Financial Sanctions and Dollar Dominance in Trade Invoicing*

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Abstract

This paper documents signs that the long-standing dominance of the U.S. dollar in global trade invoicing is beginning to falter, a development associated with U.S. financial sanctions. Theoretically, I propose a novel mechanism linking invoicing currency choice to sanctions: invoicing in dollars exposes firms to uncertainty about cash flows under sanctions, which can lead creditors to limit working-capital lending. Firms balance this financing risk against the benefits of dollar invoicing in terms of liquidity and reduced exchange-rate risk. Using country-level import invoicing data, I find a small but measurable shift away from dollar invoicing. Stronger U.S. sanctions against Chinese entities are associated with larger declines in dollar invoicing shares and larger increases in renminbi invoicing shares in countries with greater pre-existing import exposure to China. For a 10-percentage-point higher initial share of imports from China, a one-standard-deviation increase in sanctions is associated with a 0.8-percentage-point larger decline in the dollar invoicing share and a 0.2-percentage-point larger rise in the renminbi invoicing share on impact. Consistent with the mechanism, suggestive evidence indicates that the renminbi response is weaker where imports from China are concentrated in industries with lower working-capital needs. These findings suggest that sanctions can weaken dollar dominance by making dollar invoicing a source of financing risk.

Keywords: Invoicing Currency; Financial Sanctions; Dollar Dominance; RMB Internationalization; Working Capital Finance.
JEL Codes: F33; F41; F51; G32.

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1 Introduction

A central question in dollar dominance is no longer only why firms use the dollar, but when they choose to leave it. International trade is increasingly multipolar, yet invoicing and trade finance remain heavily dollar-centered. The standard explanation is that the dollar network is privately valuable: it provides liquidity, hedging opportunities, and strategic complementarities, and it shapes international transmission of shocks, exchange-rate pass-through, trade adjustment, and firms' financing conditions.[1] Recent geopolitical fragmentation, however, has made visible a cost that is largely absent from normal-time theories of dollar dominance: the same network that creates liquidity also creates jurisdictional exposure. This paper examines how U.S. financial sanctions against Chinese entities relate to currency choice along trade links to China. I show that this financial fragmentation does not produce broad de-dollarization. Instead, they induce "*selective*" switching from the U.S. dollar (USD) to Chinese renminbi (RMB) invoicing in financially vulnerable supply-chain links.

Financial sanctions matter for trade because they operate through the same private intermediaries that make the dollar network valuable. A U.S. designation can freeze assets under U.S. jurisdiction, prohibit transactions with targeted entities, and restrict access to dollar clearing. Even when a transaction is not itself prohibited, banks and correspondent institutions may respond by tightening compliance screening, delaying or rejecting payments, reducing credit limits, or withdrawing from relationships that create sanctions exposure. These de-risking responses propagate along supply chains because trade payments, receivables, and short-term trade finance are jointly processed through the banking system. For firms relying on dollar settlement, sanctions can therefore affect the reliability of dollar cash flows as a source of working-capital finance.[2] Choosing the dollar inherently means choosing exposure to the enforcement network embedded in dollar clearing and credit infrastructure.

This financing channel motivates examining two predetermined characteristics of each destination: its import exposure to China and the working-capital intensity (WCI) of its import basket from China. The first measures how strongly a country is connected to Chinese supply chains. The second summarizes how costly a disruption to dollar-linked working-capital finance would be for the goods it imports from China. Together, they provide a way to examine whether the relationship between sanc-

[1] For canonical contributions, see Corsetti et al. (2010) and Gopinath and Itskhoki (2022).

[2] Although invoicing and settlement currencies are conceptually distinct, they are closely aligned in many trade transactions (Friberg and Wilander, 2008; Boz et al., 2022). I therefore interpret invoicing currency as the observed contractual margin that proxies for exposure to currency-specific payment, settlement, and trade-finance frictions.

tions and currency choice varies with the financing needs of the goods traded.

I examine these relationships using a country-year panel of FX-neutral import invoicing shares from 2012 to 2023. The empirical challenge is that U.S. sanctions against Chinese entities are a common global shock across destinations, so I interact time-varying U.S. financial sanctions against Chinese entities with each country's pre-sanctions import exposure to China fixed in 2010. I then allow the exposure gradient to differ between destinations with working-capital intensity of the country's predetermined import basket from China. The use of FX-neutral invoicing shares is critical, as it removes mechanical valuation effects from exchange-rate movements, ensuring the estimates capture active changes in currency choice rather than passive changes in the dollar value of a fixed basket.

The baseline estimates indicate economically meaningful differences in sanctions responses across destinations. In the fully controlled specification, a one-standard-deviation increase in the log sanctions index is associated with a contemporaneous RMB response that is about 0.19 percentage points (pp) higher and a USD response that is about 0.84 pp lower for a 10 pp higher baseline import share from China.[3] The heterogeneity specification estimates a negative RMB exposure interaction for destinations in the bottom decile of working-capital intensity, indicating a flatter exposure gradient in this group. This describes how the sanctions response varies with China exposure. Accordingly, the low-WCI result remains compatible with a nonzero RMB response. These patterns are consistent with a role for financing requirements in currency choice.

To interpret these patterns, I introduce a geopolitical risk-management motive into the invoicing-currency problem. In standard models, firms choose the currency that best replicates flexible prices under normal-time macroeconomic volatility. In my model, I relax the implicit political neutrality of the currency network. Chinese exporters jointly choose an invoicing currency and a preset price, balancing the dollar's normal-time financing and hedging benefits against sanctions-sensitive credit access. Greater sanctions intensity raises the probability of a stress state in which usable credit under dollar invoicing is impaired. RMB invoicing sacrifices the dollar's normal-time advantages but insulates the firm from this dollar-specific credit impairment.

For exporters that prefer dollar invoicing at low risk and RMB invoicing at sufficiently high risk, the model yields a switching threshold logic. Working-capital dependence lowers this threshold when

[3] These magnitudes use $\widehat{\theta}^{(c)} \times 10 \times 1.36$, with the coefficients in columns (3) and (6) of Table 1 and the sanctions-index change used in Section 3.

it increases the cost of impaired credit more than the normal-time dollar advantage. Dollar-priced inputs raise the threshold when their contribution to the normal-time hedging advantage exceeds their effect on the cost of sanctions risk. Aggregation uses baseline trade-value weights to connect these firm-level choices to destination-level invoicing shares. China exposure scales the affected trade, while the concentration of trade at thresholds near the prevailing risk level determines the local response to a further risk increase. Lower thresholds imply earlier RMB adoption and the size of a subsequent response depends on how much trade remains near the switching margin. This distinction connects the conditional firm-level mechanism to the aggregate heterogeneity examined in the empirical analysis.

The quantitative implementation specifies a trade-value-weighted distribution of switching thresholds, generates RMB invoicing shares, and applies the same regression used in the empirical analysis. At the reference probability normalization and threshold dispersion, removing the working-capital loading prevents the model from reproducing the negative low-WCI exposure differential, including after re-fitting the remaining parameters to the same coefficient targets. This comparison describes the contribution of working-capital heterogeneity within the reference specification. A separate retrospective exercise calibrates the WCI and no-WCI models on 2012–2019 observations and evaluates their 2020–2023 paths conditional on realized sanctions. Under that design, including WCI modestly reduces both level root mean square error (RMSE) and RMSE after removing country and year effects, relative to the re-fitted no-WCI model and the initial-share benchmark.

Related Literature. —This paper contributes most directly to the literature on endogenous invoicing-currency choice by showing that currency choice is a pricing and hedging decision and also a choice over exposure to currency-specific payment and credit infrastructure. Classic theory views the invoicing currency as a device for managing normal-time macroeconomic risk: firms choose the currency that best replicates the flexible-price optimum they would set under the real-hedging motive (Devereux et al., 2004; Engel, 2006; Gopinath et al., 2010).[4] A more recent literature extends this real-hedging logic to dominant-currency pricing (DCP) through strategic complementarity in market structure (Bacchetta and Van Wincoop, 2005; Goldberg and Tille, 2008; Devereux et al., 2017), input-currency matching

[4]Prior to the development of endogenous currency-choice frameworks, the invoicing currency was treated as an exogenous. Classic exchange-rate overshooting and dynamic general-equilibrium models (Dornbusch, 1976; Obstfeld and Rogoff, 1995) typically relied on producer-currency pricing (PCP), while studies emphasizing pricing-to-market and incomplete pass-through (Krugman, 1986; Dornbusch, 1987) assumed local-currency pricing (LCP).

through intermediate sourcing (Novy, 2006; Devereux et al., 2010; Chung, 2016; Gopinath et al., 2020; Mukhin, 2022), and a trade-finance channel (Gopinath and Stein, 2021). These theoretical models are supported by a large empirical literature using transaction-, firm-, and country-level data on invoicing currency choice (Donnenfeld and Haug, 2003; Kamps, 2006; Goldberg and Tille, 2016; Devereux et al., 2017; Licandro and Mello, 2019; Amiti et al., 2022; Corsetti et al., 2022; Chowdhry, 2024; Yoshida et al., 2024; Benguria and Wagner, 2024). This literature emphasizes the factors that explain the dollar's dominant invoicing status and its persistence for roughly a century, while the forces that could let a new currency rise, or threaten the dollar's dominance, have received comparatively little attention (Arslanalp et al., 2022). As the dollar displaced sterling in the interwar period once dollar financial infrastructure developed (Chiţu et al., 2014), the recent build-out of financial infrastructure together with geopolitical fragmentation has renewed attention to new currencies and payment systems.

This paper also relates to a rapidly growing body of work on the effects of resurging geopolitical fragmentation on the global financial system. Recent theoretical work formalizes the tension between a hegemonic currency issuer's ability to weaponize its central position (Clayton et al., 2026a,b) and the endogenous emergence of alternative payment systems that blunt this power (Cipriani et al., 2023; Clayton et al., 2025). Through complementarities across financial systems, geopolitical fragmentation is associated with a declining dollar role in trade invoicing and borrowing (Berthou, 2023; Boz et al., 2025; Corsetti et al., 2025; Chupilkin et al., 2025; Garofalo et al., 2025), in central bank reserves (Bianchi and Sosa-Padilla, 2025; Sanusi et al., 2025), and in cross-border lending (Efing et al., 2023; Niepmann and Shen, 2025; Pradhan et al., 2025; Reinhardt et al., 2025), and with adjustments in exchange rates (Itskhoki and Mukhin, 2025; Lorenzoni and Werning, 2023).

A primary contribution of this paper is to connect country-level evidence on sanctions and invoicing with a model of exporters' financing decisions. The model builds on the three-period working-capital framework of Bahaj and Reis (2026) and their complementarity between the currency of working capital and the currency of invoicing; they study a "pull" force, central-bank swap lines that reshape the distribution of RMB funding costs. Corsetti et al. (2025) study a closely related "push" away from the dollar, operating through reputational and secondary-sanction risk and the cost of dollar payment circuits when exporting to a sanctioned destination. I examine a complementary dollar push factor operating through Chinese exporters' access to working-capital finance. In the model, usable credit depends on the invoicing currency and the realized geopolitical state, and firms take this dependence

into account when choosing currency and price. This interpretation is consistent with evidence that the currency composition of credit responds to sanctions (Garofalo et al., 2025). Building on the literature on earnings-based borrowing constraints (Lian and Ma, 2021; Drechsel, 2023; Caglio et al., 2021), the model makes currency-specific credit impairment part of the invoicing decision. Currency choice therefore determines both the denomination of revenues and exposure to a particular source of funding risk.

Finally, this paper speaks to the literature on RMB internationalization. Despite significant policy efforts by the Chinese government, the RMB's global invoicing share remains low, with highly heterogeneous adoption across countries. Existing research highlights that central bank swap lines and offshore clearing banks are critical factors explaining this cross-country variation (Song and Xia, 2020; Perez-Saiz and Zhang, 2023; Han et al., 2026). For example, Benguria and Novy (2025) document using Argentine firm-level data that RMB import invoicing surged toward fifty percent when a bilateral swap line relieved a local dollar shortage, before receding as the constraint eased. Using broader SWIFT data, Bahaj and Reis (2026) find that swap lines increase both the probability and the intensity of RMB usage. Regional economic integration and geopolitical alignment also shape adoption patterns (Koosakul et al., 2024), with RMB usage accelerating most prominently in geoeconomically fragmented regions. This paper complements these pull factors by examining a potential push factor associated with U.S. sanctions against Chinese entities. The empirical exposure relationship remains after controlling for observed swap-line access, RMB clearing banks, and trade agreements, while the model explains how less reliable dollar credit can strengthen the incentive to use RMB.

Roadmap. The remainder of the paper is organized as follows. Section 2 describes the data construction, including the FX-neutral invoicing shares, U.S. sanctions data, measures of China trade exposure, and working-capital dependence. Section 3 presents the empirical framework and main results, documenting how the relationship between China-directed sanctions and invoicing shares varies with baseline trade exposure and working-capital intensity. Section 4 develops a theoretical model of invoicing currency choice under sanctions-sensitive working-capital constraints. Section 5 implements the threshold mechanism quantitatively and compares specifications with and without working-capital heterogeneity. Section 6 concludes.

2 Data

The empirical analysis uses an unbalanced annual panel of 88 destination countries over 2012–2023. The outcome variables are country-level import invoicing shares adjusted for contemporaneous exchange-rate valuation effects. I merge these outcomes with three constructed measures: (i) annual counts of new U.S. Specially Designated Nationals and Blocked Persons (SDN) designations associated with China; (ii) the share of each destination country's total imports accounted for by imports from China in 2010, which measures predetermined trade exposure to China; and (iii) the working-capital intensity of the country's 2010 import basket from China, which captures predetermined heterogeneity in exposure to the proposed financing channel. The empirical design combines the common time-varying sanctions measure with these predetermined cross-country exposures.

Global Invoicing Data. I obtain annual country-level invoicing shares from the Global Currencies Database developed by Boz et al. (2025), which extends the dataset of Boz et al. (2022) by expanding country coverage and adding invoicing shares for the Chinese renminbi (RMB).[5] The underlying series were collected from national authorities through data requests conducted by the International Monetary Fund (IMF) and the European Central Bank (ECB). The full database covers 106 economies at an annual frequency over 1995–2024, although coverage varies across countries, years, trade flows, and invoicing currencies. The empirical analysis uses import invoicing observations from 2012 to 2023. Iran, Taiwan, and Hong Kong are not covered, while Russia is observed only through 2021. These coverage limitations reduce the representation of economies with particularly close trade.

Because the reported invoicing shares are value-based, exchange-rate movements can mechanically change a currency's recorded share even when the underlying currency composition of trade remains unchanged. For example, an appreciation of the RMB increases the USD value of a RMB-denominated trade flow and can therefore raise the recorded RMB share even if the underlying currency composition of trade does not change. To reduce this mechanical valuation component, I revalue currency-specific invoiced trade using fixed exchange rates and refer to the resulting measure as the FX-neutral invoicing share.

[5]The data are publicly available from the Chițu–Eichengreen–Mehl Global Currencies Database. I use the version downloaded on September 22, 2025.

Let $s^c_{i,t}$ denote the raw import invoicing share of currency c in country i and year t:

$$s^c_{i,t} \equiv \frac{\text{Value of imports invoiced in currency } c}{\text{Total import value}},$$

where both values refer to country i's imports from all origins in year t and are measured in current U.S. dollars. The annual average nominal exchange rate, $\mathcal{E}_{c,t}$, is expressed in units of currency c per USD.[6] Under this convention, an increase in $\mathcal{E}_{c,t}$ represents a depreciation of currency c against the USD. Holding exchange rates fixed at their base-year values, the FX-neutral invoicing share is defined as

$$\tilde{s}^c_{i,t} = \frac{s^c_{i,t}\left(\mathcal{E}_{c,t}/\mathcal{E}_{c,b}\right)}{\sum_{k\in\mathcal{C}_{i,t}} s^k_{i,t}\left(\mathcal{E}_{k,t}/\mathcal{E}_{k,b}\right)}. \tag{1}$$

Here, b denotes the base year. For country i in year t, $\mathcal{C}_{i,t}$ comprises the USD, the euro, the RMB, and a residual group containing all remaining currencies used to invoice its imports. Because the currency composition of the residual group is unobserved, I approximate its valuation adjustment using country i's home-currency exchange rate. The baseline specification sets $b = 2019$. For the corrected import-share construction, Table A.1 compares the 2019-based measure with alternatives using 2010 exchange rates and the arithmetic averages of annual exchange rates over 2007–2010 and 2012–2023. The table reports correlations in levels and after removing country and year fixed effects, using a common country–year sample for each currency. The 2012–2023 average yields measures very similar to the 2019 benchmark.

