.459)
Adj.–R^2	0.725	0.533	0.175	0.606	0.430	0.281
No. obs.	495	495	495	495	495	495

Notes: This table reports time-series regressions of off-gold, on-gold, and off-minus-on-gold portfolio excess returns on the sovereign bond market factor, the London market factor, and the British term premium factor in the IV sample. The sovereign bond market factor is the excess return on the aggregate sovereign and colonial bond portfolio. The London market factor is the excess return on the value-weighted portfolio of all British and foreign equities traded on the London Stock Exchange. The British term premium factor is the difference between the return on British consols and the short-term interest rate proxied by the 28-day banker's bill rate. The sample period is 1870–1907, with $T = 495$ four-week observations. Mean excess returns and estimated alphas are annualized by multiplying the average four-week return by 13×100. Heteroskedasticity- and autocorrelation-robust t-statistics based on Newey–West standard errors are reported in parentheses, with the lag length determined by the Bartlett-kernel rule, $\lfloor 4\,(T/100)^{2/9} \rfloor$, which equals 5 for the regression sample. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

Figure A.2: Issuer alphas and duration of gold-adherence

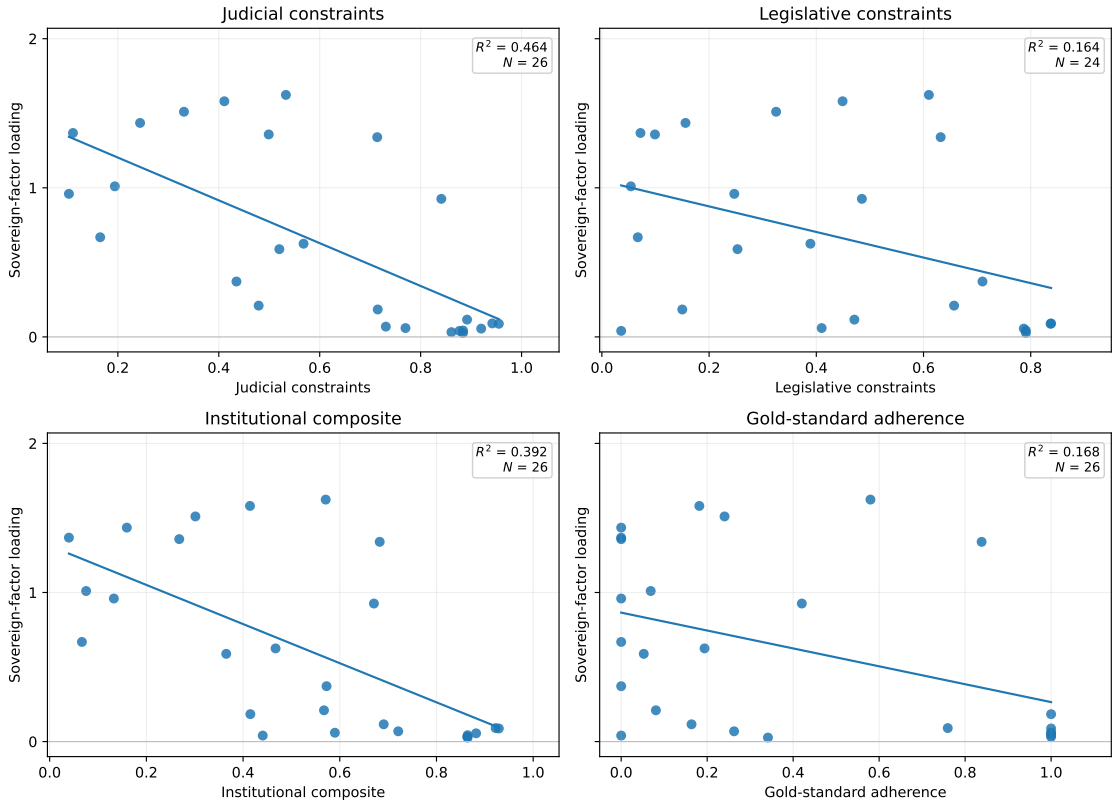


Notes: The figure plots each issuer's annualized three-factor alpha against the share of four-week sample periods in which the issuer was on gold. The left panel uses the full sample. The right panel excludes the three highest- and three lowest-alpha observations. The red line is the OLS fit.

A.3 Additional cross-sectional evidence

Appendix Figure A.3 reports the full set of bivariate relations behind the main institutional scatterplot. The figure shows that the negative relation with sovereign-factor exposure is strongest for judicial constraints and the institutional composite, weaker for legislative constraints, and less distinct for gold-standard adherence.

Figure A.3: Institutional measures, gold-standard adherence, and sovereign-factor exposure



Notes: The figure plots judicial constraints, legislative constraints, the institutional composite, and gold-standard adherence against the estimated sovereign-factor loading by issuer. Gold-standard adherence is measured as the share of four-week sample periods in which the issuer adhered to gold.

Appendix Table A.2 reports the corresponding sovereign-factor loading regressions in the OLS sample available for each institutional measure and compares them with the final regression sample. The coefficients have the same sign and similar magnitudes across the two samples, indicating that the final-sample restriction is not generating the baseline relation.

Appendix Table A.3 reports within-group regressions in the OLS sample (Panel A, $N = 47$) and the IV sample (Panel B, $N = 26$), splitting each by empire status to assess how much of the pooled institutional coefficient reflects the sovereign/colony contrast and how much reflects

Table A.2: Sovereign-factor exposure in the maximal institutional sample

	Maximal sample		Final sample	
	Coef.	N	Coef.	N
Judicial constraints	-1.139***	47	-1.306***	26
	(0.238)		(0.241)	
Legislative constraints	-0.826***	45	-0.771**	24
	(0.247)		(0.349)	
Institutional composite	-1.147***	47	-1.312***	26
	(0.222)		(0.235)	
Institutional composite, strict	-1.115***	45	-1.247***	24
	(0.225)		(0.250)	
Gold-standard adherence	-0.416**	53	-0.600***	26
	(0.164)		(0.223)	

Notes: This table reports bivariate cross-sectional regressions of the estimated sovereign-factor loading, $\hat{\beta}_{i,\mathrm{SOV}}$, on the indicated institutional measure or gold-standard adherence. The OLS sample uses all issuers with non-missing sovereign-factor loadings and the indicated regressor. The IV sample restricts the regressions to the 26 issuers used in the main cross-sectional specifications. Judicial constraints, legislative constraints, and the composite institutional measures enter in median-scaled form. Heteroskedasticity-robust standard errors are reported in parentheses, using the HC3 leverage adjustment of MacKinnon and White (1985). $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

within-group institutional variation.

Appendix Table A.7 reports the full factor-loading regressions summarized in Figure 2.

Table A.3: Sovereign-factor exposure within sub-samples: independent sovereigns and British colonies

	Pooled		Independent sovereigns		British colonies/empire	
	Coef.	N	Coef.	N	Coef.	N
Panel A. OLS sample						
Judicial constraints	-1.139***	47	-0.934***	36	-0.279**	11
	(0.238)		(0.270)		(0.130)	
Legislative constraints	-0.826***	45	-0.914***	36	-0.061	9
	(0.247)		(0.255)		(0.110)	
Institutional composite	-1.147***	47	-1.012***	36	-0.155	11
	(0.222)		(0.253)		(0.131)	
Inst. composite, strict	-1.115***	45	-1.012***	36	-0.141	9
	(0.225)		(0.253)		(0.135)	
Gold-standard adherence	-0.501***	47	-0.379	36	0.024	11
	(0.174)		(0.279)		(0.040)	
Panel B. IV sample						
Judicial constraints	-1.306***	26	-0.583	16	-0.232	10
	(0.241)		(0.513)		(0.307)	
Legislative constraints	-0.771**	24	-0.695	16	-0.060	8
	(0.349)		(0.503)		(0.107)	
Institutional composite	-1.312***	26	-0.740	16	-0.127	10
	(0.235)		(0.550)		(0.156)	
Inst. composite, strict	-1.247***	24	-0.740	16	-0.115	8
	(0.250)		(0.550)		(0.163)	
Gold-standard adherence	-0.600***	26	0.055	16	0.015	10
	(0.223)		(0.736)		(0.041)	

Notes: This table reports bivariate cross-sectional regressions of the estimated sovereign-factor loading, $\hat{\beta}_{i,\mathrm{SOV}}$, on each institutional measure or gold-standard adherence, run separately within independent sovereigns and within British colonies/imperial borrowers, with the pooled regression for comparison. Panel A uses the OLS sample of 47 issuers (36 independent sovereigns and 11 British colonies/imperial borrowers); Panel B uses the IV sample of 26 issuers (16 independent sovereigns and 10 British colonies/imperial borrowers). The institutional measures enter on the OLS-sample min–max scale used in Table 3; the within-class coefficients measure the slope per unit on that scale and are not rescaled to span the unit interval within each sub-sample, so the within-class fitted gaps between actual extremes of each sub-sample are smaller than the table coefficients. Heteroskedasticity-robust standard errors are reported in parentheses, using the HC3 leverage adjustment of MacKinnon and White (1985). The interquartile range of the institutional composite is 0.42 within independent sovereigns and 0.29 within British colonies/imperial borrowers in the OLS sample, so identifying variation in the composite is structurally compressed within the imperial subsample. Inference for sub-sample columns should be read with caution at small N. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

Table A.4: Sample comparison: kept-26 vs. dropped-27 issuers

	Kept (final 26)		Dropped (27)		Difference	
	Mean	SD	Mean	SD	$\hat{\Delta}$	t
Sovereign-factor loading	0.630	0.597	0.409	0.485	0.221	1.47
London-market loading	0.438	0.606	0.166	0.450	0.272*	1.85
Term-premium loading	0.051	0.576	0.074	0.206	-0.023	-0.19
Annualized alpha (%)	0.684	3.566	2.034	3.465	-1.350	-1.40
Average excess return (%)	3.140	2.832	3.347	3.023	-0.207	-0.26
Judicial constraints	0.576	0.311	0.583	0.296	-0.007	-0.08
Legislative constraints	0.438	0.313	0.532	0.314	-0.094	-1.00
Institutional composite	0.521	0.285	0.558	0.294	-0.037	-0.43
Gold-standard adherence	0.392	0.408	0.477	0.376	-0.085	-0.79
N	24–26		21–27			

Notes: This table reports cross-sectional means and standard deviations of the key analysis variables for the kept-26 final-sample issuers and the 27 issuers excluded by the settler-mortality and other historical-control requirements. $\hat{\Delta}$ is the difference in means (kept minus dropped); the reported t-statistic is from a two-sample Welch test with unequal variances. Sample sizes vary across rows for the institutional measures because V-Dem coverage of legislative and composite indices is incomplete for some excluded issuers; the row labelled N shows the range across rows. The factor loadings and alpha are estimated from issuer-level time-series regressions on the same three factors used in Table 3. Annualized alpha and average excess return are reported in percentage points. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$ on the difference.

Table A.5: Generated-regressor inference for the cross-sectional regressions of estimated factor loadings

	Coef.	HC3 SE	Bootstrap SE	Bootstrap 95% CI
Judicial constraints	-1.139***	(0.238)	(0.127)	[-1.384, -0.887]
Legislative constraints	-0.826***	(0.247)	(0.120)	[-1.052, -0.569]
Institutional composite	-1.147***	(0.222)	(0.136)	[-1.403, -0.865]
Inst. composite, strict	-1.115***	(0.225)	(0.138)	[-1.381, -0.827]

Notes: This table compares two inference approaches for the cross-sectional regressions of estimated sovereign-factor loadings on the indicated institutional measure in the OLS sample of 47 issuers with non-missing baseline institutional composite. HC3 SE is the heteroskedasticity-robust standard error reported in Table 3. Bootstrap SE is the standard deviation across 500 block-bootstrap replications; each replication resamples non-overlapping blocks of 13 consecutive four-week periods (roughly one year) with replacement, re-estimates each issuer's three-factor time-series betas on the resampled time series, and reruns the cross-section. This procedure incorporates both cross-sectional heteroskedasticity and the first-stage estimation noise in $\hat{\beta}_{i,\mathrm{SOV}}$ (the generated-regressor channel discussed in Lewellen et al., 2010). $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$ on the HC3 standard error.

Table A.6: Bootstrap inference for the pricing decomposition

	Baseline (pp)	Bootstrap SE	Bootstrap 95% CI
London market factor	1.644	(1.101)	[0.120, 4.302]
Sovereign factor	2.939	(1.232)	[0.694, 5.489]
Term premium factor	0.003	(0.187)	[-0.419, 0.367]
Alpha	-1.350	(2.728)	[-7.251, 3.147]
Fitted total	3.235	(3.671)	[-4.342, 10.187]
Raw-return gap	3.931	(3.933)	[-4.167, 11.612]

Notes: This table reports bootstrap inference for the pricing decomposition in Table 5 at the OLS sample of 47 issuers with non-missing baseline institutional composite. Each entry is the 1-unit-move contribution to the institutional raw-return gap; scaling the entries by 0.41 delivers the interquartile-range figures discussed in the main text. Bootstrap SEs and percentile-based 95% confidence intervals are from 500 date-block bootstrap replications (non-overlapping 13-period blocks). The bootstrap propagates time-series estimation noise in factor means and issuer-level betas through the decomposition. The sovereign-factor contribution's confidence interval excludes zero; the alpha gap is imprecise.