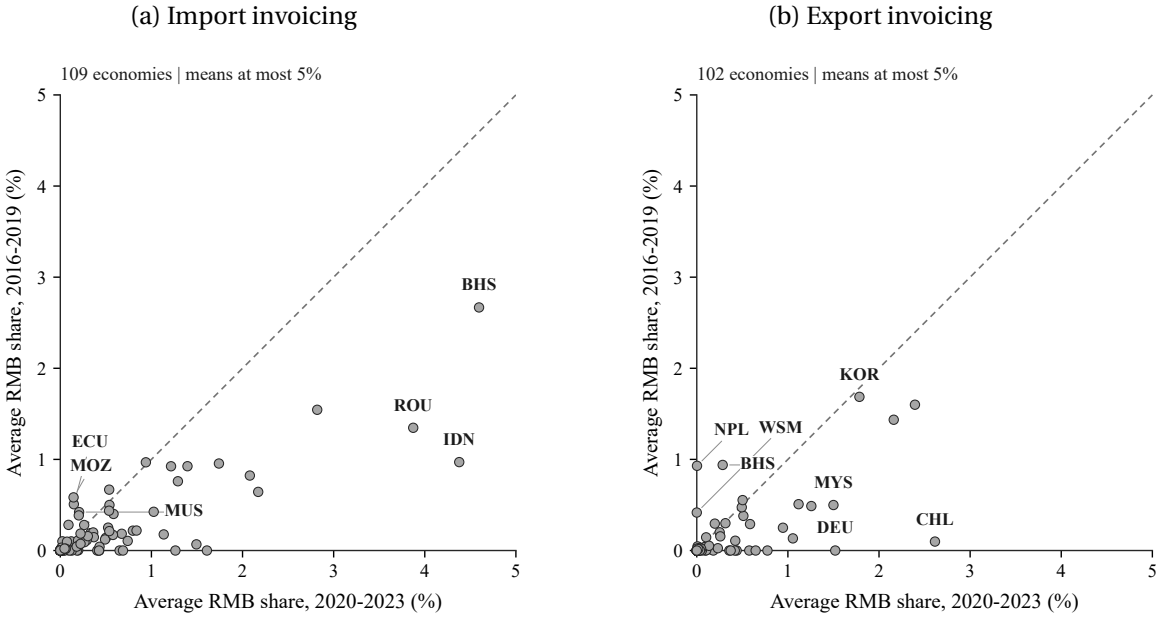
Table A.2 compares raw and FX-neutral invoicing shares over 2012–2023 using a common country–year sample within each trade flow. The USD accounts for the largest average share among the reported currencies: its unweighted average FX-neutral share is 51.42% for imports and 52.58% for exports. These descriptive averages are consistent with the prominent role of dollar invoicing documented in the dominant currency literature (Gopinath et al., 2020; Mukhin, 2022; Boz et al., 2025). The average RMB share remains below 1% in both flows, at 0.45% for imports and 0.33% for exports. Figure A.1 provides a complementary cross-country comparison of import invoicing shares and bilateral trade exposure in 2023.

Although the average RMB invoicing share remains small for most economies, Figure 1 shows considerable cross-country variation and an increase between 2016–2019 and 2020–2023 for a majority

[6] Annual average exchange rates are obtained from the IMF's International Financial Statistics (IFS).

of economies. Among the economies displayed in the figure, the average RMB import invoicing share increases in 71 of 109 economies, while the export share increases in 60 of 102 economies. In the full descriptive sample reported in Figure A.2, the corresponding counts are 73 of 111 economies for imports and 61 of 103 economies for exports. These averages use available observations within each period, so the years entering the comparison can differ across economies.

Figure 1: Changes in FX-neutral RMB invoicing shares



Notes: The figure compares country-level average RMB invoicing shares in 2016–2019 and 2020–2023. Panel a reports import invoicing shares, and Panel b reports export invoicing shares. Each point represents an economy with at least one nonmissing observation in both periods. The invoicing shares are adjusted using 2019 exchange rates as described in equation (1). The dashed 45-degree line represents no change between the two periods; observations below the line experienced an increase in their average RMB invoicing share. Selected economies are labeled by their ISO country codes. To preserve the scale of the figure, Mongolia and the Russia are omitted from the import panel, while Mongolia is omitted from the export panel. Figure A.2 reports the corresponding figures using the full sample. The underlying invoicing data are from Boz et al. (2025).

U.S. Financial Sanctions. I use the SDN data published by the U.S. Treasury's Office of Foreign Assets Control (OFAC) to construct two distinct measures of sanctions intensity.[7] The distinction reflects the empirical setting: the model centers on Chinese exporters' invoicing decisions, whereas the country-level outcome records the currencies used to invoice destination country i's imports. The first measure is *China-directed sanctions intensity*, defined from new SDN designations assigned to China.

[7] Ahn and Ludema (2020) use U.S. and EU sanctions records, including SDN designations, to study firms exposed to Russia-related sanctions. For China, Dou et al. (2026) examine financing constraints using a broader set of U.S. sanctions lists that includes SDN. Laudati and Pesaran (2023) construct an alternative sanctions-intensity measure from U.S. Treasury designation and removal records for Iran, alongside their main newspaper-based measure.

It captures the common exporter-side pressure central to the paper. The second is *destination-directed sanctions intensity*, defined from new SDN designations assigned to destination country i. Such designations may disrupt dollar payment and trade-finance conditions in the destination market, thereby changing the relative attractiveness of USD versus RMB invoicing to Chinese exporters. I include this measure separately so that the coefficient on China-directed sanctions exposure is not conflated with sanctions pressure operating through the destination country itself.

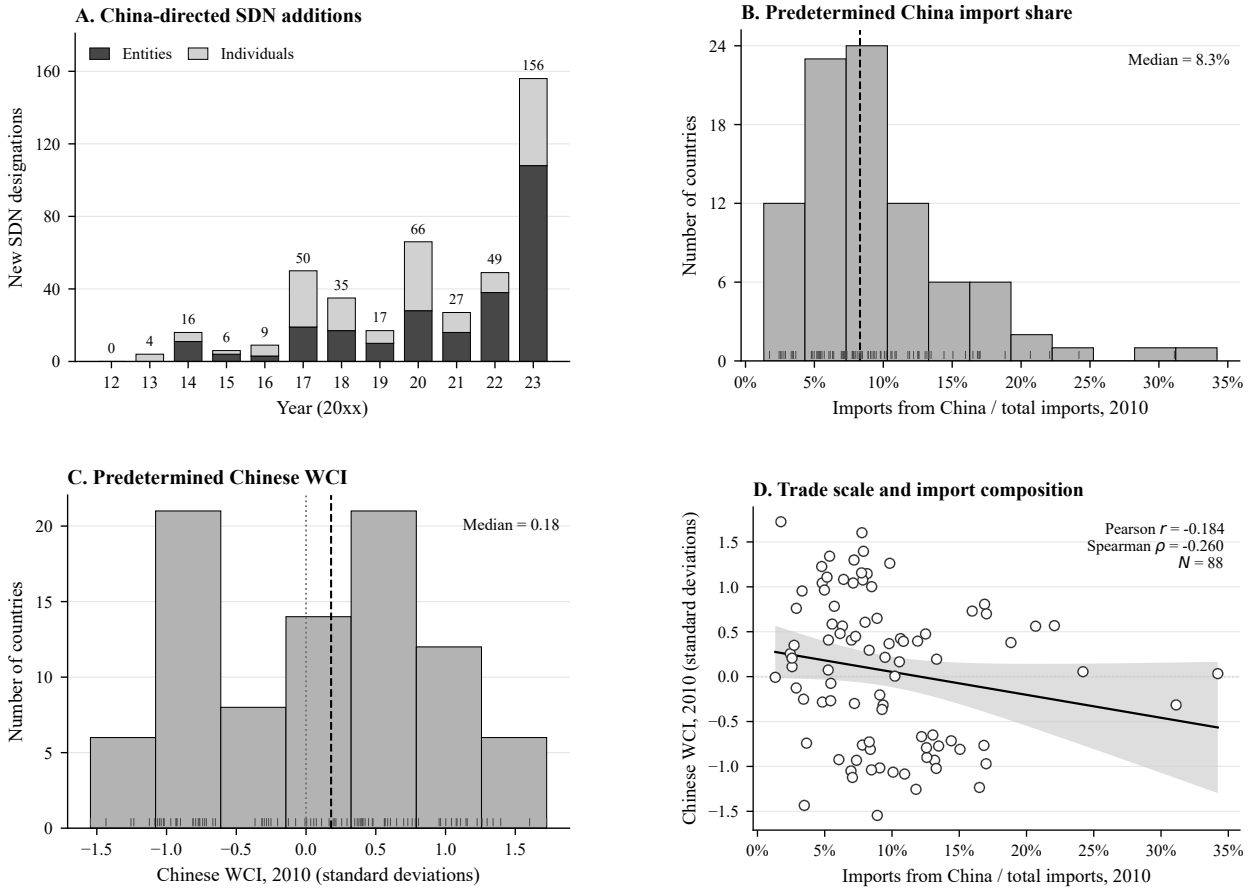
An SDN designation is a blocking sanction. It requires U.S. persons to freeze any property and interests in property of the designee that come within U.S. jurisdiction or their possession or control, and generally prohibits them from dealing with the designee. These restrictions extend to financial intermediation. U.S. financial institutions cannot open accounts for an SDN and must block transfers in which the designee has an interest. Designation therefore impairs the target's access to U.S.-linked dollar clearing and associated cross-border payment services. More broadly, restrictions on access to the institutions and infrastructure supporting international payments have become an increasingly prominent form of financial sanctions. Because key networks such as SWIFT have few close substitutes, these restrictions can disrupt a broad range of cross-border activities, including trade financing (Cipriani et al., 2023). This common combination of asset blocking and transaction prohibitions provides a relatively consistent treatment margin with a direct nexus to U.S. financial jurisdiction. By contrast, OFAC's Consolidated Non-SDN files combine several lists whose restrictions differ across programs and are not uniformly blocking. I exclude those records from the baseline measure to preserve this common treatment margin.

The raw OFAC snapshot provides the publication date, entity identifier, entity type, sanctions-program tags, country, citizenship, and location associated with each listed record. I use the publication date to assign each record to a month and retain records classified by OFAC as either an Individual or an Entity. Vessels and aircraft are excluded because they identify blocked property rather than counterparties and cannot be assigned to countries using a comparable nationality-based rule. I assign each record to its directly reported country; when that field is missing, I use citizenship and then location. Within the 2012–2023 analysis window, 683 of $10,309$ individual and entity records (6.6 percent) lack usable information in all three geographic fields and are therefore excluded. Country names are subsequently harmonized and mapped to ISO-3 codes. To prevent duplicate counting, I retain one observation per identifier, publication month, assigned country, and person type. I aggregate

the cleaned records to the country-month level, fill zero-designation months with zeros, and sum the monthly counts to obtain annual flows of new SDN designations.

Let S_t^{CN} denote the annual number of new SDN records assigned to China, and let $S_{i,t}^{\mathrm{D}}$ denote the corresponding count assigned to destination country i. I transform these counts as $g_t = \ln(1 + S_t^{\mathrm{CN}})$ and $d_{i,t} = \ln(1+S_{i,t}^{\mathrm{D}})$. Panel a of Figure 2 displays the annual China-directed SDN additions underlying S_t^{CN}, separately for entities and individuals from 2012 to 2023. The series provides the common time-series variation in sanctions intensity used in the exposure design.

Figure 2: Sources of Identifying Variation



Notes: Panel A reports annual additions to the U.S. SDN list directed at China between 2012 and 2023, separately for entities and individuals. Panels B–D use the 88 destination countries included in the fully controlled RMB specification. Panel B shows the cross-country distribution of the share of imports from China in 2010. Panel C shows the distribution of the standardized Chinese working-capital-intensity measure. Panel D plots the two predetermined exposure measures and reports their cross-country association (Pearson $r = -0.184$, Spearman $\rho = -0.260$, and $N = 88$); the fitted line is from a bivariate OLS regression, with the shaded area denoting its 95 percent confidence interval. The corresponding R^2 is 0.034. The weak linear association is consistent with the China import share primarily capturing the scale of bilateral trade and the WCI capturing the industry composition of Chinese imports.

China Import Exposure. I construct the baseline China import share from annual bilateral trade records in UN Comtrade. For destination country i, the measure is

$$E_{i,2010} = \frac{M_{i,\text{CN},2010}}{M_{i,\text{World},2010}}, \tag{2}$$

where $M_{i,\text{CN},2010}$ is the value of country i's imports reported from China in 2010 and $M_{i,\text{World},2010}$ is its total imports from all origins in the same year. The measure is expressed as a percentage and is fixed at its 2010 value. Panel b of Figure 2 shows its distribution across the 88-country matched analytical sample.

Working Capital Intensity and Exposure. To examine whether the invoicing response to sanctions varies with short-term liquidity needs before sales are collected, I construct a new, predetermined measure of the working-capital intensity embedded in each destination country's imports from China.[8] Because the relevant suppliers are Chinese firms, I construct the industry ratios from Chinese A-share listed firms' balance sheets and income statements. This allows the measure to reflect payment terms and supplier and customer financing arrangements observed in China, without assuming that industry rankings estimated from U.S. firms carry over to China. The resulting measure remains a proxy based on listed firms' observed operating-finance practices rather than a direct measure of the financing constraint faced by the marginal exporter.

The construction proceeds in three steps. First, I obtain balance-sheet and income-statement data for Chinese A-share listed firms from Eastmoney through AKShare (King, 2022). For firm f, net operating working capital (NOWC) is accounts receivable plus inventories, less accounts payable and advances from customers. Dividing this amount by sales gives

$$w_{f,t} = \frac{\text{Receivables}_{f,t} + \text{Inventories}_{f,t} - \text{Payables}_{f,t} - \text{Customer Advances}_{f,t}}{\text{Sales}_{f,t}}. \tag{3}$$

where the baseline retains firms listed by the end of fiscal year $t = 2010$ with positive sales and non-missing values for all four NOWC components.

The sign of $w_{f,2010}$ has a direct operating-finance interpretation. When $w_{f,2010} > 0$, the funds tied

[8] For sectoral financial measures constructed from U.S. firms, see Manova et al. (2015). Working-capital management using Chinese listed-company data has also been studied by He et al. (2017). The measure developed here combines Chinese industry working-capital intensity with destination-specific import composition.

up in receivables and inventories exceed the operating finance supplied through accounts payable and customer advances. The firm must finance this gap using internal liquidity or external sources such as bank credit and trade finance. A positive ratio therefore indicates an operating funding requirement, not necessarily realized borrowing or a binding financial constraint. Conversely, when $w_{f,2010} < 0$, financing provided by suppliers and customers exceeds the funds tied up in receivables and inventories, so trading partners finance the firm's operating cycle on net. A negative ratio may reflect favorable payment terms, customer prepayments, or supply-chain bargaining power.