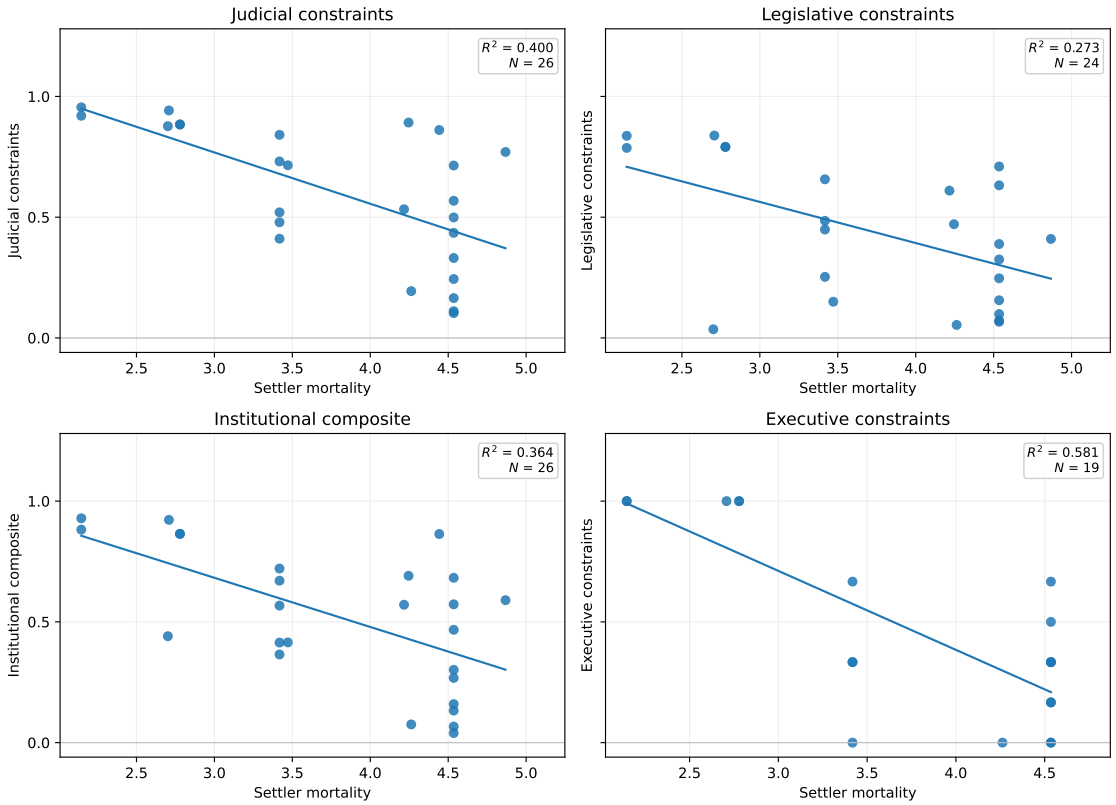
Table A.7: Institutions and factor loadings

Panel A. Sovereign-factor loading	(1)	(2)	(3)	(4)	(5)
Judicial constraints	-1.306***				
	(0.241)				
Legislative constraints		-0.771**			
		(0.349)			
Institutional composite			-1.312***		
			(0.235)		
Institutional composite, strict				-1.247***	
				(0.250)	
Gold-standard adherence					-0.600***
					(0.223)
adj.-R^2	0.442	0.126	0.366	0.324	0.133
N	26	24	26	24	26
Panel B. London-market loading	(1)	(2)	(3)	(4)	(5)
Judicial constraints	-0.935*				
	(0.494)				
Legislative constraints		-0.493			
		(0.416)			
Institutional composite			-0.913*		
			(0.527)		
Institutional composite, strict				-0.852	
				(0.543)	
Gold-standard adherence					-0.648***
					(0.242)
adj.-R^2	0.199	0.020	0.150	0.116	0.157
N	26	24	26	24	26
Panel C. Term-premium loading	(1)	(2)	(3)	(4)	(5)
Judicial constraints	-0.050				
	(0.679)				
Legislative constraints		-0.019			
		(0.517)			
Institutional composite			-0.044		
			(0.722)		
Institutional composite, strict				-0.041	
				(0.752)	
Gold-standard adherence					0.019
					(0.278)
adj.-R^2	-0.041	-0.045	-0.041	-0.045	-0.041
N	26	24	26	24	26

Notes: This table reports cross-sectional regressions of estimated issuer-level factor loadings on institutional quality and gold-standard adherence in the IV sample. Panels A, B, and C use the sovereign-factor loading, London-market loading, and term-premium loading as dependent variables, respectively. Columns (1)–(5) correspond to judicial constraints, legislative constraints, the baseline institutional composite, the strict institutional composite, and gold-standard adherence. Institutional measures enter in median-scaled form. Heteroskedasticity-robust standard errors are reported in parentheses, using the HC3 leverage adjustment of MacKinnon and White (1985). $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

Appendix Figure A.4 plots updated log settler mortality against four institutional measures by issuer: judicial constraints, legislative constraints, and the institutional composite—three of the four first-stage specifications reported in Table 6—together with Polity executive constraints as a supplementary measure. The strict-composite first stage in Table 6 is omitted because it is near-identical to the baseline composite by construction.

Figure A.4: Settler mortality and institutional quality



Notes: The figure plots updated log settler mortality against each institutional measure by issuer. The institutional measures are judicial constraints, legislative constraints, the institutional composite, and executive constraints.

A.4 Proofs of Propositions 1 and 2

The proofs use the loss decomposition, the institutional-quality and monotonicity-in-z conditions on each component, the conditional-independence stipulation $L_i^T \perp f_{\text{SOV}} \mid D$, and the covariance and level conditions, all stated in Section 3.

Proof of Proposition 1. The factor-driven component of the gross return on issuer i's bond is

$$R_i^F(z, \eta_i; I_i) = \frac{1 - L_i^F(z, \eta_i; I_i)}{q_i(I_i)}.$$

Because the constant 1 contributes nothing to covariances and q_i depends only on I_i,

$$\operatorname{Cov}\big(R_i^F, f_{\text{SOV}}\big) = -\frac{1}{q_i(I_i)} \operatorname{Cov}\big(L_i^F, f_{\text{SOV}}\big),$$

and the sovereign-factor loading is

$$\beta_{i,\text{SOV}} = \frac{\operatorname{Cov}(R_i^F, f_{\text{SOV}})}{\operatorname{Var}(f_{\text{SOV}})} = -\frac{\operatorname{Cov}(L_i^F, f_{\text{SOV}})}{q_i(I_i) \operatorname{Var}(f_{\text{SOV}})}.$$

Differentiating in I_i,[21]

$$\frac{\partial \beta_{i,\text{SOV}}}{\partial I_i} = -\frac{1}{q_i(I_i) \operatorname{Var}(f_{\text{SOV}})} \frac{\partial \operatorname{Cov}(L_i^F, f_{\text{SOV}})}{\partial I_i} + \underbrace{\frac{\operatorname{Cov}(L_i^F, f_{\text{SOV}})\, q_i'(I_i)}{q_i(I_i)^2 \operatorname{Var}(f_{\text{SOV}})}}_{\text{price effect}}.$$

Under the covariance condition on L_i^F, $\partial \operatorname{Cov}(L_i^F, f_{\text{SOV}})/\partial I_i \geq 0$, so the first term is weakly negative. For the price effect, $\operatorname{Cov}(L_i^F, f_{\text{SOV}}) \leq 0$ (factor-driven losses rise when the sovereign factor is lower) and $q_i'(I_i) \geq 0$.[22] The price effect is therefore weakly negative.[23] Both terms are weakly negative, so $\partial \beta_{i,\text{SOV}}/\partial I_i \leq 0$, with strict inequality when the covariance condition on L_i^F binds

[21] The proof requires $q_i(I_i)$ differentiable in I_i and $\operatorname{Var}(f_{\text{SOV}}) > 0$, treated as exogenous to I_i. The latter is innocuous when the weight of issuer i in the value-weighted factor is small. Both conditions are satisfied in the present setting.

[22] From the price equation $q_i(I_i) = \mathbb{E}[m(z)(1 - L_i(z, \eta_i; I_i))]$, $\partial q_i/\partial I_i = -\mathbb{E}[m(z)\, \partial L_i/\partial I_i]$, which is non-negative given $m(z) \geq 0$ under no-arbitrage and $\partial L_i^c/\partial I_i \leq 0$ for $c \in \{F, T\}$ by the institutional-quality conditions on the loss components.