Second, I assign firms to 2-digit ISIC Revision 3.1 industries using the official China Securities Regulatory Commission (CSRC) 2012Q4 listed-company assignments and a CSRC-to-ISIC concordance that I construct. This concordance places the firm accounts and bilateral trade data in a common industry classification. This is a retrospective taxonomy applied to fiscal-year 2010 accounts. The broad CSRC C39 (Computers, Communication Equipment, and Other Electronic Equipment Manufacturing) does not have a unique ISIC counterpart. I use the more detailed legacy CSRC code to form an initial assignment and then review the principal activity of 60 ambiguous firms. For industry s, the baseline working-capital intensity is the revenue-weighted mean of the winsorized firm ratios:

$$\text{NOWC}_{s,2010} = \frac{\sum_{f \in s} \text{Sales}_{f,2010}\, \widetilde{w}_{f,2010}}{\sum_{f \in s} \text{Sales}_{f,2010}}, \tag{4}$$

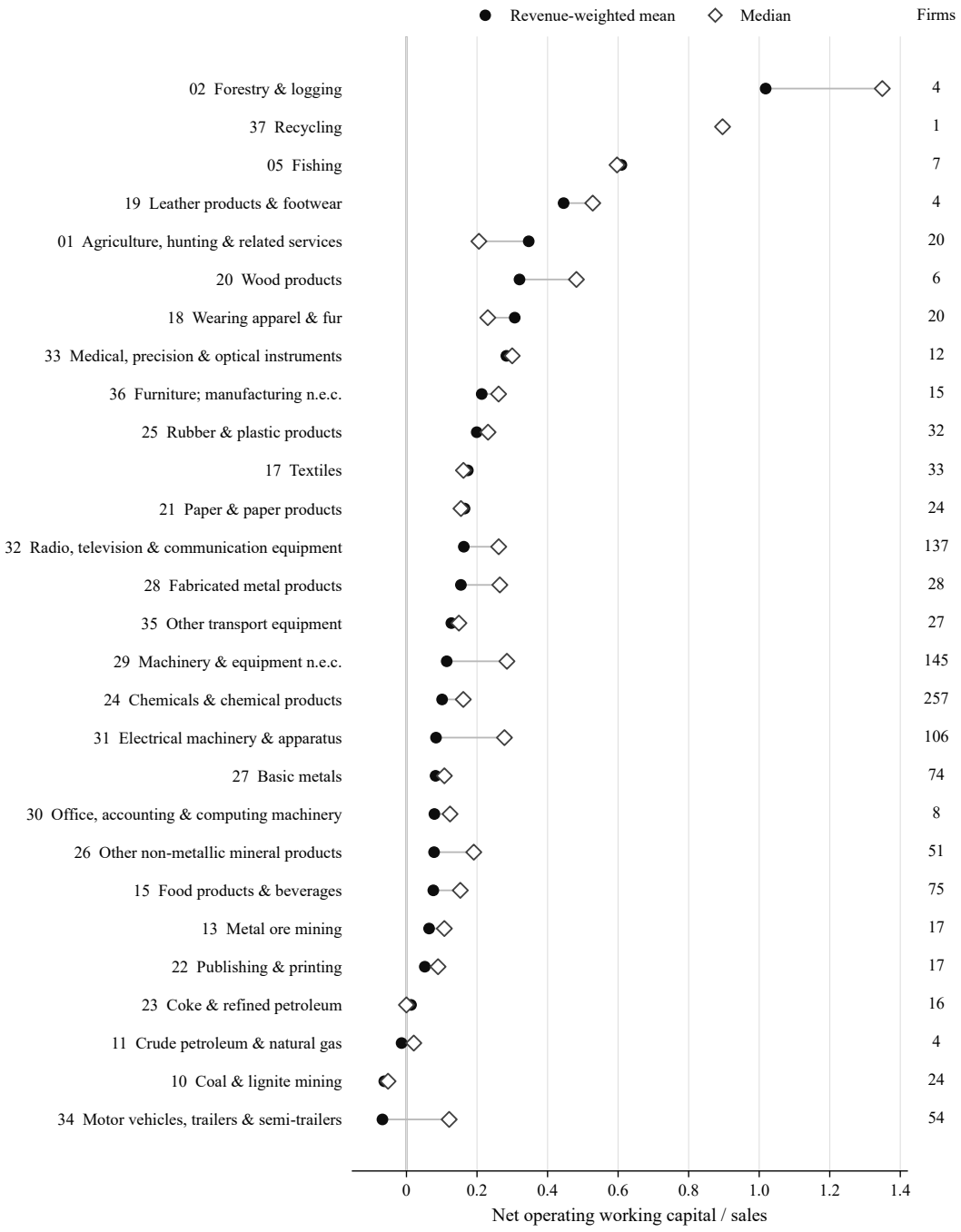
where $\widetilde{w}_{f,2010}$ denotes the winsorized firm ratio at the 1st and 99th percentiles of the ISIC goods-sector firm distribution, preventing service sectors from determining the goods-sector cutoffs.[9] The revenue weights make the statistic representative of listed-firm sales within an industry, while the industry median shown in Figure 3 documents sensitivity to large firms.

Figure 3 shows substantial cross-industry dispersion. To assess the sensitivity of measured industry working-capital requirements, I compare the baseline NOWC-to-sales measure with an inventory-to-sales proxy and alternative choices of aggregation, accounting period, sample composition, and outlier treatment. The inventory-to-sales proxy follows the intuition of Manova et al. (2015). Table A.5 defines these alternatives and reports the number of available industries and their Pearson and Spearman correlations with the baseline. The comparisons indicate similar cross-industry patterns across

[9]The resulting baseline contains $1,214$ firms in 28 ISIC divisions. The underlying Eastmoney endpoint provides current-vintage values for historical reporting periods, which may incorporate subsequent restatements, rather than an archive of values exactly as reported in 2011. De-listed firms are also underrepresented.

the alternative measures.

Figure 3: Net Operating Working Capital across 2-Digit ISIC Industries



Notes: This figure reports cross-industry variation in net operating working capital at the 2-digit ISIC (Revision 3.1) level. The ratios are constructed from financial statements of Chinese A-share listed firms in 2010. Filled circles report revenue-weighted industry means of firm-level NOWC-to-sales ratios winsorized using the pooled 1st- and 99th-percentile cutoffs across all eligible goods-sector firms in 2010; open diamonds report medians of the corresponding raw ratios. The cutoffs are common across industries. Industries are ordered by the revenue-weighted mean. The rightmost column reports the number of firms.

Finally, I map each destination country's 2010 HS6 imports from China into 2-digit ISIC Revision

3.1 industries and combine the resulting import shares with Equation 4. Let $M^{\text{CN}}_{i,s,2010}$ denote country i's imports from China in industry s, and let $\mathcal{S}_i$ denote the mapped goods industries for which the final Chinese WCI is available. I normalize the weights over this covered basket,

$$\text{WC}_{i,2010} = \sum_{s \in \mathcal{S}_i} \omega_{i,s,2010}\, \text{NOWC}_{s,2010}, \qquad \omega_{i,s,2010} = \frac{M^{\text{CN}}_{i,s,2010}}{\sum_{r \in \mathcal{S}_i} M^{\text{CN}}_{i,r,2010}}. \tag{5}$$

Thus, $\text{WC}_{i,2010}$ measures the average working-capital intensity of country i's 2010 imports from China, weighting each covered industry by its share in that country's covered Chinese import basket. A higher value indicates that these imports are concentrated in industries with greater net operating working-capital requirements relative to sales. China's share of country i's total imports separately measures its overall trade exposure to China.

Coverage of the industry measure is high. ISIC divisions 12, 14, and 16 lack an eligible 2010 Chinese WCI observation, but together account for only 0.210% of import value in the crosswalk-eligible analytical trade sample. The median destination-country unmatched share is 0.085%.[10] To strictly isolate the impact of U.S. financial sanctions or invoicing choices during the 2012–2023 estimation period, both the industry intensities and bilateral import weights are fixed in 2010. Its interpretation nonetheless remains exposure-based: identification requires that the predetermined import composition not proxy for other country characteristics that differentially load on subsequent sanctions shocks. The resulting country index is standardized across countries. Panel c of Figure 2 reports its distribution across the same 88-country analytical sample. Both the underlying industry intensities and the bilateral industry weights are measured using the sectoral composition of countries' 2010 imports from China.

Panel d of Figure 2 compares the two 2010 country-level measures. Their Pearson correlation is -0.184, their Spearman rank correlation is -0.260, and the corresponding bivariate R^2 is 0.034 ($N = 88$). The limited raw association shows that the measures are not mechanical substitutes in the matched sample. The China import share measures the overall scale of imports from China, whereas the WCI summarizes the industry composition of those imports. This descriptive separation supports including the two measures as conceptually distinct exposures.

[10] These percentages use the analytical trade universe after the upstream product and industry crosswalks. Relative to raw 2010 imports from China, 98.77% of value reaches a non-missing Chinese WCI. The most of the additional loss occurs in the HS-to-industry mapping of bilateral trade data.

Additional Covariates. The empirical specifications progressively incorporate controls motivated by evidence on trade composition and geopolitics in currency invoicing (Boz et al., 2025), and on China-related trade and monetary institutions in international RMB use (Bahaj and Reis, 2026; Perez-Saiz and Zhang, 2023). I first add contemporaneous institutional indicators and then lagged measures of trade exposure, import composition, and political alignment.

The first group consists of three contemporaneous institutional indicators. The free trade agreement (FTA) indicator equals one from the entry-into-force year of a trade agreement with China, including a regional agreement. It captures policy-driven trade integration that may affect invoicing through trading relationships and transaction costs beyond the observed bilateral import share.

The swap-line indicator equals one when the country has an active People's Bank of China (PBoC) facility at year-end, including coverage through the shared PBoC–ECB arrangement for euro-area members. It captures access to an RMB liquidity backstop that can reduce funding risk and support RMB-denominated trade finance (Bahaj and Reis, 2026). This indicator measures facility availability, rather than actual drawings or the size of the announced ceiling.

The clearing-bank indicator equals one from the year in which an official RMB clearing bank is designated in the country. It captures local infrastructure that can lower the costs and operational frictions of clearing and settling cross-border RMB payments (Perez-Saiz and Zhang, 2023).[11]

The second group captures annual trade exposure, import composition, and geopolitical alignment. These three variables enter the regressions with a one-year lag. The time-varying China import share is the value of country i's imports from China divided by its total merchandise imports from all origins, expressed in percent. It accounts for changes in trade composition that may raise aggregate RMB invoicing even if currency choices within a given trading relationship are unchanged. This annual control is distinct from the fixed 2010 China import share used to construct sanctions exposure.

The commodity-import share is the sum of the World Bank's World Development Indicators import shares for fuels, ores and metals, and agricultural raw materials, each measured as a percentage of total merchandise imports from all origins. Following Boz et al. (2025), I use this share as a proxy for trade in relatively homogeneous goods. It accounts for import composition in markets where standardized

[11] I manually collect FTA entry-into-force dates from the China FTA Network of China's Ministry of Commerce, and swap-line agreements and clearing-bank designations from PBoC announcements and RMB internationalization reports. Renewals and changes in announced maximum amounts are recorded in the underlying swap-line data. The common PBoC–ECB ceiling is not allocated to individual euro-area members. Details of the three institutional indicators are reported in Table A.3.

pricing and vehicle-currency use tend to favor the dollar (Goldberg and Tille, 2008).

Using the time-varying ideal points of Bailey et al. (2017), I construct a political-alignment indicator equal to one when country i's ideal-point distance to China is smaller than its distance to the United States in year t, and zero otherwise. This control accounts for geopolitical preferences and relationships that may influence currency use alongside sanctions (Boz et al., 2025).

Together, these controls allow the sanctions-related invoicing response to be estimated conditional on observed institutional arrangements, trade composition, and political alignment.

Table A.4 reports summary statistics for the variables used in the fully controlled baseline and working-capital heterogeneity specifications over 2012–2023. Outcomes and time-varying regressors, including lagged variables, are summarized over the observations entering the regressions. This distinguishes variation across country–year observations, across the 88 sample countries, and across the 12 sample years. The EUR specification uses a smaller sample because euro-area countries are excluded.

3 US Sanctions and Spillover Effects

3.1 Exposure Design and Identification

This section examines whether increases in U.S. financial sanctions directed at China are followed by differential changes in the currencies used to invoice destination countries' imports. I combine the FX-neutral import invoicing shares and sanctions measures described in Section 2 with two country-level exposures fixed in 2010. The outcomes cover imports from all origins and are examined separately for the RMB, USD, EUR, and residual currency categories. I estimate dynamic panel regressions with country and year fixed effects and complement them with local projections (Jordà, 2005) to trace horizon-specific adjustment.

Let S_t^{CN} and $S_{i,t}^{\text{D}}$ denote the annual numbers of new U.S. SDN designations assigned to China and destination country i, respectively, including entities and individuals. Define the corresponding log-transformed sanctions measures as[12]

$$g_t = \ln(1 + S_t^{\text{CN}}) \qquad \text{and} \qquad d_{i,t} = \ln(1 + S_{i,t}^{\text{D}}).$$

[12]Using the inverse hyperbolic sine (IHS) transformation for both sanctions measures preserves the main findings: the RMB exposure coefficient remains positive and significant at the 1% level, and the USD coefficient remains negative and significant at the 5% level in all three specifications (Appendix Table B.4).

Because g_t is common to all destination countries, its level is absorbed by year fixed effects. I interact it with the predetermined exposure $E_{i,2010}$, the share of country i's total imports sourced from China in 2010, measured in percent:

$$\text{ChineseRisk}_{i,t} \equiv E_{i,2010} \times g_t. \tag{6}$$

In this interaction, g_t provides the time variation in sanctions intensity, and $E_{i,2010}$ provides the fixed cross-sectional exposure weight. This strategy mirrors macroeconomic approaches for identifying the heterogeneous transmission of aggregate shocks (Ottonello and Winberry, 2020; Crouzet and Mehrotra, 2020).

This interaction provides the identifying variation in the baseline exposure design. Its coefficient measures whether invoicing shares change more strongly in destinations with greater initial import exposure to China when China-directed sanctions intensity increases, conditional on the lagged outcome and included covariates. It does not identify the average global effect of g_t, which remains absorbed by year fixed effects. I also include $d_{i,t}$ separately to distinguish this exposure-related response from sanctions pressure directed at the destination itself.

A causal interpretation requires that, conditional on the fixed effects and included covariates, countries with different values of $E_{i,2010}$ would not otherwise have experienced systematically different changes in invoicing shares in the same years in which g_t changed. Fixing exposure in 2010 prevents subsequent invoicing choices or sanctions realizations from mechanically changing the exposure measure. The main threat is another aggregate event that coincides with China-directed sanctions and affects countries differentially according to their initial trade exposure to China. Year fixed effects absorb the component of such events that is common across countries, but not heterogeneous effects correlated with $E_{i,2010}$.

I next examine whether sanctions-related invoicing responses vary with the working-capital intensity of imports from China. Conditional on the initial China import share, $E_{i,2010}$, a higher $\text{WC}_{i,2010}$ indicates that the destination's 2010 import basket from China is concentrated in covered industries with greater net operating working-capital requirements relative to sales. If sanctions increase the risk that payments on dollar-invoiced exports are delayed or blocked, banks may restrict the working-capital credit available before production. This mechanism motivates testing whether RMB invoicing rises more, and USD invoicing falls more, among destinations whose initial Chinese import baskets have

greater working-capital requirements. The working-capital index combines industry measures derived from Chinese A-share listed firms' balance sheets and income statements with destination-specific import weights. It measures the operating funding needs embedded in import composition; these data do not establish whether the exporters serving each destination face binding credit constraints. A response pattern matching the proposed prediction would therefore be consistent with the financing channel, but would not by itself identify that channel. The weak raw correlation between $E_{i,2010}$ and $\text{WC}_{i,2010}$ in Figure 2 documents their limited overlap in the sample.

3.2 Panel Linear Regressions

I estimate dynamic panel regressions over 2012–2023. The sample begins after the expansion of cross-border RMB trade settlement and ends in 2023 because invoicing data for 2024 remain incomplete.[13] The regressions relate changes in import invoicing shares to China-directed sanctions pressure interacted with predetermined trade exposure, conditional on destination-directed sanctions, observed covariates, and country and year fixed effects.

Specification. Building upon the established empirical frameworks of Goldberg and Tille (2008) and Boz et al. (2022, 2025), I specify the following equation for each currency $c \in \{\text{RMB}, \text{USD}, \text{EUR}, \text{Residual}\}$:

$$\begin{aligned}\text{Invoicing}_{i,t}^{(c)} = \alpha_i^{(c)} + \beta_t^{(c)} + \xi^{(c)}\text{Invoicing}_{i,t-1}^{(c)} + \gamma^{(c)\prime} X_{i,t-1} + \psi^{(c)\prime} Z_{i,t} \\ + \delta^{(c)} d_{i,t} + \theta^{(c)}\text{ChineseRisk}_{i,t} + \epsilon_{i,t}^{(c)}. \end{aligned} \tag{7}$$

The dependent variable is the FX-neutral import invoicing share, measured in percentage points. To account for the pronounced inertia characteristic of currency invoicing, Equation 7 includes the lagged dependent variable ($\text{Invoicing}_{i,t-1}^{(c)}$). $d_{i,t} = \ln(1 + S_{i,t}^{\text{D}})$ controls for SDN designations directed at the destination country capturing idiosyncratic sanctions risk. As defined in Equation 6, $\text{ChineseRisk}_{i,t} = E_{i,2010} g_t$ combines the common China-directed sanctions measure with country i's predetermined import exposure to China. The vector $X_{i,t-1}$ contains the lagged import share from China, the lagged

[13] In July 2009, China initiated a pilot program for cross-border RMB trade settlement, initially restricted to Shanghai and four cities in Guangdong province. It was not until the end of 2011 that the international use of the RMB was authorized nationwide. The extension of the coverage of cross-border RMB trade settlement to all provinces and municipalities was officially implemented via the "Notice on Expansion of the List of Regions for Cross-border RMB Settlement," jointly issued by the PBoC and the other five regulatory authorities in August 2011.

homogeneous-goods import share, and lagged political alignment with China. The time-varying China import share accounts for changes in bilateral trade intensity that may accompany changes in RMB use; it is distinct from the fixed 2010 exposure entering ChineseRisk. The vector $Z_{i,t}$ contains contemporaneous indicators for an FTA with China, access to an RMB swap line, and an official RMB clearing bank. The coefficients therefore describe responses conditional on these observed trade and institutional characteristics.[14] Finally, α_i and β_t absorb country and time fixed effects, respectively. Standard errors are clustered by country.