[23] The price-effect channel introduces an interpretive caveat. Because R_i^F is deflated by the endogenous price $q_i(I_i)$, cross-sectional differences in $\hat{\beta}_{i,\text{SOV}}$ partly reflect cross-sectional differences in price levels rather than payoff covariances. The sign of the comparative static $\partial \beta_{i,\text{SOV}}/\partial I_i \leq 0$ is unaffected, but the price-level correction influences its magnitude. A direct check would re-estimate the sovereign factor loading using yields rather than returns, or using excess returns scaled by duration.

with strict inequality. □

Remark 1 (Baseline three-factor regression). *The proof concerns the slope on f_{SOV} derived from R_i^F. In the issuer-level baseline regression $r_{i,t}^e = \alpha_i + \beta_{i,\mathrm{SOV}} f_{\mathrm{SOV},t} + \beta_{i,\mathrm{MKT}} f_{\mathrm{MKT},t} + \beta_{i,\mathrm{TERM}} f_{\mathrm{TERM},t} + u_{i,t}$, which lacks an explicit tail-state dummy D_t, the slope absorbs an additional channel from L_i^T even under conditional independence, because $\mathbf{1}\{D=1\}\, L_i^T$ correlates with f_{SOV} through the binary state D. With D binary,*

$$\mathrm{Cov}\big(\mathbf{1}\{D=1\}\, L_i^T,\, f_{\mathrm{SOV}}\big) = p(1-p)\,\mathbb{E}[L_i^T \mid D=1]\,\big(\mathbb{E}[f_{\mathrm{SOV}} \mid D=1] - \mathbb{E}[f_{\mathrm{SOV}} \mid D=0]\big) \le 0,$$

where $p = \Pr(D=1)$. The level condition implies $\partial\,\mathbb{E}[L_i^T \mid D=1]/\partial I_i \le 0$, so $\partial\,\mathrm{Cov}(\mathbf{1}\{D=1\}\, L_i^T,\, f_{\mathrm{SOV}})/\partial I_i \ge 0$. The additional channel reinforces the comparative static $\partial\beta_{i,\mathrm{SOV}}/\partial I_i \le 0$ in the baseline regression rather than working against it.

Proof of Proposition 2. By the decomposition $L_i = L_i^F + \mathbf{1}\{z < z^*\}\, L_i^T$, the gross return decomposes as $R_i = R_i^F - \mathbf{1}\{D=1\}\, L_i^T/q_i(I_i)$. Under the conditional-independence stipulation $L_i^T \perp f_{\mathrm{SOV}} \mid D$, L_i^T is orthogonal to f_{SOV} given D, so it does not contribute to the sovereign-factor loading and instead identifies the tail-state intercept in the linear factor representation:

$$\alpha_{i,\mathrm{tail}} = -\frac{\mathbb{E}[L_i^T \mid D=1]}{q_i(I_i)}.$$

Differentiating in I_i,

$$\frac{\partial \alpha_{i,\mathrm{tail}}}{\partial I_i} = -\frac{1}{q_i(I_i)}\,\frac{\partial\,\mathbb{E}[L_i^T \mid D=1]}{\partial I_i} + \frac{\mathbb{E}[L_i^T \mid D=1]\, q_i'(I_i)}{q_i(I_i)^2}.$$

Under the level condition, $\partial\,\mathbb{E}[L_i^T \mid D=1]/\partial I_i \le 0$, so the first term is weakly positive. For the second term, $L_i^T \ge 0$ and $q_i'(I_i) \ge 0$ (as derived in the footnote to the proof of Proposition 1), so it is weakly positive. Both terms are weakly positive, giving $\partial\alpha_{i,\mathrm{tail}}/\partial I_i \ge 0$, with strict inequality when the level condition binds with strict inequality. □

Remark 2 (Strength of the conditions and microfoundation). *The cross-partial restriction $\partial^2 L_i^c/\partial I_i\,\partial z \ge 0$, $c \in \{F,T\}$—that the institutional reduction in losses is concentrated state-by-state in bad aggregate states—is a strictly stronger structural condition than the covariance and level conditions adopted in the*

proofs. It places a sign restriction on each component pointwise in (I_i, z), whereas the covariance and level conditions are integrated conditions on the joint distribution of L_i^c and the relevant aggregate object. The proofs use the weaker integrated conditions because the empirical analysis tests them. The cross-partial enters the empirical work as a stronger statement nested in the same panel specification: applied to L_i^F, it implies that the institutional slope on $\hat{\beta}_{i,\mathrm{SOV}}$ should steepen in left-tail realizations of f_{SOV}, and applied to L_i^T, it implies that the institutional effect on tail-state intercepts is concentrated in the most extreme tail realizations. The empirical analysis estimates a panel specification that nests both signatures.

The covariance condition on L_i^F can be derived from the cross-partial restriction under explicit structural assumptions on the joint distribution of $(L_i^F, z, f_{\mathrm{SOV}})$. Decompose

$$\mathrm{Cov}(L_i^F, f_{\mathrm{SOV}}) = \mathbb{E}\big[\mathrm{Cov}(L_i^F, f_{\mathrm{SOV}} \mid z)\big] + \mathrm{Cov}\big(\mathbb{E}[L_i^F \mid z],\, \mathbb{E}[f_{\mathrm{SOV}} \mid z]\big)\,.$$

Suppose (a) $\mathbb{E}[f_{\mathrm{SOV}} \mid z]$ is increasing in z, capturing the idea that the sovereign factor proxies aggregate sovereign-credit conditions; and (b) conditional on z, L_i^F and f_{SOV} are independent, which holds when the residual $f_{\mathrm{SOV}} - \mathbb{E}[f_{\mathrm{SOV}} \mid z]$ is independent of η_i — a natural small-issuer condition for a value-weighted aggregate. Under (a) and (b), the first term is zero, and differentiating in I_i gives

$$\frac{\partial\, \mathrm{Cov}(L_i^F, f_{\mathrm{SOV}})}{\partial I_i} = \mathrm{Cov}\left(\frac{\partial\, \mathbb{E}[L_i^F \mid z]}{\partial I_i},\, \mathbb{E}[f_{\mathrm{SOV}} \mid z]\right).$$

The cross-partial restriction $\partial^2 L_i^F / \partial I_i\, \partial z \geq 0$ implies that $\partial\, \mathbb{E}[L_i^F \mid z]/\partial I_i$ is itself increasing in z. Combined with (a), Chebyshev's covariance inequality (two functions monotone in the same direction with respect to a common scalar) yields a non-negative right-hand side, and the covariance condition follows. The same structure delivers $\mathrm{Cov}(L_i^F, f_{\mathrm{SOV}}) \leq 0$ used in the proof's price-effect step. Under $\partial L_i^F/\partial z \leq 0$, $\mathbb{E}[L_i^F \mid z]$ is decreasing in z, so the relevant Chebyshev pairing has opposite-monotone functions. The level condition on L_i^T admits a parallel microfoundation. The cross-partial $\partial^2 L_i^T / \partial I_i\, \partial z \geq 0$ implies that $\partial\, \mathbb{E}[L_i^T \mid z]/\partial I_i$ is increasing in z, so conditioning on $z < z^*$ delivers $\partial\, \mathbb{E}[L_i^T \mid z < z^*]/\partial I_i \leq 0$.