The coefficient $\theta^{(c)}$ measures the exposure gradient in the response to the intensity of U.S. sanctions against Chinese entities. For a change Δg and baseline exposure E, the contemporaneous exposure-related component is $\theta^{(c)} E \Delta g$.[15] Both invoicing shares and E are expressed in percentage points, so a country importing 10 percent of its goods from China has $E = 10$.

Baseline Results. Table 1 reports three specifications for each currency. Columns (1), (4), (7), and (10) include the lagged dependent variable, both sanctions measures, and country and year fixed effects. The second column within each currency group adds the institutional indicators, and the third adds the lagged trade and political-alignment controls. Euro-area countries are excluded from the EUR regressions. The fully controlled specifications contain 831 country-year observations from 88 countries for RMB, USD, and Residual, and 640 observations from 69 countries for EUR.

The reported estimates show a positive exposure-related response of RMB invoicing and a negative response of USD invoicing. Across the three specifications, the ChineseRisk coefficient ranges from 0.014 to 0.017 for RMB and from -0.062 to -0.053 for USD. The RMB coefficients are statistically significant at the 1 percent level, and the USD coefficients at the 5 percent level. The RMB and USD patterns persist after the observed policy and trade controls are included. The opposite signs for RMB and USD are consistent with a reallocation of invoicing away from the dollar and toward the RMB. The estimates for EUR and Residual are imprecise and do not establish a systematic response.

These findings relate to evidence that financial and geopolitical frictions shape currency choice. Using detailed Turkish export data, Corsetti et al. (2025) document a shift from dollar to lira invoicing

[14] As a robustness check against potential simultaneity bias, lagged policy variables ($Z_{i,t-1}$) was considered. The results show that the signs and significance levels of the main coefficients are fully robust to this alternative specification.

[15] The specific choice of the baseline year for exposure does not materially alter the baseline results, as the import share from China exhibits high persistence over time(**??**).

Table 1: US Financial Sanctions on Chinese Entities and Global Invoicing Currencies

	RMB			USD			EUR			Residual		
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
ChineseRisk	0.017***	0.016***	0.014***	-0.053**	-0.059**	-0.062**	0.011	0.014	0.018	0.050	0.051	0.048
	(0.005)	(0.006)	(0.005)	(0.024)	(0.026)	(0.027)	(0.025)	(0.023)	(0.022)	(0.034)	(0.037)	(0.038)
Idio. Sanc.	-0.026	-0.023	-0.024	-0.187	-0.180	-0.152	0.288	0.291	0.270	0.181	0.182	0.183
	(0.041)	(0.041)	(0.043)	(0.258)	(0.255)	(0.264)	(0.199)	(0.205)	(0.213)	(0.266)	(0.270)	(0.278)
Lag. Dep. Var.	0.459***	0.454***	0.438***	0.509***	0.510***	0.514***	0.579***	0.573***	0.565***	0.497***	0.496***	0.501***
	(0.140)	(0.135)	(0.121)	(0.061)	(0.062)	(0.064)	(0.078)	(0.079)	(0.076)	(0.098)	(0.101)	(0.105)
LR Effects	0.031***	0.030***	0.025***	-0.108**	-0.121**	-0.127***	0.027	0.033	0.041	0.100	0.100	0.095
	(0.010)	(0.010)	(0.008)	(0.045)	(0.047)	(0.049)	(0.061)	(0.055)	(0.053)	(0.062)	(0.068)	(0.070)
Contemp. Ctrl.		✓	✓		✓	✓		✓	✓		✓	✓
Lagged Ctrl.			✓			✓			✓			✓
Year & Cty FE	✓	✓	✓	✓	✓	✓	✓	✓	✓	✓	✓	✓
Obs	849	849	831	849	849	831	658	658	640	849	849	831
Countries	91	91	88	91	91	88	72	72	69	91	91	88
Within R^2	0.175	0.177	0.194	0.278	0.279	0.281	0.327	0.328	0.328	0.251	0.252	0.252

Notes: This table presents the baseline dynamic panel fixed-effects estimation results. The dependent variables are the FX-neutralized import invoicing shares of the respective currencies (RMB, USD, EUR, and Residual). **ChineseRisk** is $E_{i,2010}g_t$, where $E_{i,2010}$ is the country's 2010 import share from China, measured in percent, and g_t is the natural logarithm of one plus annual China-directed U.S. SDN designations, including entities and individuals. **Idio. Sanc.** is $d_{i,t}$, the natural logarithm of one plus annual U.S. SDN designations directed at the importing country. **LR Effects** reports the long-run multiplier of the sanctions shock, computed via the Delta method based on the lagged dependent variable (**Lag. Dep. Var.**). Contemporaneous controls encompass indicators for bilateral FTAs, currency swap lines, and official RMB clearing banks. Lagged controls include the import share from China, the import share of homogeneous goods, and a political alignment index. All specifications include country and year fixed effects. Standard errors, clustered at the country level, are reported in parentheses. * p<0.10, ** p<0.05, *** p<0.01.

and greater reliance on cash-in-advance payments in trade with Russia following the 2022 invasion of Ukraine and Western sanctions. In a different policy setting, Bahaj and Reis (2026) document increased RMB use following the introduction of PBoC swap lines, while Perez-Saiz and Zhang (2023) find that swap lines and offshore clearing banks facilitate RMB use in cross-border payments with China. My analysis complements these studies by examining how annual import invoicing shares respond to China-directed sanctions pressure across countries with different predetermined import exposures to China. The positive RMB and negative USD coefficients persist after conditioning on the observed institutional indicators.

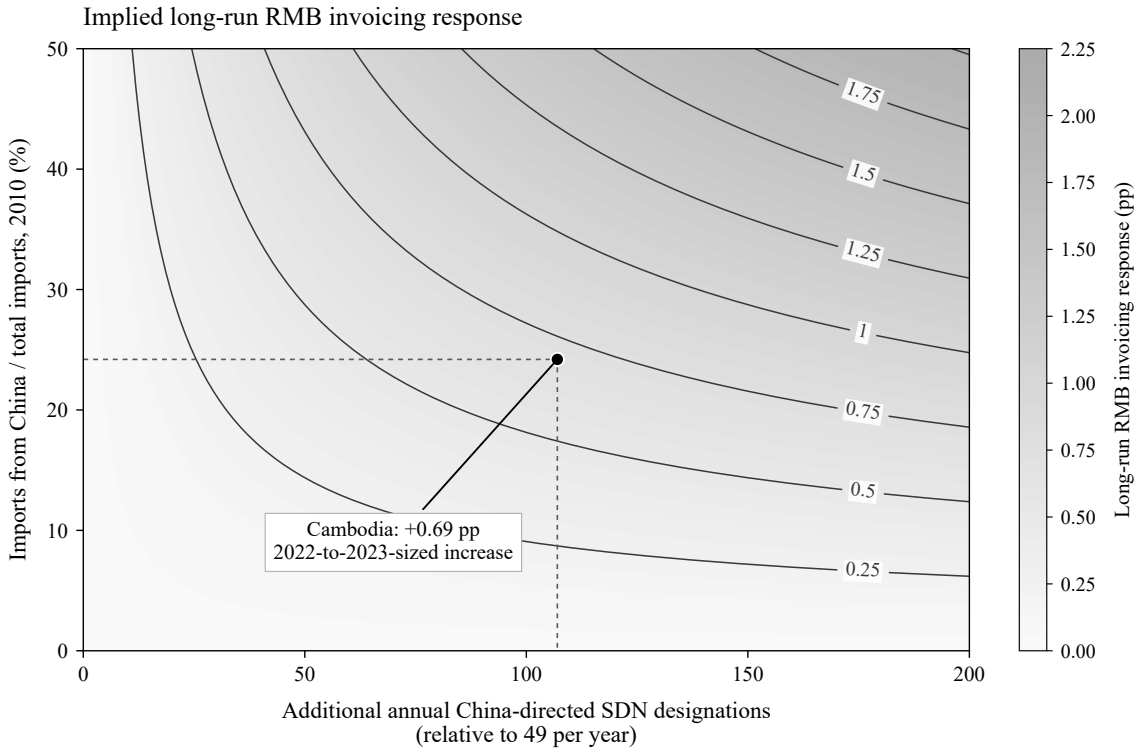
The lagged dependent variable also links the contemporaneous estimates to the adjustment implied by persistence in invoicing shares. Holding other covariates fixed, the long-run exposure-related response to a sustained increase Δg converges to

$$\mathrm{LR}^{(c)}(E,\Delta g)=\frac{\theta^{(c)}}{1-\xi^{(c)}}E\Delta g, \qquad |\xi^{(c)}|<1. \tag{8}$$

The LR Effects row in Table 1 reports the unscaled long-run multiplier $\theta^{(c)}/(1-\xi^{(c)})$. For a one-period

increase in sanctions intensity, the associated response instead decays according to the estimated autoregressive coefficient; the same multiplier summarizes the sum of those responses over subsequent periods.

Figure 4: Model-Implied Long-Run RMB Invoicing Response to a Sustained Increase in U.S. Sanctions



Notes: This figure uses the estimated dynamic panel coefficients from Column (3) of Table 1. The vertical axis measures the share of a country's imports originating in China in 2010, expressed in percent. The horizontal axis measures additional annual China-directed SDN designations, relative to the 2022 total of 49. For an increase of k designations, the simulation compares a reference path with 49 new designations each year against a path with $49+k$ new designations each year, starting from the same initial conditions and holding other factors constant. Shading and contours report the resulting long-run difference in the RMB invoicing share, in percentage points: $\left\{\hat{\theta}E/(1-\hat{\xi})\right\}[\ln(50+k)-\ln(50)]$ where E is baseline import exposure, $\hat{\theta}$ is the ChineseRisk coefficient, and $\hat{\xi}$ is the lagged dependent-variable coefficient. They illustrate the conditional exposure-related component of the estimated relationship.

To retain the normalization used in the reported results, consider a benchmark one-standard-deviation change $\Delta g = 1.36$ and a country with $E = 10$. The parsimonious estimates imply contemporaneous changes of approximately 0.23 percentage points for RMB and -0.72 percentage points for USD. Under a sustained change of the same size, the fully controlled estimates imply long-run changes of approximately 0.34 and -1.73 percentage points, respectively. These are conditional implications of the

estimated dynamic relationship.[16]

Figure 4 illustrates how the estimated long-run RMB relationship varies with baseline China exposure and the assumed change in sanctions intensity, based on the estimates in Column (3) of Table 1. The larger predicted response at greater exposure follows directly from the interaction specification. The figure is therefore an illustration of the estimated relationship rather than independent evidence of a nonlinear response or a forecast for individual countries.

Cambodia illustrates the economic magnitude of the estimates. China accounted for approximately 24.2% of Cambodia's imports in 2010, the fixed exposure used in the regressions. Consider two hypothetical scenarios with the same initial conditions. In the reference scenario, annual China-directed SDN designations remain at 49. In the higher-sanctions scenario, they rise to 156 and remain at that level each year. These counts correspond to the observed annual totals in 2022 and 2023, respectively. Holding other factors constant, the fully controlled specification implies an exposure-related difference in Cambodia's RMB invoicing share of approximately 0.39 percentage points between the two scenarios in the first year. Because invoicing shares are persistent, this difference gradually widens to approximately 0.69 percentage points if the higher annual designation rate continues. The 0.69-percentage-point figure is the total long-run difference. For perspective, the long-run difference is approximately 37% of Cambodia's observed FX-neutral RMB invoicing share of about 1.9% in 2023.

Heterogeneity by Working-Capital Intensity. I further examine whether the sanctions-related invoicing response varies with the working-capital intensity of imports from China. The proposed financing mechanism suggests that larger operating-liquidity needs can strengthen exporters' incentives to reconsider dollar invoicing when U.S. financial sanctions increase the risk of disruptions to dollar-linked credit and payments.

For the specification reported in Table 2, let $D_i^{\text{WC}} = \mathbb{I}\{\text{WC}_{i,2010} \leq Q_{0.10}\}$ indicate membership in

[16] Given the relatively short time dimension of the panel, the standard two-way fixed effects (TWFE) estimator may underestimate the autoregressive coefficient $\xi^{(c)}$ due to Nickell (1981) bias. To address this, Table B.1 reports estimates using the bias-corrected least squares dummy variable (LSDVC) estimator developed by Kiviet (1995). The LSDVC specification yields comparable long-run effects. While while the USD coefficient remains negative but is less precisely estimated, the positive RMB coefficient remains statistically significant

Table 2: U.S. Financial Sanctions and Global Invoicing Currencies: Working-Capital Intensity

	RMB	USD	EUR	Residual
	(1)	(2)	(3)	(4)
ChineseRisk	0.014***	-0.062**	0.018	0.047
	(0.005)	(0.027)	(0.022)	(0.038)
ChineseRiskD_i^{WC}	-0.014*	0.326***	-0.219***	-0.047
	(0.008)	(0.091)	(0.077)	(0.103)
$g_t D_i^{\text{WC}}$	0.143**	-2.066**	1.727**	-0.060
	(0.065)	(0.879)	(0.825)	(0.985)
Lag. Dep. Var.	0.438***	0.515***	0.564***	0.500***
	(0.121)	(0.065)	(0.078)	(0.105)
Contemp. Ctrl.	✓	✓	✓	✓
Lagged Ctrl.	✓	✓	✓	✓
Year & Cty FE	✓	✓	✓	✓
Observations	831	831	640	831
Countries	88	88	69	88
Within R^2	0.194	0.282	0.328	0.252

Notes: The table presents the effects of working-capital intensity Let E_i denote the fixed 2010 China import share, measured in percent, and let $g_t = \ln(1 + \text{SDN}_{\text{CN},t})$, where the count includes annual China-directed additions. $\text{ChineseRisk}_{i,t} = E_i g_t$. D_i^{WC} identifies the bottom decile of the original WCI reference distribution. Policy controls are the China FTA, swap-line, and RMB clearing-bank indicators. Trade controls are lagged China import and homogeneous-goods shares; political alignment is also lagged. All columns additionally control for destination-directed sanctions and the lagged dependent variable, and include country and year fixed effects. The EUR sample excludes euro-area countries under the original sample rule. Country-clustered standard errors appear in parentheses. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

the bottom decile of the original working-capital index distribution. I estimate

$$\begin{aligned}\text{Invoicing}_{i,t}^{(c)} = {} & \alpha_i^{(c)} + \beta_t^{(c)} + \xi^{(c)}\text{Invoicing}_{i,t-1}^{(c)} + \gamma^{(c)\prime} X_{i,t-1} + \psi^{(c)\prime} Z_{i,t} \\ & + \delta^{(c)} d_{i,t} + \theta^{(c)}\text{ChineseRisk}_{i,t} + \kappa^{(c)}\left(\text{ChineseRisk}_{i,t} D_i^{\text{WC}}\right) \\ & + \lambda^{(c)}\left(g_t D_i^{\text{WC}}\right) + \epsilon_{i,t}^{(c)}. \end{aligned} \tag{9}$$

All specifications retain the baseline controls, country and year fixed effects, and country-clustered standard errors. Country fixed effects absorb the time-invariant terms D_i^{WC} and $E_{i,2010} D_i^{\text{WC}}$, while year fixed effects absorb the common sanctions index g_t.

The additional interaction $g_t D_i^{\text{WC}}$ allows the group difference in sanctions responses to have an intercept as well as an exposure-dependent component. Accordingly, $\kappa^{(c)}$ measures the difference in

exposure gradients. For a one-unit increase in g_t, the conditional short-run response difference between the low-WCI group and the remaining countries at the same baseline exposure E is $\Delta R^{(c)}(E) = \lambda^{(c)} + \kappa^{(c)} E$. A negative value of $\Delta R^{(c)}(E)$ indicates a smaller response in the low-WCI group at that exposure.

In Table 2, the RMB interaction coefficient is -0.014 indicating a flatter exposure gradient in the low-WCI group. This does not imply a uniformly smaller RMB response, because the estimated $\lambda^{(\text{RMB})}$ is positive. At higher exposure values, the fitted contrast $\widehat{\Delta R}^{(\text{RMB})}(E)$ becomes negative, a direction consistent with the proposed financing-channel interpretation.[17] The USD and EUR interaction coefficients are positive and negative, respectively. These estimates suggest that the relationship between sanctions pressure and invoicing varies with the financing requirements of the goods imported from China. Their interpretation is limited by the country-level nature of the data and the small low-WCI estimation group.

Appendix C provides complementary evidence on timing through year-specific exposure profiles and local projections. The projections relate sanctions pressure to changes in invoicing shares from $t-1$ to $t+h$ for $h = 0, \ldots, 3$, including a comparison across WCI groups. Longer horizons, however, require follow-up outcomes that are unavailable for the most recent sanctions observations. At $h = 3$, the 2022 and 2023 reference years are excluded because the corresponding 2025 and 2026 outcomes are not observed. The unbalanced panel also produces substantial changes in country and reference-year coverage across horizons and estimation windows. Differences across projections therefore reflect changes in sample composition as well as possible dynamic adjustment. These exercises complement the panel evidence but do not establish the full adjustment path following the latest sanctions escalation.