Remark 3 (Multivariate factor specification)**.** *The proof of Proposition 1 is stated for the univariate loading $\beta_{i,\mathrm{SOV}}^{\mathrm{u}} = \mathrm{Cov}(R_i^F, f_{\mathrm{SOV}})/\mathrm{Var}(f_{\mathrm{SOV}})$. The empirical specification estimates the partial loading $\beta_{i,\mathrm{SOV}}$ from a multivariate regression on $(f_{\mathrm{SOV}}, f_{\mathrm{MKT}}, f_{\mathrm{TERM}})$, which by the Frisch–Waugh–Lovell theorem equals the slope on the residualized sovereign factor $\tilde{f}_{\mathrm{SOV}} = f_{\mathrm{SOV}} - \lambda_{\mathrm{MKT}} f_{\mathrm{MKT}} - \lambda_{\mathrm{TERM}} f_{\mathrm{TERM}}$,*

where $(\lambda_{\mathrm{MKT}}, \lambda_{\mathrm{TERM}})$ *depend on the joint distribution of the factors but not on* I_i*. The comparative static* $\partial\beta_{i,\mathrm{SOV}}/\partial I_i \leq 0$ *extends to the multivariate loading under either of two sufficient conditions: (i) institutions do not differentially affect* $\mathrm{Cov}(L_i^F, f_{\mathrm{MKT}})$ *or* $\mathrm{Cov}(L_i^F, f_{\mathrm{TERM}})$*, in which case the partial-coefficient identity reduces the multivariate* $\partial\mathrm{Cov}(L_i^F, \tilde{f}_{\mathrm{SOV}})/\partial I_i$ *to a positive multiple of* $\partial\mathrm{Cov}(L_i^F, f_{\mathrm{SOV}})/\partial I_i$*; or (ii)* $\tilde{f}_{\mathrm{SOV}}$ *inherits monotonicity in* z *from* f_{SOV}*, in which case the microfoundation argument above applies directly to* $\tilde{f}_{\mathrm{SOV}}$*. Figure 2 provides empirical support for (i). The institutional coefficient is concentrated in the sovereign-factor loading, is small for the London market loading, and is indistinguishable from zero for the term premium loading.*

A.5 Institutional mechanisms and credible commitment

This appendix collects the historical evidence and the mechanism-by-mechanism discussion behind the reduced-form mapping from institutional quality I_i to the payoff-loss function $L_i(z, \eta_i; I_i)$. The formal proof of Proposition 1 appears in Appendix A.4. The model does not separately parameterize an enforcement penalty. Instead, mechanisms that raise the political, legal, or market costs of nonpayment are associated with lower creditor losses, L_i, particularly in adverse aggregate states. The empirical analysis uses V-Dem's judicial-constraints and legislative-constraints indices and Polity's executive-constraints index as proxies for I_i. These variables are interpreted as capturing mechanisms that constrain sovereign discretion, support enforcement, and improve the credibility of repayment commitments.

Contemporary sources provide direct evidence that investors priced institutional constraints in the London sovereign bond market. Withers (1916, p. 40) observed that international lending required that political and social conditions had gained sufficient stability for long-term contracts to be credible, and that the growth of British foreign investment was driven in part by the greater political stability that developed in borrowing countries. The same source noted that Mexico's default demonstrated that the political stability investors had believed it to have achieved was illusory, linking default risk directly to the failure of institutional credibility (Withers, 1916, p. 42). Commentary from the period more broadly framed sovereign creditworthiness in terms of the protection of creditor claims and the credibility of repayment commitments (Corporation of Foreign Bondholders, 1877, 1894; Powell, 1910, p. 15), with borrowing costs varying directly with repayment reputation—from Britain borrowing at 3 percent to Honduras, described as a hardened defaulter, unable to obtain credit on any terms. As early as 1874, the Corporation of Foreign Bondholders identified agreed principles governing the liability of sovereign states as a prerequisite for international lending (Mauro and Yafeh, 2003, p. 73). London investors were also well informed about political and institutional developments in borrowing countries, which were meticulously reported in the financial and general press, partly in response to investor demand for this information as a key input in investment decisions (Mauro et al., 2006, pp. 16–17).

A first mechanism operates through constraints on executive authority. Stronger executive constraints limit the government's ability to unilaterally suspend payments, alter contractual

terms, or divert resources from debt service. In the model, this corresponds to a reduction in $L_i(z, \eta_i; I_i)$ by raising the expected cost of default through political and legal channels.

A second mechanism operates through legal and administrative capacity: the extent to which contractual claims can be enforced and revenues can be credibly committed to debt service. The V-Dem judicial-constraints index used in the regressions captures this mechanism indirectly by measuring the strength of judicial independence and procedural checks on the executive, which together govern whether revenue allocations to debt service are protected from ex post discretion. Historically, sovereign borrowing arrangements often relied on the assignment of specific revenue streams to debt repayment and, in some cases, on delegation to intermediary institutions with partial control over those revenues. These arrangements limited ex post discretion and reduced expected creditor losses associated with nonpayment. Contemporary investment manuals made this pricing logic explicit: Powell (1910, pp. 147–151) classified sovereign bonds according to a hierarchy of political and constitutional development, in which stronger constitutional authority and the absence of prolonged political upheaval placed issuers in higher investment classes, while weak or inchoate constitutional arrangements degraded them, with persistent defaulters falling outside the range of investment classification altogether (Powell, 1910, pp. 158–159).

A within-country illustration of the delegation channel is the third charter of the Banco do Brasil, which from 1853 to 1866 received a monopoly over note issue and the responsibility for managing the exchange rate against an explicit Treasury commitment to refrain from new currency issues. Using monthly secondary-market prices on Brazilian *apólices* in Rio de Janeiro and London-traded Brazilian sovereign bonds across the imperial era, Summerhill (2008) estimates ordinary least squares specifications that include controls for the funded debt stock, the deficit, political uprisings, the Paraguayan War, and the period of suspended payments on the Portuguese loan; the indicator for the delegation period enters with coefficients of approximately -180 basis points on the London risk premium and -70 basis points on Rio *apólice* yields, both significant at conventional levels. A Johansen cointegrating equation that pairs the delegation indicator with controls for the Portuguese loan suspension and the debt stock yields an analogous long-run effect of roughly -128 basis points on the domestic perpetuity yield. The same study identifies persistent structural breaks in both yield series whose timing aligns with political-institutional events—including the Additional Act of 1834, the resolution of regional uprisings, and the wartime re-

sumption of Treasury currency issue in 1866—rather than with fundamental shifts in fiscal capacity. The pattern is consistent with the interpretation that institutional constraints on sovereign discretion, including delegation to bodies with separate governance, move the risk premium beyond what borrower fundamentals alone would predict.

A third mechanism operates through the common external enforcement environment in which all issuers borrow, rather than through cross-sectional variation in I_i. Investors in international capital markets developed institutions to collect information, monitor sovereign behavior, and coordinate responses to defaults. Organizations such as the Corporation of Foreign Bondholders improved creditor bargaining power and, in some cases, restricted access to new borrowing following default. The CFB's own framing of its mission captures this enforcement logic directly: its 1873 Annual Report stated that bondholders had a duty to preserve common action, "maintaining and promoting credit to the honest, and inflicting penalties on dishonest governments" (Mauro and Yafeh, 2003, p. 43). Because the CFB and similar coordination institutions were market-wide features common to all issuers in this sample, they shape the baseline cost of default embedded in L_i but do not generate the cross-sectional differences across issuers identified by the empirical analysis. The external-enforcement and domestic-constraints channels are complementary: external enforcement raises the equilibrium cost of default that all sovereigns face, and domestic constraints determine how credibly different sovereigns can precommit to bearing that cost.

The London Stock Exchange operated as a parallel enforcement institution. From 1827 onward, the Committee for General Purposes refused to recognize new issues from sovereigns in arrears on previous London obligations, and the rule remained in the Exchange's rule book throughout the sample period (Attard, 2013). The threat of denied quotation also induced observable institutional adjustment by borrowers seeking market access. A documented instance is the consolidation of New Zealand's provincial debts into a single colonial stock after the Committee declined to recognize provincial issues. Borrower-side adjustment of this kind is consistent with the reduced-form mapping from I_i to L_i in Section 3, in which institutional quality affects expected creditor losses in part because issuers operate under credible threats of market exclusion.

These mechanisms imply that the institutional variables used in the empirical analysis should be interpreted as proxies for a broader set of constraints and enforcement technologies that affect repayment incentives. The reduced-form relation between institutional quality and sovereign bor-

rowing costs reflects the extent to which these mechanisms increase the credibility of repayment commitments, rather than capturing institutional characteristics per se (North and Weingast, 1989; Eichengreen et al., 2021).

This interpretation also motivates the inclusion of competing explanations—such as geography, legal origin, and colonial history—as controls or indices in the empirical specification, since these factors may affect borrowing costs through channels that are distinct from the commitment mechanisms captured by I_i.

A.6 Data description

Foreign sovereign and colonial bond returns

Bond prices are sampled at 28-day intervals from the official quotation lists published in the *Money Market Review* and *The Economist* from 1866 through 1907, the last year in which those sources reported quotations at the level of detail required for the return calculations. These bond-return data are drawn from the same underlying London-market security-return database used in a study of more than 500,000 four-week returns on British and foreign securities from 1866 to 1907 (Chabot and Kurz, 2010). I combine prices, coupon information, and issuance data to construct realized four-week holding-period returns, adjusting for coupon payments and default events. Default dates follow the annual reports of the Council of Foreign Bondholders and standard historical default chronologies (Winkler, 1933; Suter, 1990). To limit transcription and classification errors, each bond return is linked to an audit trail that covers the quoted price, coupon schedule, issue characteristics, and any default-related adjustments. The regression sample runs from January 1870 through December 1907 and contains 495 four-week observations.

British and foreign equity returns

The equity sample consists of 1,808 firms from 52 countries that were traded on the London Stock Exchange at some point between 1866 and 1907. Price data come from the official Friday lists in the *Money Market Review*, while dividend histories and shares outstanding are drawn from the *Investors Monthly Manual* and *The Economist*. These data are used to construct 28-day holding-period returns. Additional details appear in prior work using the same underlying security-return database (Chabot and Kurz, 2010). The regression sample again runs from January 1870 through December 1907 and contains 495 four-week observations.

Risk-free rate

I use short-term London bank bills with maturities of 30-60 days and geometrically scale the annualized rate to obtain a 28-day risk-free rate. When the shortest-maturity bill is unavailable, I use scaled versions of the two- and three-month bills. The data are from *The Economist*.

Gold-standard adherence

Gold-standard adherence is measured by an indicator that equals 1 if a country remained on gold

for the full four-week return interval and 0 otherwise. The underlying gold-standard data are from prior return-based work and are updated here to incorporate additional countries, revised bond classifications, and corrected adherence dates (Alquist and Chabot, 2011).

Settler mortality

I use the alternative European settler mortality series ("Naval Stations 2"), capped at 250 deaths per 1,000 per annum, from Acemoglu et al. (2012). This series is used in updated versions of the analysis (Johnson, 2025). All specifications use the logarithm of mortality per 1,000 individuals.

V-Dem institutional measures

I use two institutional measures from the Varieties of Democracy (V-Dem) project: the legislative constraints on the executive index, denoted in the V-Dem codebook as `v2xlg_legcon`, and the judicial constraints on the executive index, denoted as `v2x_jucon`. The legislative constraints index measures the extent to which the legislature and oversight agencies can question, investigate, and exercise oversight of the executive. The judicial constraints index measures the extent to which the executive respects the constitution and complies with court rulings, and the extent to which the judiciary can act independently. Both variables are interval indices on a 0–1 scale constructed from underlying expert-coded indicators using V-Dem's Bayesian measurement model. In the empirical analysis, I aggregate the annual observations from 1870–1907 to form country-level measures that are matched to the bond-return sample (V-Dem Institute, 2026).[24] From these two indices, I construct two composites used in the cross-sectional regressions. The baseline composite is the simple average of the judicial and legislative constraints indices, using whichever index is available if the other is missing. The strict composite is the same average but restricts the sample to issuers with both indices non-missing.

Executive constraints

The executive constraints variable measures the extent of institutionalized constraints on chief executives' decision-making powers, whether imposed by legislatures, the judiciary, party elites, ruling councils, or other accountability groups. It is operationalized as a 7-category scale and typically used as an indicator of horizontal accountability or de jure checks on executive power,

[24] V-Dem Country-Year Dataset, version 16, V-Dem Institute, University of Gothenburg, available at https://www.v-dem.net/.

with higher values indicating stronger institutional constraints on the executive.[25]

Legal origin

The legal-origin data classify countries according to the source of company law or commercial code—English, French, Socialist, German, or Scandinavian—following the law-and-finance classification (La Porta et al., 1998).

British colony

This indicator equals one if the bond was issued by a British colony, possession, or other public borrower within the British Empire, and zero otherwise.

Geographic and disease environment

I use mean annual temperature, the share of land area within 100 kilometers of the coast, and absolute latitude from standard geography controls, as reported in subsequent development-accounting work (Gallup et al., 1999; McArthur and Sachs, 2001). I also use the country-level malaria ecology index (ME), which measures the stability of malaria transmission based on climatic conditions and the biological characteristics of regionally dominant malaria vectors (Kiszewski et al., 2004). I transform the malaria variable as $\ln(1 + \mathrm{ME}_j)$ for country j. To construct the geography-and-disease index, I standardize the component variables to have a mean of 0 and unit variance, orient them so that higher values correspond to higher modern income, and then take their arithmetic average.

Historical state capacity

The log of population density in 1500 is taken from research that treats population density at the onset of European colonization as a proxy for precolonial prosperity and broader social complexity (Acemoglu et al., 2002).

Share of countries in external default

Aggregate sovereign default stress is derived from the Reinhart–Rogoff annual series on the share of sovereign defaults in external defaults across all countries in their sample (Reinhart and Rogoff, 2009). The series is available annually from 1800 through 2014 and provides a broad measure of

[25] Polity5 Project, Center for Systemic Peace, dataset covering 1800–2018, available at https://www.systemicpeace.org/polityproject.html.

global sovereign credit distress, rather than a count limited to issuers in the London bond-return sample. I use the 1870–1907 overlap with the return data.

A.7 Risk factor construction

The empirical analysis employs a linear factor model in the spirit of arbitrage pricing models, in which a small number of pervasive risk factors drives excess returns. The factors are constructed from the same underlying London market data used in the main analysis. The factors serve as controls in time-series regressions of bond returns, capturing common sources of variation across securities. They are used to construct risk-adjusted returns for issuer portfolios, which form the basis of the paper's cross-sectional analysis.

A.7.1 London market factor

The London market factor is defined as the excess return on the value-weighted portfolio of all British and foreign equities traded on the London Stock Exchange. This factor captures broad fluctuations in economic activity and aggregate risk conditions that affect asset prices. The use of an equity-based market factor follows standard practice in empirical asset pricing and reflects the absence of a comprehensive bond market index for the period.