3.3 Robustness: The Russia Effect and Placebo Tests

In this subsection, I assess the sensitivity of the panel estimates to three changes in sample or specification: excluding Russia, replacing baseline China exposure with exposure to Germany, and replacing current China-directed sanctions intensity with its two-year lead. These exercises examine whether the reported invoicing patterns depend on Russia's observations and whether similar relationships appear

[17] At China import exposures of 20% and 30%, the estimated low-WCI-minus-remaining-group contrasts are -0.135 and -0.273 percentage points per unit increase in g_t, respectively.

under alternative exposure and timing measures.

Excluding Russia. A particular concern is that Russia's own adjustment to Western financial sanctions may disproportionately influence the estimated China-exposure relationship. Existing research documents a limited and uneven international role for the RMB. Its use in cross-border payments is associated with trade and institutional ties to China (Perez-Saiz and Zhang, 2023), while evidence on trade invoicing suggests a limited role as a vehicle currency in transactions between third countries (Chowdhry, 2024; Boz et al., 2025). Against this background, longstanding sanctions against Russia could have encouraged greater RMB use in its trade with China. If this adjustment covaries with China-directed sanctions intensity, it could influence the estimated coefficient on $E_{i,2010}g_t$. The baseline specification already adjusts for sanctions directed at the importing country through $d_{i,t}$, with a common coefficient across countries. This adjustment and the other included controls may nevertheless leave Russia-specific responses incompletely captured, particularly if their magnitude or timing differs from the relationship imposed by the model.

I therefore re-estimate the fully controlled baseline specification across three samples: the original sample, a sample excluding Russia, and a sample jointly excluding Russia and four selected trading partners with important trade links to Russia, including Brazil, India, Turkey, and Kazakhstan, to assess whether the positive RMB and negative USD relationships depend on its observations.[18] In Table 3, Columns (1)–(3) report the RMB estimates, and Columns (4)–(6) report the USD estimates. All columns retain the same controls and country and year fixed effects. Russia contributes invoicing observations only for 2012–2021, so its own invoicing changes following the 2022 invasion of Ukraine do not enter these regressions directly.

Excluding Russia alone leaves the RMB and USD coefficients nearly unchanged, at approximately 0.014 and -0.062, respectively. Under the broader exclusion, the RMB coefficient remains similar at approximately 0.015, while the USD coefficient remains negative but declines in absolute magnitude to approximately -0.052. Thus, the positive RMB and negative USD relationships persist across these samples, with some attenuation of the USD relationship under the broader exclusion. Table B.2 reports the corresponding estimates for all four currency categories using the same three sample definitions. Table B.3 reports the corresponding Russia-only exclusion for the working-capital heterogeneity spec-

[18] Iran and Syria are also absent from the baseline estimation sample.

ification in Equation 9. The main coefficient patterns and the estimated group contrasts at a common baseline China import share of 10 percent remain broadly similar.

Table 3: Baseline Estimates under Alternative Country Exclusions

	RMB			USD		
	Full	Excl. Russia	Excl. five	Full	Excl. Russia	Excl. five
	(1)	(2)	(3)	(4)	(5)	(6)
$E_i g_t$	0.014***	0.014***	0.015***	-0.062**	-0.062**	-0.052*
	(0.005)	(0.005)	(0.006)	(0.027)	(0.027)	(0.026)
Destination sanctions	-0.024	-0.025	-0.026	-0.152	-0.112	-0.261
	(0.043)	(0.045)	(0.048)	(0.264)	(0.276)	(0.282)
Lagged dependent variable	0.438***	0.433***	0.428***	0.514***	0.515***	0.513***
	(0.121)	(0.120)	(0.119)	(0.064)	(0.065)	(0.070)
p: $\theta = 0$	0.0090	0.0097	0.0094	0.0256	0.0245	0.0509
Observations	831	821	784	831	821	784
Countries	88	87	83	88	87	83
Within R^2	0.194	0.188	0.189	0.281	0.281	0.279

Notes: This table compares the fully controlled baseline specification in Equation 7 using the original sample, excluding Russia only, and jointly excluding Russia, Brazil, India, Turkey, and Kazakhstan ("Excl. five"). The estimation period is 2012–2023; Russian observations cover 2012–2021. The dependent variables are FX-neutral RMB and USD import invoicing shares, in percentage points. E_i is the fixed 2010 China import share, in percent. g_t and the destination sanctions measure are the natural logarithms of one plus annual U.S. SDN designations directed at China and the importing country, respectively, including entities and individuals. All columns include country and year fixed effects, the lagged dependent variable, destination-directed sanctions, contemporaneous FTA, RMB swap-line and RMB clearing-bank indicators, and lagged China import share, homogeneous-goods import share, and political alignment. Standard errors in parentheses are clustered by country. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

Exposure placebo. I next examine whether the baseline pattern also appears when China-directed sanctions intensity is interacted with exposure to a different major supplier. Let E_i^{DEU} denote the share of country i's imports originating in Germany in 2010, expressed in percent. I replace $E_i g_t$ in Equation 7 with

$$P_{it}^{\text{DEU}} = E_i^{\text{DEU}} g_t,$$

retaining the lagged dependent variable, destination-directed sanctions, baseline covariates, and country and year fixed effects. Germany is excluded from this specification.

Columns (3) and (4) of Table 4 report coefficients of approximately -0.002 for RMB and 0.035 for USD, with country-clustered standard errors of 0.002 and 0.022, respectively. Neither coefficient is statistically distinguishable from zero at conventional levels. The Germany-exposure specification therefore does not reproduce the baseline combination of a positive RMB gradient and a negative USD gradient. This finding is consistent with the relevance of China-specific trade exposure, although Ger-

many exposure need not capture all alternative sources of differential invoicing trends.

Timing placebo. Finally, I replace the contemporaneous interaction $E_i g_t$ with

$$P_{it}^{\text{lead}} = E_i g_{t+2}.$$

The outcome remains the invoicing share in year t, and the other controls and fixed effects are unchanged. This specification examines whether sanctions intensity dated two years ahead is associated with current invoicing shares through baseline China exposure.

Table 4: US Financial Sanctions on Chinese Entities and Global Invoicing Currencies (Placebo Tests)

	Baseline		**Exposure Placebo**		**Temporal Placebo**	
	RMB	**USD**	**RMB**	**USD**	**RMB**	**USD**
	(1)	(2)	(3)	(4)	(5)	(6)
ChineseRisk	0.014^{***}	-0.062^{**}	-0.002	0.035	-0.004	-0.016
	(0.005)	(0.027)	(0.002)	(0.022)	(0.004)	(0.033)
Idio. Sanc.	-0.024	-0.152	-0.027	-0.182	-0.032^{*}	-0.113
	(0.043)	(0.264)	(0.046)	(0.271)	(0.017)	(0.284)
Lag. Dep. Var.	0.438^{***}	0.514^{***}	0.471^{***}	0.511^{***}	0.980^{***}	0.563^{***}
	(0.121)	(0.064)	(0.130)	(0.063)	(0.307)	(0.057)
LR Effects	0.025^{***}	-0.127^{***}	-0.004	0.072	-0.179	-0.036
	(0.008)	(0.049)	(0.004)	(0.050)	(2.917)	(0.073)
Contemp. Ctrl.	✓	✓	✓	✓	✓	✓
Lagged Ctrl.	✓	✓	✓	✓	✓	✓
Year & Cty FE	✓	✓	✓	✓	✓	✓
Obs	831	831	819	819	751	751
Countries	88	88	87	87	88	88
Within R^2	0.194	0.281	0.169	0.277	0.493	0.316

Notes: This table presents placebo tests to validate the identification strategy of the baseline dynamic panel fixed-effects regressions. The dependent variables are the import invoicing shares of the RMB and USD. Columns (1) and (2) replicate the baseline results for comparison. Columns (3) and (4) report the **Exposure Placebo**, which tests for spurious network trends by replacing the destination country's 2010 baseline import share from China with the baseline import share from Germany ($\overline{\text{Exposure}}_{i,2010}^{\text{DEU}}$). Columns (5) and (6) report the **Temporal Placebo**, which tests for pre-existing trends and anticipation effects by substituting the contemporaneous sanctions shock with a 2-year leading value of the global U.S. sanctions against Chinese entities ($\text{sanction}_{cn,t+2}^{\log}$). **Idio. Sanc.** denotes the log-transformed idiosyncratic U.S. sanctions directly targeting the destination country. **LR Effects** reports the long-run multiplier of the sanctions shock, computed via the delta method based on the lagged dependent variable (**Lag. Dep. Var.**). Contemporaneous controls encompass indicators for bilateral FTAs, currency swap lines, and official RMB clearing banks. Lagged controls include the import share from China, the import share of homogeneous goods, and a political alignment index. All specifications include country and year fixed effects. Standard errors, clustered at the country level, are reported in parentheses. * p<0.10, ** p<0.05, *** p<0.01.

Columns (5) and (6) of Table 4 report coefficients of approximately -0.004 for RMB and -0.016 for USD, with standard errors of 0.004 and 0.033, respectively. These estimates do not provide statistically significant evidence of the lead-exposure relationship in this specification. The placebo replaces the

current interaction with its lead; it does not estimate their separate contributions in a joint specification.

The placebo comparisons also involve changes in sample coverage. The exposure and timing specifications use 819 and 751 observations, respectively, compared with 831 in the baseline columns. Differences across specifications therefore reflect both the alternative regressor and the available observations. These checks provide evidence on specific sources of sensitivity.

The next section develops a model of a Chinese exporter to examine a financing mechanism consistent with the empirical patterns. In the model, the invoicing currency determines exposure to sanctions-sensitive disruptions in the usable amount of working-capital credit. The firm weighs the dollar's normal-time financing and hedging benefits against the risk of impaired dollar-linked credit access. Working-capital requirements increase the cost of such disruptions, while dollar-priced intermediate inputs create an incentive to retain dollar revenues as an operational hedge.

4 Theoretical Model for Currency Choice

This section develops a three-period partial-equilibrium model of invoicing-currency choice with working-capital financing. A Chinese exporter jointly chooses its invoicing currency and a preset price to maximize expected profits. Its choice between RMB and USD invoicing balances the normal-time benefits of dollar invoicing against greater exposure to sanctions-related disruptions in credit access. This trade-off provides an interpretation of the positive China-exposure gradient for RMB invoicing, the negative gradient for USD invoicing, and the differences across working-capital groups documented in Section 3. The model derives a currency-switching threshold and specifies the additional conditions needed to connect firm choices to the country-level estimates.

The contribution is to make the reliability of credit access depend on the invoicing currency choice in an otherwise familiar preset-price problem. Existing work establishes the role of hedging and input-currency matching in currency choice (Engel, 2006; Goldberg and Tille, 2008; Chung, 2016) and the complementarity between trade finance and invoicing (Gopinath and Stein, 2021; Bahaj and Reis, 2026). The central mechanism here concerns a reduction in usable credit capacity in a sanctions-stress state. Its cost depends on how much production must be financed before revenues arrive and on the alternatives available when usable credit line is impaired. This also distinguishes the proposed mech-

anism from the payment-circuit and reputational frictions studied by Corsetti et al. (2025). The model takes sanctions risk and financial infrastructure as given.

4.1 Environment, Timing, and Currency Choice

A risk-neutral Chinese exporter j in sector s serves a single destination i. It chooses producer-currency pricing (PCP), denoted $c = e$ and interpreted as RMB invoicing, or dominant-currency pricing (DCP), denoted $c = v$ and interpreted as USD invoicing. The destination's own currency need not be either of these. The binary choice isolates the RMB–USD margin. Throughout the firm-level analysis, destination i is fixed. I retain i on destination demand and its shifters, and suppress it on firm-specific prices, profits, values, and switching thresholds.

At $t = 0$, the firm observes the sanctions-risk environment τ, takes its predetermined credit limit $\bar{\mathcal{B}}_j$ as given, and chooses an invoicing currency and a price fixed in that currency. I treat this limit as a firm-level benchmark reflecting established financing capacity, including historical cash flows, not a transaction-by-transaction receivable collateral. At $t = 1$, the macroeconomic state $\mathbf{X}$, including exchange rates, borrowing rates, and destination demand shifters, is realized. Given the preset price, this state determines the quantity demanded and the production and financing conditions facing the firm. The credit state $\xi \in \{g, b\}$ is also realized, with g and b denoting the good and bad states, respectively. Given the chosen invoicing currency c, usable credit is $\mathcal{A}^c_{j,\xi} = \Theta^c_\xi \bar{\mathcal{B}}_j$, where Θ^c_ξ is the credit-availability factor associated with currency c in state ξ. Having observed both macroeconomics state and credit state, the firm chooses its input mix and financing to minimize the cost of fulfilling contractual demand, and then produces. At $t = 2$, it collects export revenue, repays borrowing, and pays wages. Exchange rates are uncertain at $t = 0$ but do not change further after $t = 1$.

Although the credit state is realized at $t = 1$, the firm assesses its likelihood when choosing its invoicing currency and preset price at $t = 0$. The observed sanctions-risk environment τ summarizes public information available at that date about the risk of a subsequent disruption to dollar-related credit and payment services. I represent the relationship between this information and the future credit state by

$$\Pr(\xi = b \mid \tau) = \pi(\tau), \qquad \pi'(\tau) > 0. \tag{10}$$

In the benchmark, firms know the mapping $\pi(\cdot)$ and use $\pi(\tau)$ as the conditional probability of the

bad state in their expected-profit calculations. A higher τ therefore makes a future credit disruption more likely, while the realized state remains unknown when the contract is priced. The firm chooses its invoicing currency and price in anticipation of this possibility; after the state is realized, it adjusts production and financing to the credit actually available.

For the sanctions comparative static, I let τ change this probability while holding fixed the joint distribution of $\mathbf{X}$, each firm's reference credit limit $\bar{\mathcal{B}}_j$, and the state-specific credit-availability factors Θ_ξ^c. I also assume that ξ is independent of $\mathbf{X}$ conditional on τ. These restrictions isolate a change in the likelihood of impaired credit access: neither the macroeconomic distribution nor the amount of usable credit within a given state changes with τ.

I assume that export contracts are settled in the currency in which they are invoiced. The exporter commits to the associated payment arrangements and lending relationships before the credit state is observed. I treat these arrangements as fixed over the contract. Expected disruptions to settlement in the chosen currency can therefore affect the credit banks make available.

Disclosed lending agreements illustrate the contingent nature of credit access. Article 10 of the agreement between Multi-Fineline Electronix (Suzhou No. 2) and Shanghai Pudong Development Bank permits the lender to adjust the credit limit and its duration, or terminate credit provision, under circumstances including:

> 2. "There happened or has potential critical financial risks at Party A' s region;"
>
> 4. "Party A is or will be under significant business difficulties or risks;"
>
> *Source:* Comprehensive Credit Line Agreement (2006), Article 10(2) and (4). Party A is the borrower.

These provisions encompass anticipated difficulties, allowing credit access to be adjusted before an actual repayment failure. The agreement between Shenzhen Highpower Technology and Bank of China also specifies triggering circumstances and the credit restrictions available to the lender.[19]

> 5. "Party [A] is or will be under significant business difficulties or risks [⋯]"
>
> (2) "Reduce, temporarily pause or permanently terminate Party A's Credit limit in part or in all"

[19] The disclosed English version of Clause 10(5) refers to Party B, the lender. I interpret this as an apparent party-designation error: Clause 7 uses the same borrower-distress wording with Party A, as does Article 10(4) of the Shanghai Pudong Development Bank agreement quoted above. I therefore replace B with [A] in the quotation. This editorial correction is based on contextual comparison.

Source: Comprehensive Credit Line Contract (2018), Clause 10, item (5) and remedy (2). Party A is the borrower.

These contractual rights motivate the distinction between the firm's reference credit limit and the amount it can actually draw. Under the assumed link between invoicing currency and payment infrastructure, sanctions-related disruptions to dollar payment processing may weaken lenders' confidence in the timely arrival of the firm's operating cash inflows and its ability to service debt. Lenders may consequently restrict drawings under the existing facility before production. I represent this possible response by $\Theta_b^v < \Theta_g^v$. For the benchmark model, I assume that the payment infrastructure supporting RMB invoicing remains operational under this particular disruption and that its associated credit availability is unchanged, $\Theta_b^e = \Theta_g^e$.[20]

An illustrative exporter. Consider a hypothetical Chinese manufacturer exporting electric vehicles to Korea. Its predetermined credit limit is RMB 10 million, reflecting its established financing capacity. At $t = 0$, the firm observes τ and chooses whether to invoice in USD or RMB, together with a price fixed in that currency. For illustration, usable credit under USD invoicing is RMB 10 million in the good state and RMB 5 million in the bad state; under RMB invoicing, it is RMB 8 million in either state.