A.7.2 British term premium factor

The British term premium factor is constructed as the difference between the return on long-term British government debt (consols) and the short-term interest rate. The short rate is proxied by the yield on 28-day banker's bills. This factor captures variation in the slope of the yield curve and is motivated by evidence that term premia are an important determinant of bond returns.

A.7.3 Sovereign bond market factor

The sovereign bond market factor is defined as the excess return on the value-weighted portfolio of sovereign and colonial bonds in the sample. This factor captures common shocks to sovereign credit conditions in the London market. In the context of the paper's main results, it is the most direct empirical proxy for the systematic component of sovereign risk.

The factor is constructed from the same four-week return panel used to estimate issuer-level returns, with portfolio weights based on market values. Because it aggregates across issuers, the factor captures movements that are common to sovereign debt as an asset class rather than

country-specific repayment events. A high sovereign-factor loading indicates that an issuer's bonds comove strongly with broad sovereign bond market shocks. A central empirical question in the paper is whether issuers with weaker commitment institutions loaded more heavily on the sovereign factor.

A.7.4 Leave-one-out sovereign bond market factor

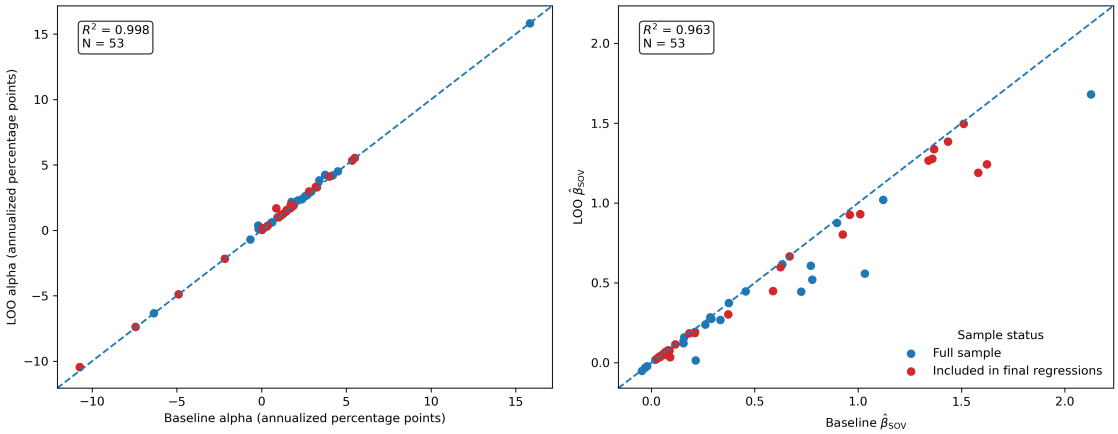
Because the sovereign bond market factor is constructed from the same asset class as the issuer portfolios used in the time-series regressions, I re-estimate each issuer's alpha and sovereign-factor loading using a leave-one-out sovereign factor that excludes the issuer under consideration. Appendix Figure A.5 plots the baseline estimates against the corresponding leave-one-out estimates. The annualized alphas remain highly stable across leave-one-out samples, with the points lying close to the 45-degree line. The leave-one-out sovereign-factor loadings are highly correlated with the baseline sovereign-factor loadings, but they are not identical. For issuers with positive baseline loadings, the leave-one-out estimates tend to lie below the baseline estimates, consistent with the mechanical upward component introduced when an issuer's own return enters the aggregate sovereign bond market factor. The relevant implication for the paper is that the cross-sectional ordering is largely preserved after removing this mechanical component, which supports the baseline sovereign-factor construction used in the main analysis.[26]

A.7.5 Alternative sovereign bond market factor constructions

The baseline sovereign bond market factor is value-weighted across sovereign and colonial issuer portfolios. This construction reflects an investor perspective, but it raises two related concerns. Large issuers could dominate the factor, or it could primarily reflect the imperial segment of the London sovereign bond market. To assess these possibilities, I re-estimate issuer-level

[26] The simultaneous stability of $\hat{\alpha}$ and the larger movement in $\hat{\beta}_{\mathrm{SOV}}$ may at first appear inconsistent. Two features reconcile them under the alpha identity $\hat{\alpha}_i = \bar{r}_i - \sum_k \hat{\beta}_{i,k}\bar{f}_k$. First, the leave-one-out sovereign factor has a different sample mean from the baseline factor because the issuer's own returns are removed from the value-weighted aggregate; for the highest-weight issuers, this displaces the baseline factor mean by 0.1–0.3 percentage points annualized. Second, the leave-one-out regression also re-estimates the global-market and term-premium loadings, which adjust, partially offsetting the change in the sovereign-factor contribution. Reproducing the per-issuer $\Delta\hat{\alpha}$ from these three components yields a near-exact match with the actual estimates ($R^2 \approx 0.88$ between predicted and actual $\Delta\hat{\alpha}$); the largest $|\Delta\hat{\alpha}|$ is below 1 percentage point even for issuers with $|\Delta\hat{\beta}_{\mathrm{SOV}}| \approx 0.4$–$0.5$. The decomposition is reproduced in the appendix notebook `appendix_tables_figures.ipynb` and saved in `loo_alpha_decomposition.csv`.

Figure A.5: Baseline and leave-one-out issuer alphas and sovereign bond market factor loadings



Notes: The left panel plots each issuer's annualized three-factor alpha from the baseline model against the corresponding estimate using a leave-one-out sovereign bond market factor that excludes the issuer's own returns. The right panel plots the baseline sovereign-factor loading against its leave-one-out counterpart. Blue markers are issuers in the full return sample; red markers are issuers in the final cross-sectional regressions. Dashed lines are 45-degree lines.

sovereign-factor loadings using alternative factor constructions: leave-one-out, equal-weighted, value-weighted foreign-bond, and equal-weighted foreign-bond factors. The leave-one-out column is included here for direct comparison with the other alternative factors. I then repeat the main cross-sectional regressions in the maximal institutional sample and in the IV sample.

Appendix Table A.8 reports the results. The estimated institutional coefficient remains negative across all factor constructions and both samples. The estimates are not identical, as expected, because the factors use different issuer weights and market segments, but the central institutional coefficient does not depend on the value-weighted all-issuer baseline factor. Gold-standard adherence does not materially account for this relation when included alongside the institutional composite.