At $t = 1$, exchange rates, borrowing rates, demand conditions, and the credit state are realized. For a given macroeconomic state and the preset USD price, suppose the cost-minimizing input plan at borrowing rates requires RMB 8 million of external funding. In the good state, the firm can draw the full amount under its existing facility. In the bad state, usable credit covers only RMB 5 million, leaving a funding shortfall of RMB 3 million if the input plan is unchanged. The firm can obtain emergency finance at an additional cost and adjust its input mix to reduce expenditure requiring advance payment. Its actual emergency borrowing may therefore be smaller than RMB 3 million while contractual demand is still fulfilled.

At $t = 2$, the firm receives export revenue, repays borrowing, and pays wages. Anticipating these outcomes, it chooses the currency and preset price that maximize expected profits at $t = 0$. A higher τ increases the probability of restricted USD credit access and hence the expected cost associated with that exposure.

[20] Discussing RMB settlement, Adam Wolfe, quoted in the *Financial Times* on October 23, 2025, emphasized China's ability to continue trading: "no matter what happens, it can still trade." The remark illustrates the rationale for maintaining alternative payment channels.

4.2 Demand and Normal Financing Costs

Let $\mathcal{E}_{ed}$ and $\mathcal{E}_{ev}$ denote units of RMB per unit of the destination currency and per USD, respectively. An increase in either rate therefore indicates a depreciation of the RMB against the corresponding currency. These exchange rates are components of macroeconomic state $\mathbf{X}$ realized at $t = 1$. For the two invoicing choices $c \in \{e, v\}$, where e denotes RMB and v denotes USD, define $a_e = 1$ and $a_v = \mathcal{E}_{ev}$, so that a_c converts one unit of the invoicing currency into RMB. A preset price P^c denominated in c therefore yields RMB unit revenue $p^c = a_c P^c$ and a destination-currency price $a_c P^c / \mathcal{E}_{ed}$. For a given preset price, the invoicing currency determines how exchange-rate movements translate into destination-currency prices and demand.

The firm sells a differentiated variety to destination i and faces constant-elasticity demand. Its price is fixed in the chosen invoicing currency c before the macroeconomic state is realized. Conditional on $c \in \{e, v\}$, the preset price, and $\mathbf{X}$, demand is

$$Y_i^c(P^c, \mathbf{X}) = \left(\frac{a_c P^c}{\mathcal{E}_{ed} Q_i} \right)^{-\epsilon}, \qquad \epsilon > 1, \tag{11}$$

where ϵ is the demand elasticity and $Q_i > 0$ is a destination-demand shifter, taken as given by the firm. For a given preset price and macroeconomic state, the firm fulfills the same demand in either credit state. Sanctions risk can nevertheless affect quantities through the price chosen ex ante.

I now formalize the firm' s production technology and cost structure. Firm j in sector s operates with a constant-returns-to-scale Cobb–Douglas technology. Combining an intermediate input bundle X_j and labor L_j, output is given by

$$Y_j = X_j^{\alpha} L_j^{1-\alpha}, \qquad 0 < \alpha < 1. \tag{12}$$

where α is the intermediate-input share of total production cost.

The exporters supplying intermediate inputs to firm j choose the prices and invoicing currencies of those inputs. Firm j takes these upstream pricing decisions as given. Its exposure to dollar-priced inputs is summarized by ω_v, which is predetermined given its sourcing structure. The intermediate bundle aggregates RMB- and USD-priced inputs with fixed Cobb–Douglas weights $1 - \omega_v$ and ω_v. The

input bundle's pre-financing price in RMB is

$$\bar{P}_X^e(\mathbf{X}) = (P_x^e)^{1-\omega_v}(\mathcal{E}_{ev}P_x^v)^{\omega_v}, \qquad 0 \leq \omega_v \leq 1. \tag{13}$$

where P_x^e and P_x^v are the prices set by intermdeidate inputs exporters in RMB and USD, respectively. ω_v fixes expenditure weights rather than physical input proportions.[21]

A sector-specific fraction $\nu_s \in (0,1]$ of the input bundle's pre-financing price requires external financing before revenues are collected with the remaining fraction $1-\nu_s$ funded internally. The sectoral financing requirement is driven by its technological characteristics, such as cash conversion cycle and time-to-build (Figure 3). I assume that this financing requirement applies uniformly across invoicing currencies of intermediate goods. Therefore, the borrowing currency mechanically mirrors this exogenous currency composition of its intermediate input bundle. This alignment of borrowing shares with input expenditure weights is a maintained financing assumption. Let i_e and i_v denote the corresponding net borrowing rates for RMB and USD over the financing period, respectively. The average borrowing rate is $i_j = (1-\omega_v)i_e + \omega_v i_v \geq 0$. The effective unit price of the intermediate bundle, inclusive of financing costs, is therefore

$$\tilde{P}_X = \bar{P}_X^e(1+\nu_s i_j). \tag{14}$$

Absent an credit shortfall, cost minimization gives the state-contingent marginal cost

$$MC_j^0(\mathbf{X}) = \left(\frac{\tilde{P}_X}{\alpha}\right)^{\alpha}\left(\frac{W}{1-\alpha}\right)^{1-\alpha}, \tag{15}$$

where W represents the nominal wage denominated in RMB. The cost-minimizing factor demands are

$$X_j^0(Y,\mathbf{X}) = \frac{\alpha MC_j^0}{\tilde{P}_X}Y, \qquad L_j^0(Y,\mathbf{X}) = \frac{(1-\alpha)MC_j^0}{W}Y. \tag{16}$$

The working-capital principal requirement is the amount borrowed before production to finance the externally funded share of the intermediate-input bill. Evaluated at the normal-financing input choice,

[21] For $0 < \omega_v < 1$, the normalization $X_j = [x_j^e/(1-\omega_v)]^{1-\omega_v}[x_j^v/\omega_v]^{\omega_v}$ yields the stated unit cost $\bar{P}_X^e(\mathbf{X})$ where x_j^e and x_j^v are quantities of RMB- and the USD-priced inputs.

it is

$$\begin{aligned}\mathcal{W}_j^0(Y,\mathbf{X}) &\equiv \nu_s \bar{P}_X^e(\mathbf{X}) X_j^0(Y,\mathbf{X}) \\ &= \nu_s \bar{P}_X^e \frac{\alpha MC_j^0}{\tilde{P}_X} Y \\ &= \underbrace{\frac{\nu_s \alpha}{1+\nu_s i_j} MC_j^0}_{\kappa_j^0(\mathbf{X})} Y. \end{aligned} \tag{17}$$

The denominator follows from $\tilde{P}_X = \bar{P}_X^e(1+\nu_s i_j)$. Wages are paid after revenues are received and therefore do not enter this advance-financing requirement. Appendix D.1 derives the normal cost function, factor demands, and principal requirement. These are normal-financing benchmarks in the good state. When credit is insufficient, the firm may adjust its input mix, so its actual principal requirement need not equal $\mathcal{W}_j^0$.

4.3 Usable Credit and the Cost of Meeting Demand

The normal-financing benchmark describes how much credit the firm would need to produce at minimum cost. I now allow the credit available at $t = 1$ to fall short of that requirement. The firm responds by combining emergency finance with adjustments to its input mix, while continuing to meet demand.

From the credit limit to usable funds. At $t = 0$, the exporter takes a predetermined credit limit $\bar{\mathcal{B}}_j > 0$ as given. It reflects established borrowing capacity, including firm scale, historical cash flows, and creditworthiness, and remains fixed over the production cycle. The role of earnings in corporate borrowing motivates this specification (Lian and Ma, 2021; Drechsel, 2023).[22]

At $t = 1$, banks may allow the firm to draw only part of this limit. Let Θ_ξ^c denote the share approved and drawable through the lending and payment arrangements associated with invoicing currency c in credit state ξ. Usable credit, measured in RMB, is

$$\mathcal{A}_{j,\xi}^c = \Theta_\xi^c \bar{\mathcal{B}}_j, \qquad 0 \le \Theta_\xi^c \le 1. \tag{18}$$

[22] For U.S. nonfinancial firms, Lian and Ma (2021) estimate that approximately 80% of debt by value is based predominantly on operating cash flows. This evidence motivates the firm-level credit limit; it does not determine the usable-credit shares of exporters.

Thus $\bar{\mathcal{B}}_j$ measures established capacity, while $\mathcal{A}^c_{j,\xi}$ measures the funds available for the current production decision.[23]

The benchmark considers a disruption to dollar-related credit and payment services during which the exporter retains access to RMB financing and settlement. The state-specific shares are

$$\Theta^e_g = \Theta^e_b = \bar{\Theta}^e, \qquad \Theta^v_g = \bar{\Theta}^v, \qquad \Theta^v_b = \underline{\Theta}^v, \tag{19}$$

with

$$0 \leq \underline{\Theta}^v < \bar{\Theta}^e < \bar{\Theta}^v \leq 1. \tag{20}$$

Dollar invoicing therefore provides more usable ordinary credit in the good state and less in the bad state. The normal-time ordering $\bar{\Theta}^v > \bar{\Theta}^e$ represents the complementarity between invoicing and bank financing emphasized by Gopinath and Stein (2021). The reversal in the bad state represents a stress scenario in which disruptions to dollar payment services reduce the willingness or ability of banks to finance dollar-invoiced exports.[24]

This ranking orders credit capacity, rather than profits. Extra capacity lowers costs only when it relaxes a financing shortfall. When ordinary credit is sufficient under both currencies, normal-time differences in profits instead reflect the preset-price problem, including the input-currency matching motive examined below. The unchanged PCP share ($\bar{\Theta}^e$) isolates the specified disruption to dollar services.

Financing the shortfall and adjusting production. For an input quantity X_j, the external principal requirement is $\nu_s \bar{P}^e_X X_j$. The firm draws credit up to $\mathcal{A}$ and obtains any remaining principal through emergency finance. The shortfall is

$$G_j(X_j, \mathcal{A}) = \left[\nu_s \bar{P}^e_X X_j - \mathcal{A}\right]_+, \qquad [z]_+ \equiv \max\{z, 0\}. \tag{21}$$

[23] For an institutional example, see Article 4(aa) of Guideline (EU) 2016/65, as amended by Guideline (EU) 2026/690. It specifies additional valuation markdowns of 16 percent for USD- or GBP-denominated marketable assets and 26 percent for JPY-denominated assets. For the exchange-rate-risk rationale, see Adler et al. (2023). These provisions concern collateral for Eurosystem credit operations and serve as an institutional illustration of currency-dependent financing conditions.

[24] Brüggen et al. (2025) discuss how sanctions can reduce the availability of a settlement currency. DeMarco and Walker (2025) document a decline in Chinese banks' dollar cross-border lending after 2022Q1 alongside evidence of a shift toward RMB lending, and discuss funding costs and geopolitical considerations. These findings motivate the scenario; the credit-access ranking remains a model assumption.

Emergency finance is available at an additional cost $\Psi_s(G)$ above the ordinary financing charges already included in $\tilde{P}_X$. Assume $\Psi_s(0) = \Psi_s'(0) = 0$, $\Psi_s'(G) > 0$ for $G > 0$, and $\Psi_s''(G) \geq 0$, with finite cost for finite G.

The firm can reduce its need for advance funding by using fewer intermediates and more labor, whose wages are paid after revenues arrive. It chooses this input mix after observing both states. For required output Y and macroeconomic state $\mathbf{X}$, the minimum total cost is

$$C_j(Y, \mathbf{X}; \mathcal{A}) = \min_{X_j, L_j \geq 0} \left\{ \tilde{P}_X X_j + W L_j + \Psi_s\big(G_j(X_j, \mathcal{A})\big) : X_j^{\alpha} L_j^{1-\alpha} \geq Y \right\}, \tag{22}$$

where $Y = Y_i^c(P^c, \mathbf{X})$ is fixed by the preset price and the realized macroeconomic state. The credit state changes the cost of fulfilling this demand through both financing charges and the chosen input mix.

Two benchmarks clarify the size of the cost increase. Without a shortfall, cost is $MC_j^0 Y$. If the firm keeps its normal input mix despite a shortfall, it additionally pays the emergency-finance cost on $[\mathcal{W}_j^0(Y, \mathbf{X}) - \mathcal{A}]_+$. Since this input mix remains feasible,

$$MC_j^0 Y \leq C_j(Y, \mathbf{X}; \mathcal{A}) \leq MC_j^0 Y + \Psi_s\Big([\mathcal{W}_j^0(Y, \mathbf{X}) - \mathcal{A}]_+\Big). \tag{23}$$

The actual cost incorporates the firm's ability to adjust inputs. Moreover, more usable credit weakly reduces cost at every input choice and hence also at the optimum. For the same required output and macroeconomic state,

$$C_j(Y, \mathbf{X}; \mathcal{A}_{j,b}^v) \geq C_j(Y, \mathbf{X}; \mathcal{A}_{j,g}^v). \tag{24}$$

Appendix D.2 derives the input choices and these cost properties. They provide the link from a disruption in credit access to the expected profits used in the currency decision.

4.4 Currency Values and the Switching Threshold

The firm now compares the expected profits from RMB and USD invoicing, allowing for the production decisions just described. The key timing restriction is that it chooses one price in the selected currency at $t = 0$, before either state is observed. Throughout this comparison, only the bad-state probability π varies; the macroeconomic distribution and all other model parameters remain fixed.

Expected profits at a common preset price. For either invoicing currency $c \in \{e, v\}$, realized profit in credit state ξ, measured in RMB, equals revenue less minimized production and financing costs:

$$\Pi^c_{j,\xi}(P^c, \mathbf{X}) = a_c P^c Y^c_i(P^c, \mathbf{X}) - C_j\Big(Y^c_i(P^c, \mathbf{X}), \mathbf{X}; \mathcal{A}^c_{j,\xi}\Big), \tag{25}$$

where $a_e = 1$ and $a_v = \mathcal{E}_{ev}$ convert revenue into the same unit of account. For each currency, define expected profit conditional on the credit state at a given preset price as

$$J^c_{j,\xi}(P^c) = \mathbb{E}_{\mathbf{X}}\left[\Pi^c_{j,\xi}(P^c, \mathbf{X})\right], \qquad c \in \{e, v\}, \quad \xi \in \{g, b\}.$$

The expectation averages over the common macroeconomic distribution. For each currency, both credit states are evaluated at the same preset price.

Under PCP, $\mathcal{A}^e_{j,g} = \mathcal{A}^e_{j,b}$. Realized profit is therefore identical across credit states at a given price and macroeconomic state, implying

$$J^e_{j,g}(P^e) = J^e_{j,b}(P^e) \equiv J^e_j(P^e).$$

Under DCP, usable credit differs between the two states, so the state subscript is retained. At a given preset dollar price P and macroeconomic state, revenue and demand are identical across credit states. The expected loss from impaired dollar credit is therefore the increase in minimized total cost:

$$\begin{aligned} \mathcal{L}_j(P) &= J^v_{j,g}(P) - J^v_{j,b}(P) \\ &= \mathbb{E}_{\mathbf{X}}\Big[C_j(Y^v_i(P, \mathbf{X}), \mathbf{X}; \mathcal{A}^v_{j,b}) - C_j(Y^v_i(P, \mathbf{X}), \mathbf{X}; \mathcal{A}^v_{j,g})\Big] \geq 0. \end{aligned} \tag{26}$$

The inequality follows from Equation 24. Because inputs can adjust, $\mathcal{L}_j(P)$ includes the resulting changes in input expenditure as well as emergency-finance charges.

Choosing the price and invoicing currency. Under PCP, usable credit is the same in both credit states, so its optimized value is independent of π. Under DCP, the firm weights the two state-specific

profits by their probabilities and chooses a common price:

$$V_j^e = \max_{P^e} J_j^e(P^e), \tag{27}$$

$$\begin{aligned} V_j^v(\pi) &= \max_{P^v} \left\{ (1-\pi) J_{j,g}^v(P^v) + \pi J_{j,b}^v(P^v) \right\} \\ &= \max_{P^v} \left\{ J_{j,g}^v(P^v) - \pi \mathcal{L}_j(P^v) \right\}. \end{aligned} \tag{28}$$

The last line expresses the trade-off: expected good-state profit minus the probability-weighted loss from impaired credit. Let $P_j^{v,*}(\pi)$ denote an optimal DCP price. It can differ across risk environments, although it remains fixed across realized states within a contract. Assume a fixed compact interval of positive feasible prices and finite, continuous state-specific expected profits. These conditions ensure that the optimized values exist and are continuous (See Appendix D.3).

The exact switching threshold. Define the DCP advantage as $\Delta_j(\pi) = V_j^v(\pi) - V_j^e$. The firm prefers DCP when this difference is positive and PCP when it is negative. Increasing π lowers expected DCP profit at every fixed price, and choosing the best price preserves this ordering.

Proposition 1 (Sanctions risk and currency choice) *The optimized DCP value is continuous, convex, and weakly decreasing in π. If $\Delta_j(0) > 0$ and $\Delta_j(1) < 0$, there is a unique cutoff $\pi_j^* \in (0,1)$ satisfying $\Delta_j(\pi_j^*) = 0$. The firm prefers DCP below this cutoff and PCP above it.*

The endpoint conditions describe firms that prefer dollars without bad-state risk but prefer RMB when the bad state is certain, after optimizing prices in each case. The optimized DCP value is the highest of the fixed-price profit lines in π; this makes it convex and, together with the endpoint signs, gives a unique crossing. A corresponding threshold τ_j^* exists when the range of $\pi(\tau)$ contains π_j^*.

At a unique optimal price, the envelope theorem gives

$$\frac{dV_j^v(\pi)}{d\pi} = -\mathcal{L}_j(P_j^{v,*}(\pi)) \leq 0. \tag{29}$$

Thus the marginal cost of greater sanctions risk is the bad-state loss at the currently optimal price. The decline is strict wherever that loss is positive. The result accounts for price adjustment; it does not require holding the optimal price fixed as risk changes.

An interpretable approximation. Let $P^v_{j,0} = P^{v,*}_j(0)$ be the normal-time DCP price and $N_j = V^v_j(0) - V^e_j > 0$ the normal-time DCP premium. If this price were retained as risk increased, the DCP advantage would be exactly $N_j - \pi \mathcal{L}_j(P^v_{j,0})$. It would disappear at

$$\pi^{\text{lin}}_j = \frac{N_j}{\mathcal{L}_j(P^v_{j,0})}. \tag{30}$$

This ratio compares the normal-time dollar advantage with the loss from a bad credit state at the normal-time price. The firm can also choose a different preset price, so

$$\Delta_j(\pi) \geq N_j - \pi \mathcal{L}_j(P^v_{j,0}), \qquad 0 < \pi^{\text{lin}}_j \leq \pi^*_j < 1$$

under Proposition 1. The ratio is therefore the fixed-price cutoff and a lower bound on the exact cutoff. At a unique normal-time optimum, it also gives the cutoff of the first-order approximation to $\Delta_j(\pi)$ around zero. Appendix D.3 proves these statements. The comparative statics below use the exact cutoff.

4.5 Working Capital and Operational Hedging

Two firm characteristics shape the currency trade-off. Working-capital dependence ν_s determines how much of the input bill must be financed before revenue arrives. The dollar-priced input share ω_v determines how closely input costs move with dollar revenues. To assess either characteristic, one must compare its effect on the normal-time dollar advantage with its effect on the cost of sanctions risk.

Comparing switching cutoffs across firm characteristics. Let z denote either working-capital dependence or the dollar-priced input share, holding the other characteristic and the remaining primitives fixed. The bad-state probability $\pi(\tau)$ is exogenous to the firm and does not depend on z. Firm characteristics instead affect the benefits and costs of DCP at a given probability.

To separate these effects, define

$$\begin{aligned} N_j(z) &= V^v_j(0,z) - V^e_j(z), \\ H_j(\pi,z) &= V^v_j(0,z) - V^v_j(\pi,z), \end{aligned} \tag{31}$$

where $N_j(z)$ is the DCP premium at zero bad-state probability and $H_j(\pi,z)$ is the reduction in opti-

mized DCP value at probability π. The remaining DCP advantage is $\Delta_j(\pi, z) = N_j(z) - H_j(\pi, z)$.

Under the conditions ensuring a unique interior cutoff, let $\pi_j^*(z)$ denote the probability at which firm j is indifferent between DCP and PCP:

$$N_j(z) = H_j(\pi_j^*(z), z).$$

This cutoff depends on z because firm characteristics affect both the initial DCP premium and the value loss from risk. Given sanctions intensity τ, the firm strictly prefers DCP when $\pi(\tau) < \pi_j^*(z)$ and PCP when $\pi(\tau) > \pi_j^*(z)$.

Consider a regular interior cutoff: the values are differentiable nearby and $\mathcal{L}_j(P_j^{v,*}(\pi_j^*), z) > 0$. Differentiating the cutoff equality gives

$$\frac{\partial \pi_j^*}{\partial z} = \frac{\partial_z N_j(z) - \partial_z H_j(\pi_j^*, z)}{\mathcal{L}_j(P_j^{v,*}(\pi_j^*), z)}. \tag{32}$$

Since the denominator is positive, a characteristic raises the cutoff when it adds more to the normal-time premium than to the risk cost. It lowers the cutoff when the risk-cost effect is larger.

Working-capital dependence. A larger externally financed share increases the funds needed for a given input purchase. When dollar credit is impaired, this can make fulfilling demand more costly and induce an earlier switch to RMB.

Proposition 2 (Conditional working-capital amplification) *At a regular interior cutoff, higher working-capital dependence ν_s lowers π_j^* if*

$$\partial_{\nu_s} H_j(\pi_j^*, \nu_s) > \partial_{\nu_s} N_j(\nu_s).$$

One sufficient local case has zero normal net borrowing rates. Let $P_{j,0}^v$ denote the optimal DCP price when $\pi = 0$, and let $P_j^{v,*}(\pi_j^*)$ denote the optimal DCP price when the bad-state probability equals the switching cutoff. Both prices are chosen before the credit state is realized. Assume that, locally, normal credit is sufficient at the PCP optimum. For DCP, assume that good-state credit is sufficient to meet demand at either of these two prices. With input prices, credit limits, and emergency-finance technology fixed, a marginal increase in ν_s then leaves the PCP value and the DCP value at $\pi = 0$ unchanged

($\partial_{\nu_s} N_j(\nu_s) = 0$). Now consider the firm charging $P_j^{v,*}(\pi_j^*)$ when the bad credit state is realized. Suppose that meeting demand requires emergency finance for a set of macroeconomic realizations with positive probability. A higher ν_s raises minimized costs in those realizations, even after inputs adjust. Consequently, $\partial_{\nu_s} H_j > 0 = \partial_{\nu_s} N_j$ at the cutoff, where the partial derivative of H_j holds π fixed. Appendix D.4 establishes this result allowing both prices and inputs to adjust.

Dollar-priced inputs. When more inputs are priced in dollars, an RMB depreciation raises a larger part of the input bill. Dollar invoicing also raises RMB revenue per unit sold in that state, creating an operational hedge. In the normal-time pricing problem, this covariance can strengthen the incentive to use DCP. The firm may then tolerate more sanctions risk before switching.

Proposition 3 (Conditional operational hedging and DCP resilience) *At a regular interior cutoff, a higher dollar-priced input share ω_v raises π_j^* if*

$$\partial_{\omega_v} N_j(\omega_v) > \partial_{\omega_v} H_j(\pi_j^*, \omega_v).$$

This condition compares the additional normal-time dollar advantage with the additional risk cost. A higher ω_v can change both: it changes the RMB input bill, the weighted borrowing rate, and the financing shortfall even when all usable-credit limits are fixed. Thus an unchanged credit-line impairment alone does not imply a higher cutoff. Appendix D.6 derives the normal-time covariance channel and explains how it relates to the level premium N_j; Appendix D.4 proves the cutoff result. Together, the two propositions identify conditions for earlier switching through working-capital needs and later switching through operational hedging.

4.6 Aggregation and the Empirical Exposure Design

The model describes individual Chinese exporters, whereas the invoicing data are observed at the destination-country level. I connect the two by aggregating exporters' currency choices using fixed baseline trade values, consistent with the value-based measurement of import invoicing shares.

For industry s and destination i, let $F_{is}(\pi)$ denote the fraction of baseline imports from China within that industry accounted for by exporters choosing RMB at risk level π. Equivalently, F_{is} is the cumulative distribution of switching thresholds, weighted by exporters' baseline trade values within

the destination-industry pair. Let $\omega_{is,2010}$ denote industry s's share of country i's baseline imports from China, with $\sum_s \omega_{is,2010} = 1$. The implied RMB share within imports from China is

$$q_i^{\text{RMB}}(\pi) = \sum_s \omega_{is,2010} F_{is}(\pi). \tag{33}$$

This share captures currency choices for the baseline import basket.

To express this contribution as a share of total imports, let $E_{i,2010}$ denote baseline China exposure, measured in percent, and hold invoicing on imports from other origins fixed. Let $y_i^{\text{RMB,tar}}$ denote the resulting target RMB share of total imports, before any gradual adjustment. Under the model's mapping from observed sanctions to credit risk, $\pi_t = \pi(\tau_t)$, a change from $\pi_0 = \pi(\tau_0)$ to π_t implies

$$\Delta y_i^{\text{RMB,tar}} = E_{i,2010} \sum_s \omega_{is,2010} \left[F_{is}(\pi_t) - F_{is}(\pi_0) \right]. \tag{34}$$

Because $E_{i,2010}$ is expressed in percent and F_{is} is a fraction, this change is measured in percentage points. On this direct China margin, switching occurs only between RMB and USD, so their share changes are equal and opposite. The empirical coefficients need not satisfy this restriction because the observed outcomes also cover other origins and currencies.

For a local interpretation, suppose F_{is} has a density $f_{is}(u) = F'_{is}(u)$ near the current risk level. For a small probability increase $\delta > 0$, $f_{is}(u)\delta$ approximates the fraction of baseline imports from China within the destination-industry pair that switches to RMB. Holding τ_0 fixed, differentiating Equation 34 gives

$$\frac{\partial \Delta y_i^{\text{RMB,tar}}}{\partial \tau_t} = \frac{\partial y_i^{\text{RMB,tar}}(\tau_t)}{\partial \tau_t} = E_{i,2010} \pi'(\tau_t) \sum_s \omega_{is,2010} f_{is}(\pi(\tau_t)). \tag{35}$$

China exposure scales the directly affected import share, $\pi'(\tau_t)$ measures how sanctions change the bad-state probability, and the weighted density captures the concentration of baseline trade near the current switching threshold. This expression motivates the interaction between baseline China exposure and observed sanctions in Equation 7.

The WCI specification in Equation 9 examines whether the exposure gradient of the sanctions response differs between low-WCI destination countries and the remaining countries. The estimated coefficients are $\theta^{(c)}$ in the remaining group and $\theta^{(c)} + \kappa^{(c)}$ in the low-WCI group. Their difference, $\kappa^{(c)}$, is the main heterogeneity parameter of interest. For RMB invoicing, a negative $\kappa^{(\text{RMB})}$ indicates that

the sanctions response increases less with baseline China exposure in the low-WCI group. The specification also includes $\lambda^{(c)}$ to allow a group-specific component of the sanctions response that does not vary with exposure.

Interpreting this aggregate heterogeneity through the firm's financing mechanism requires attention to the distribution of switching thresholds. Holding exposure and $\pi'(\tau_t)$ fixed, a larger local response requires a greater concentration of baseline trade near the current risk level. Lower thresholds imply earlier RMB adoption, but exporters that have already switched do not switch again after a further sanctions increase. Thus, the firm-level cutoff comparative statics alone do not determine the difference in aggregate exposure gradients. Appendix D.7 formalizes the distinction between earlier adoption and a larger subsequent response. Country-level invoicing shares do not directly reveal individual exporters' thresholds or how working-capital dependence affects them. The WCI interactions therefore measure heterogeneity in aggregate responses.

5 Quantitative Calibration

This section develops a parsimonious quantitative implementation of the switching-threshold mechanism in the theoretical model. The theory characterizes exporters' currency choices through firm-specific sanctions-risk thresholds and links those choices to aggregate invoicing shares using baseline trade-value weights. To implement this aggregation with country-level data, I impose a parametric distribution of switching thresholds whose location depends on the working-capital intensity of each destination' s import basket from China. Baseline China import exposure determines the share of total imports directly affected by currency switching on the China margin, while working-capital intensity enters the distribution of switching thresholds within that margin. This distribution is an additional quantitative assumption. The quantitative mapping therefore combines the scale of China-related trade with heterogeneity in currency-switching incentives.

Using the observed country-year design, I generate RMB invoicing shares and re-estimate the same dynamic panel specifications. The primary targets are the exposure gradient θ, its low-WCI differential κ, and the auxiliary regression's lag coefficient. The exercise evaluates whether this quantitative implementation of the theoretical mechanism can reproduce the main reduced-form patterns.

5.1 Empirical Inputs and the Auxiliary Regression

The calibration uses the same data, estimation sample, WCI classification, and covariates as the RMB specification in Equation 9, with variable notation as in Section 3. For each candidate parameter vector, I generate RMB invoicing shares and re-estimate this specification, replacing both the dependent variable and its lag with their model-generated counterparts. This approach follows Ottonello and Winberry (2020), who estimate their empirical specification on a simulated panel of firms to assess how investment responses to monetary policy vary with firms' financial positions. I apply it to invoicing responses across destinations that differ in China exposure and working-capital intensity. The role of coefficients differs. In their application, the regression coefficients are untargeted and evaluate a model calibrated to other moments. Here, the empirical coefficients are the calibration targets, weighted by the inverse of their country-clustered covariance matrix.

The coefficient κ measures the difference in the exposure gradient of the sanctions response between low-WCI destinations and the remaining countries. Among the latter, a higher baseline China import share is associated with a larger conditional RMB response to sanctions at a rate of θ. For low-WCI destinations, the corresponding gradient is $\theta + \kappa$. A negative κ therefore indicates that the sanctions response increases less with baseline China import exposure in the low-WCI group. In the reported estimates, $\theta + \kappa$ is close to zero, suggesting a nearly flat relationship between baseline exposure and the sanctions response within this group. A near-zero exposure gradient remains compatible with a nonzero sanctions response through the group-specific term $\lambda g_t D_i^{\text{WC}}$. I retain this term in every regression and report λ as an auxiliary diagnostic, while the primary calibration focuses on reproducing the exposure gradients.

5.2 A Bounded Mapping from Switching Thresholds to RMB Shares

Let $e_i = E_{i,2010}/100$, and let r_{i0} and u_{i0} denote initial RMB and USD invoicing shares as fractions. For each country, initialization uses the first year with an observed RMB share from 2010 onward. This is an initial condition, not necessarily a pre-sanctions observation, since some countries enter the sample later. The model is simulated annually, including years with missing observed invoicing shares. I then estimate Equation 9 on the original empirical sample, retaining its consecutive-year observation requirement and using model-generated shares for both the dependent variable and its one-year lag.

For the quantitative implementation, I measure sanctions intensity using the same index as in the empirical analysis, $g_t = \log(1 + S_t^{\text{CN}})$. This index provides the empirical counterpart to τ in the theoretical model. Let $g_{\max} = \max_{2012 \le t \le 2023} g_t$. I map sanctions intensity into the common probability of the bad credit state as

$$\pi_t = 1 - \exp(-\chi g_t), \qquad \chi = \frac{-\log(1-\pi_{\max})}{g_{\max}}. \tag{36}$$

This mapping makes the probability increase with sanctions intensity while remaining between zero and one. It also allows the marginal increase in risk to diminish as sanctions intensify. The parameter χ scales the mapping so that the highest sanctions intensity observed during 2012–2023 corresponds to $\pi_{\max}$. Thus, $\pi_{\max}$ is an assumed probability normalization.

Exporters serving the same destination can differ in their financing conditions and incentives to retain dollar invoicing, so their switching thresholds can differ. Proposition 1 characterizes these thresholds at the exporter level. To represent this heterogeneity quantitatively, I assume a baseline normal distribution of switching thresholds weighted by baseline trade values:

$$m_i = \Lambda\big(a_0 - a_W \text{WC}_{i,2010}\big), \qquad H_i(\pi_t) = \Phi\left(\frac{\pi_t - m_i}{\sigma_\pi}\right), \tag{37}$$

where the logistic function $\Lambda(x) = (1+\exp(-x))^{-1}$ restricts the location parameter m_i to $(0,1)$, placing the center of the baseline threshold distribution within the feasible range of risk probabilities. Φ is the standard normal cumulative distribution function. Thus. evaluated at the current risk probability π_t, $H_i(\pi_t)$ gives the fraction of baseline trade assigned a switching threshold at or below that probability.

The location m_i depends on the working-capital intensity of the destination's 2010 import basket from China and the calibrated parameters, and remains fixed over time. The parameter σ_π is the standard deviation of the baseline normal distribution. With $a_W > 0$, higher working-capital intensity shifts the distribution toward lower thresholds, capturing the conditional tendency toward earlier RMB adoption. The benchmark restriction $a_W \ge 0$ imposes this ordering.