Table A.8: Institutions and sovereign-factor exposure: alternative factor constructions

Sovereign-factor loading	Baseline VW	Leave-one-out	Equal-weighted	VW foreign	EW foreign	LOO foreign	LOO colonial	HI subsample
Panel A. OLS sample								
Composite only								
Institutional composite	-1.147***	-1.045***	-2.118***	-1.062***	-1.610***	-0.967***	-0.718*	-1.135***
	(0.222)	(0.189)	(0.381)	(0.205)	(0.287)	(0.175)	(0.403)	(0.347)
adj.-R^2	0.338	0.367	0.348	0.340	0.360	0.366	0.097	0.235
N	47	47	47	47	47	47	47	47
Composite and gold-standard adherence								
Institutional composite	-1.272***	-1.101***	-1.499***	-1.168***	-1.167***	-1.011***	-0.770*	-1.134**
	(0.424)	(0.332)	(0.560)	(0.393)	(0.421)	(0.310)	(0.424)	(0.465)
Gold-standard adherence	0.139	0.062	-0.690	0.118	-0.495	0.048	0.058	-0.001
	(0.321)	(0.253)	(0.430)	(0.298)	(0.321)	(0.237)	(0.208)	(0.267)
adj.-R^2	0.329	0.354	0.372	0.329	0.381	0.353	0.077	0.217
N	47	47	47	47	47	47	47	47
Panel B. IV sample								
Composite only								
Institutional composite	-1.312***	-1.271***	-2.564***	-1.222***	-1.960***	-1.181***	-0.707	-1.200**
	(0.235)	(0.217)	(0.532)	(0.214)	(0.403)	(0.197)	(0.650)	(0.560)
adj.-R^2	0.366	0.425	0.403	0.373	0.414	0.430	0.053	0.184
N	26	26	26	26	26	26	26	26
Composite and gold-standard adherence								
Institutional composite	-1.394***	-1.375***	-2.127**	-1.292***	-1.641**	-1.273***	-1.011	-1.133
	(0.478)	(0.437)	(0.866)	(0.441)	(0.670)	(0.404)	(0.785)	(0.715)
Gold-standard adherence	0.081	0.104	-0.436	0.070	-0.318	0.092	0.303	-0.066
	(0.412)	(0.375)	(0.686)	(0.383)	(0.527)	(0.350)	(0.341)	(0.322)
adj.-R^2	0.341	0.404	0.391	0.347	0.402	0.408	0.031	0.149
N	26	26	26	26	26	26	26	26

Notes: This table reports cross-sectional regressions of estimated sovereign-factor loadings on the composite institutional measure and gold-standard adherence under alternative constructions of the sovereign bond market factor. The dependent variable is the issuer-level loading on the column-specific sovereign factor from a three-factor time-series model that also includes the London market factor and the British term premium factor. Baseline VW uses the value-weighted sovereign bond market factor. Leave-one-out excludes each issuer's own returns from the sovereign factor used to estimate that issuer's loading. Equal-weighted uses an equal-weighted sovereign bond market factor. VW foreign and EW foreign use value-weighted and equal-weighted versions of the foreign sovereign-bond factor (excluding colonial bonds). LOO foreign and LOO colonial apply the leave-one-out construction within the foreign-only and colonial-only bond panels respectively, with the full panel-specific factor used for issuers not in that panel. HI subsample uses a sovereign factor built from the upper-tercile composite-institution issuers in the IV sample (Australia, Canada, Mauritius, New Zealand, Newfoundland, Sri Lanka, Trinidad, Uruguay, and the United States), with leave-one-out applied to those issuers. The HI-subsample construction is the most stringent reflexivity test: the factor's construction does not depend on the bottom two-thirds of the institutional cross-section being tested. Panel A uses the OLS sample with nonmissing institutional composite and gold-standard adherence. Panel B uses the IV sample. The institutional composite enters as a country-level median for 1870–1907, then min–max rescaled across issuers to the unit interval; the OLS coefficients therefore measure the difference between hypothetical issuers at the cross-sectional extremes. Heteroskedasticity-robust standard errors are reported in parentheses, using the HC3 leverage adjustment of MacKinnon and White (1985). Coefficient magnitudes are not directly comparable across factor constructions because the alternative sovereign factors have different return scales. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

A.8 Sovereign factor left-tail risk robustness

Appendix Table A.9 reports compact robustness checks for the sovereign-factor left-tail diagnostic. The first three rows compare the annual default share with contemporaneous, one-year-lagged, and two-year-lagged measures of sovereign-factor downside risk. The lagged correlations are similar to or larger than the contemporaneous correlations, consistent with the default-share series measuring the stock of countries in default rather than only new default onsets. The final row excludes the 1875–1878 default wave. The relation weakens substantially, confirming that the analysis should be read as a historical diagnostic rather than as evidence of a stable annual mapping from default shares to factor drawdowns.

Table A.9: Default stress and sovereign-factor tail risk: timing and episode sensitivity

	MDD	Downside	Bad10	Roll13 MDD
Contemporaneous	0.274	0.370	0.406	0.342
Lagged one year	0.323	0.409	0.430	0.371
Lagged two years	0.301	0.388	0.454	0.415
Excluding 1875–1878	-0.144	-0.084	-0.065	-0.134

Notes: The table reports correlations between the annual share of sovereigns in default and annual sovereign-factor tail-risk measures. The contemporaneous row correlates default share in year y with the tail-risk measure in year y. The lagged rows correlate default share in year y with sovereign-factor tail risk measured in years $y-1$ and $y-2$. The final row reports the contemporaneous correlation after excluding the 1875–1878 default wave. MDD is the within-year maximum drawdown; Downside is the cumulative downside return; Bad10 is the number of four-week sovereign-factor returns below the unconditional 10th percentile; and Roll13 MDD is the maximum trailing 13-period drawdown observed within the year.

A.9 Haircut record for OLS-only issuers

Appendix Table A.10 cross-checks the OLS-only issuer set against the historical haircut archive of Meyer et al. (2022), supporting the discussion of the level-channel attenuation in Section 6.2. The OLS-only set comprises the 21 issuers in the OLS sample but outside the IV sample; of these, six experienced 10 external debt restructurings with default dates in 1870–1907. The remaining 15 had no in-window event, either because they never defaulted on external private bonds (7 sovereigns) or because their first default falls after 1907 (8 sovereigns). The institutional composite for the six OLS-only defaulters has the same mean (0.37) as for the IV defaulter group (0.36), so the level-channel attenuation in the OLS sample reflects the addition of 15 zero-loss observations rather than a shift in within-defaulter institutional composition.

Table A.10: Meyer-Reinhart-Trebesch (2022) haircut record for OLS-only issuers, 1870–1907

Country	Default	Agreement	PV haircut (%)	FV haircut (%)	Debt ($m)
Dominican Republic†	Jun 1872	1872	100.0	100.0	137
Dominican Republic	Dec 1892	1893	54.0	0.0	156
Dominican Republic	Dec 1896	1897	32.4	0.0	308
Dominican Republic	Apr 1899	1908	32.4	50.0	386
Greece	Dec 1893	1898	60.6	11.4	2,614
Liberia	Aug 1874	1898	86.2	0.0	29
Portugal	Jan 1892	1902	54.0	50.4	1,351
Spain	Jun 1872	1882	38.5	54.5	17,989
Turkey	Jan 1876	1881	91.7	58.3	19,838
Turkey	Dec 1903	1903	29.5	59.4	6,668
Mean / Total			57.9	38.4	49,476

Notes: External sovereign debt restructurings with default dates in 1870–1907 for the six OLS-only issuers (out of 21) that appear in the haircut archive of Meyer et al. (2022) with at least one in-window event. "Default" is the start month of the default spell; "Agreement" is the year the combined restructuring deal was reached, which can post-date the sample window. "PV haircut" is the present-value creditor loss following Sturzenegger and Zettelmeyer (2006), MRT's preferred measure; "FV haircut" is the face-value reduction. "Debt ($m)" is the par amount restructured in millions of 2009 USD, deflated by MRT. The remaining 15 OLS-only issuers have no in-window event in the haircut archive: 7 never defaulted on external private bonds (Belgium, Denmark, Fiji, France, Netherlands, Norway, Sweden) and 8 defaulted only after 1907 (Austria-Hungary, Bulgaria, China, Germany, Italy, Japan, Romania, Russia). † Full repudiation. Source: Meyer et al. (2022) replication archive (final-sample haircut file).