To connect switching on the China margin to aggregate invoicing shares, the benchmark allocates all initial RMB invoicing to imports from China and holds RMB invoicing on imports from other origins at zero. Let q_{i0} denote the model-imputed initial RMB invoicing share within imports from China under

the fixed baseline trade weights. I set

$$r_{i0} = e_i q_{i0}, \qquad q_{i0} = \frac{r_{i0}}{e_i}. \tag{38}$$

The implied share q_{i0} lies between zero and one for every country in the estimation sample. Additional RMB invoicing replaces USD invoicing on imports from China one for one. The switching pool cannot exceed either initially non-RMB imports from China, $e_i(1-q_{i0})$, or the initial aggregate USD share, u_{i0}. I therefore set the share of total imports available for additional switching to

$$c_i = \min\{e_i(1-q_{i0}), u_{i0}\}. \tag{39}$$

I treat this amount as eligible USD-invoiced imports from China. Both the initial RMB allocation and the switching capacity are quantitative assumptions, since the aggregate invoicing data do not identify currency use by origin.

Let t_i^0 denote country i's initialization year, and let $\pi_{t_i^0}$ be the common risk probability in that year. The initial allocation in Equation 38 anchors RMB use to the observed aggregate share r_{i0}. The calibration does not impose $q_{i0} = H_i(\pi_{t_i^0})$: the initial bilateral share is determined by the allocation rule, while H_i provides a baseline distribution for assigning thresholds to additional switching transactions.

The switching capacity c_i, determined from initial currency shares and baseline trade weights, represents trade available for RMB adoption beyond the observed initial share r_{i0}. Initial RMB use is accounted for separately and maintained throughout the simulation. The transactions assigned to c_i must therefore remain outside RMB invoicing at initialization. Thus, the fraction of this capacity whose thresholds are reached at the current risk level π_t is

$$\begin{aligned} G_i(\pi_t) &\equiv \Pr_i\left(\pi^* \le \pi_t \,\middle|\, \pi^* > \pi_{t_i^0}\right) \\ &= \begin{cases} 0, & \pi_t \le \pi_{t_i^0}, \\ \dfrac{H_i(\pi_t) - H_i(\pi_{t_i^0})}{1 - H_i(\pi_{t_i^0})}, & \pi_t > \pi_{t_i^0}, \end{cases} \end{aligned} \tag{40}$$

where a subsequent decline in risk can reduce additional RMB adoption. If risk falls to or below its initial level, the additional target share becomes zero. On the other hand, for risk above its initial level,

the numerator measures the baseline threshold mass reached since initialization. The denominator normalizes this mass relative to all thresholds above initial risk, so that $G_i(\pi_t)$ measures the fraction of the switching pool c_i converted into RMB. Thus, $c_i G_i(\pi_t)$ is the additional RMB target share.

Adding $c_i G_i(\pi_t)$ to initial RMB use gives the target share, expressed in percentage points:

$$y_{it}^{\text{tar}} = 100\{r_{i0} + c_i G_i(\pi_t)\}. \tag{41}$$

For $e_i > 0$, the corresponding RMB share within imports from China is $q_{it}^{\text{tar}} = q_{i0} + (c_i/e_i)G_i(\pi_t)$. Generated RMB shares adjust toward this target according to

$$y_{it}^{\text{mod}} = \rho y_{i,t-1}^{\text{mod}} + (1-\rho) y_{it}^{\text{tar}}, \qquad y_{i,t_i^0}^{\text{mod}} = 100 r_{i0}, \quad 0 \le \rho < 1. \tag{42}$$

The parameter ρ governs adjustment toward the current target: $\rho = 0$ implies immediate adjustment at the annual frequency. Because $0 \le G_i \le 1$, the target remains between $100 r_{i0}$ and $100(r_{i0} + c_i)$. The adjustment rule takes a weighted average of the previous share and this target, so generated RMB shares remain in the same interval.

5.3 Coefficient Matching and Parameter Selection

For given $(\pi_{\max}, \sigma_\pi)$, I choose $v = (a_0, a_W, \rho)$ to approximate $b = (\theta, \kappa, \xi)'$ using

$$Q(v; \pi_{\max}, \sigma_\pi) = \{b^{\text{mod}}(v; \pi_{\max}, \sigma_\pi) - \hat{b}\}' \widehat{V}_b^{-1} \{b^{\text{mod}}(v; \pi_{\max}, \sigma_\pi) - \hat{b}\}, \tag{43}$$

where $\hat{b}$ and $\widehat{V}_b$ are the vector of empirical estimates and the empirical country-clustered covariance matrix, respectively. Matching coefficients obtained from the same auxiliary regression follows the logic of indirect inference (Gourieroux et al., 1993; Smith Jr, 1993). Christiano et al. (2005) estimate a subset of their model's parameters by using a related minimum-distance approach to match empirical and model-implied impulse responses. Equation 43 has the same form, with panel regression coefficients as the matched objects and their full covariance block as the weight.

I evaluate a grid of $(\pi_{\max}, \sigma_\pi)$ pairs, minimizing Equation 43 at each pair from multiple starting values. The benchmark imposes $a_W \ge 0$ to implement the conditional ordering suggested by the theo-

retical mechanism. I select the pair with the smallest stored value of Q.[25]

5.4 Generated Coefficients and the Working-Capital Channel

Table 5 compares the observed regression with the reference and two versions without WCI in the threshold location. The reference reproduces θ, κ, and ξ to numerical precision. Its low-WCI exposure gradient, $\theta+\kappa$, is correspondingly close to zero. This describes how the sanctions response varies with baseline China exposure and does not imply a zero sanctions response for the group.

Table 5: Observed and generated auxiliary-regression coefficients

Statistic	**Data** (1)	**Reference** (2)	**No WCI fixed** (3)	**No WCI re-fitted** (4)
θ	0.014200	0.014200	0.024048	0.011459
	(0.005304)			
κ	-0.013870	-0.013870	0.002101	0.000883
	(0.008325)			
ξ	0.437773	0.437773	0.153060	0.502155
	(0.120671)			
λ	0.142856	0.002884	-0.016122	-0.006430
	(0.064503)			
$\theta+\kappa$	0.000330	0.000330	0.026149	0.012343
Q	—	≈ 0	14.0421	3.1559

Notes: The three targets are θ, κ, and ξ. Country-clustered standard errors are reported in parentheses below the observed-data coefficient estimates. All model columns use the same $(\pi_{\max}, \sigma_\pi)$ pair. The fixed comparison keep a_0 and ρ with $a_W = 0$; the re-fitted comparison re-optimizes a_0 and ρ. All columns retain the empirical WCI group labels. The table reports numerical fit, not a model specification test.

To assess the role of working-capital heterogeneity, I consider two restricted versions of the model that impose $a_W = 0$: one holds the remaining parameters at their reference values and the other re-fits them, as Christiano et al. (2005) do. Under this restriction, the baseline threshold location becomes $m_i = \Lambda(a_0)$ for every destination, so working-capital intensity no longer directly shifts the threshold distribution. Both comparisons retain the reference values of $(\pi_{\max}, \sigma_\pi)$, together with observed China exposure, initial currency shares, initialization years, and the original covariates.

The first comparison, labeled "No WCI: fixed," sets $a_W = 0$ while holding a_0 and ρ at their reference values. It therefore measures how the generated coefficients change when the WCI loading is removed

[25] The nonuniform grid contains 114 values of $\pi_{\max}$ between 0.001 and 0.999 and 36 values of σ_π between 0.01 and 2, giving 4,104 pairs. At each pair, bounded nonlinear least squares uses twelve starting values, with $a_0 \in [-16, 8]$, $a_W \in [0, 15]$, and $\rho \in [0, 0.95]$.

without allowing the remaining parameters to adjust. The second comparison, labeled "No WCI: re-fitted," again imposes $a_W = 0$ but re-optimizes a_0 and ρ to match the same three targets (θ, κ, ξ) using the same objective function. This allows the common threshold location and adjustment persistence to compensate for the absence of the WCI loading. Neither comparison reproduces the negative low-WCI exposure differential at the reference $(\pi_{\max}, \sigma_\pi)$ pair; the re-fitted model has $Q = 3.1559$. These results show that the WCI-dependent location helps reproduce the targeted attenuation under the reference specification.

5.5 Retrospective Temporal Evaluation and Interpretation

I next examine whether including working-capital heterogeneity improves the model's ability to account for subsequent invoicing paths. I calibrate the WCI and no-WCI models separately on 2012–2019 observations and evaluate their generated RMB shares over 2020–2023 for countries represented in the training sample.

Each model selects its parameters and $\pi_{\max}$ by matching the training-period coefficients (θ, κ, ξ), with $a_W = 0$ in the no-WCI model and $\sigma_\pi = 0.4$ in both models. Among candidates with numerically equivalent minimum losses, I select the one with the lowest training-period RMSE after removing country and year effects.

The probability mapping is scaled using the maximum sanctions index observed through 2019. I then hold this mapping and the selected parameters fixed and simulate the subsequent years using realized sanctions. No evaluation-period invoicing outcomes enter parameter selection.

Table 6: Retrospective evaluation of RMB invoicing paths, 2020–2023

Statistic	WCI	No WCI re-fitted	Initial share held
RMSE, levels (pp)	1.0668	1.1152	1.2041
RMSE, country/year effects removed (pp)	0.7005	0.7054	0.7102

Notes: Training uses 2012–2019 and evaluation uses 2020–2023. All columns use the same 286 evaluation observations from 73 countries represented in the training sample. RMSE is in percentage points; lower values indicate smaller discrepancies. The adjusted measure removes country and year indicators using the evaluation sample, without removing the other controls or lagged shares. The initial-share benchmark holds each country's initialization share fixed. The re-fitted no-WCI model selects its parameters and probability normalization separately.

Table 6 shows that the WCI model has lower evaluation-period RMSE than both the re-fitted no-WCI model and the initial-share benchmark. Relative to the re-fitted no-WCI model, level RMSE is

lower by 4.34 percent. After removing country and year effects, RMSE falls from 0.7054 to 0.7005 percentage points, a reduction of 0.70 percent. The latter comparison measures relative path discrepancies. These results provide limited evidence of improved subsequent paths under the stated training procedure and search bounds.

6 Conclusion

This paper examines how U.S. financial sanctions against Chinese entities relate to import invoicing across destinations. Using a panel of FX-neutral invoicing shares from 2012 to 2023, I find that higher sanctions intensity is associated with relatively higher RMB and lower USD shares in countries with greater baseline import exposure to China. The estimated RMB exposure gradient is flatter among destinations whose imports from China have low working-capital intensity. Thus, the evidence connects sanctions-related currency substitution to both trade exposure and the financing requirements of the goods traded.

The theoretical model provides a financing interpretation of these patterns. Chinese exporters balance the dollar's normal-time financing and hedging benefits against the risk that sanctions impair usable credit associated with dollar invoicing. Firms choose both the invoicing currency and a preset price, anticipating how they will adjust financing and production if credit access deteriorates. Greater working-capital dependence lowers the switching threshold when it raises the cost of impaired credit more than the normal-time dollar advantage. Dollar-priced inputs can strengthen the incentive to retain dollar invoicing through operational hedging. Aggregation connects these choices to country-level shares: China exposure scales the affected trade, while the distribution of switching thresholds determines how much trade responds at the sanctions risk level.

The quantitative implementation translates this mechanism into generated RMB invoicing shares and applies the same regression used in the empirical analysis. The reference calibration reproduces the targeted exposure gradients and lag coefficient. At the reference probability normalization and threshold dispersion, removing WCI from the threshold location prevents the model from reproducing the negative low-WCI exposure differential, including after re-fitting the remaining parameters. A separate retrospective exercise calibrates the models on 2012–2019 observations and evaluates 2020–2023 paths conditional on realized sanctions. In that comparison, the WCI model reduces both level

and fixed-effect-adjusted errors relative to the re-fitted no-WCI model and the initial-share benchmark. These results illustrate how working-capital heterogeneity can account for the targeted regression patterns and improve subsequent paths under the specified evaluation design.

The broader implication is that sanctions can change the appeal of an invoicing currency by changing the reliability of the credit and payment services associated with it. Currency substitution can therefore differ across trade relationships facing the same sanctions environment. The incentives to adopt RMB depend on how costly dollar-related financing disruptions are for production and on the availability of alternative financing and settlement arrangements.

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Appendices

A Data and Measurement

Table A.1: Reference Exchange Rates under the Corrected Import-Invoicing Construction

Reference exchange rates	RMB	USD	EUR	Residual
Panel A. Correlation with the 2019 measure: levels				
2007–2010 average	0.9743	0.9787	0.9967	0.9220
2010	0.9772	0.9817	0.9973	0.9328
2019	1.0000	1.0000	1.0000	1.0000
2012–2023 average	0.9998	0.9997	1.0000	0.9990
Panel B. Correlation with the 2019 measure: Country and year fixed effects removed				
2007–2010 average	0.9627	0.8819	0.9620	0.8735
2010	0.9664	0.8915	0.9700	0.8805
2019	1.0000	1.0000	1.0000	1.0000
2012–2023 average	0.9997	0.9988	0.9995	0.9976
Observations	819	819	640	819
Countries	87	87	69	87

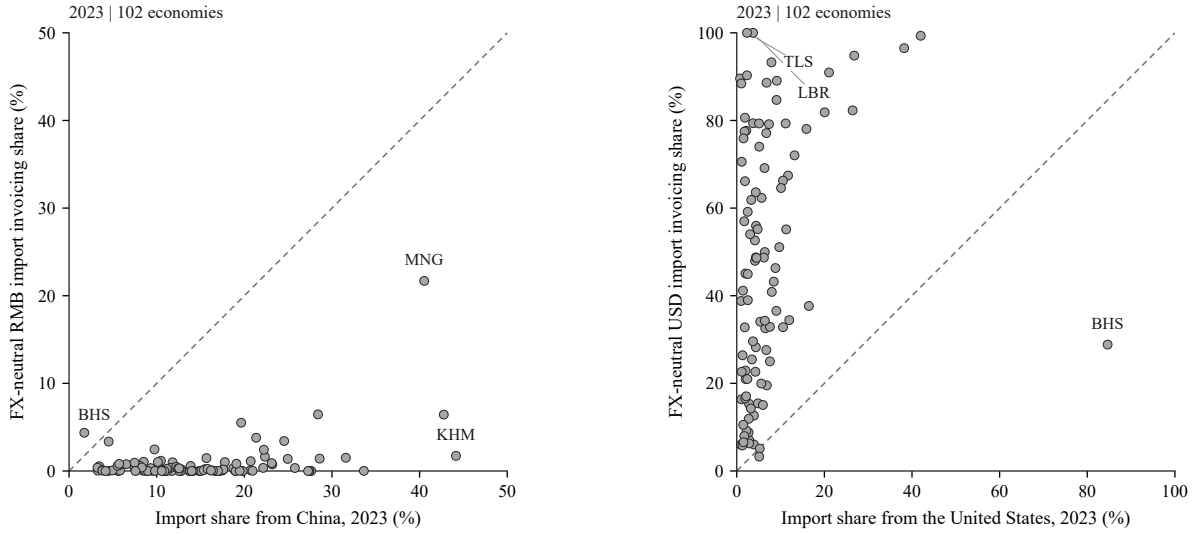
Notes: All correlations are Pearson correlations with the corresponding 2019-reference measure. Period reference rates are arithmetic means of annual currency-per-USD exchange rates, requiring complete observations throughout the reference window. Panel A uses invoicing-share levels. Panel B uses residuals from separate least-squares projections of each measure on country and year indicators; these are not residuals from the fully controlled regression. The country–year sample is identical across all reference choices and both the original and corrected invoicing constructions, within each currency, over 2012–2023.

Table A.2: Summary Statistics of Invoicing Currency Shares: Raw vs. FX-neutral

	Raw						**FX-neutral**					
	Export			Import			Export			Import		
Statistic	USD	EUR	RMB	USD	EUR	RMB	USD	EUR	RMB	USD	EUR	RMB
N (Obs.)	1,080	1,080	1,080	1,159	1,159	1,159	1,080	1,080	1,080	1,159	1,159	1,159
N (Countries)	117	117	117	121	121	121	117	117	117	121	121	121
Mean	52.13	35.08	0.34	50.84	34.23	0.46	52.58	34.80	0.33	51.42	33.88	0.45
SD	33.97	35.54	1.90	28.72	32.11	1.90	33.93	35.38	1.87	28.75	31.94	1.82
Min	0.00	0.00	0.00	0.67	0.00	0.00	0.00	0.00	0.00	0.69	0.00	0.00
P25	20.24	1.83	0.00	23.93	4.49	0.00	20.96	1.83	0.00	25.09	4.36	0.00
Median	47.50	15.38	0.00	48.00	20.87	0.02	48.11	15.17	0.00	48.26	20.41	0.02
P75	87.17	73.26	0.04	78.60	66.85	0.25	87.80	72.65	0.04	79.12	65.19	0.24
Max	100.00	99.46	31.20	100.00	95.26	32.43	100.00	99.43	31.71	100.00	95.50	29.61

Notes: This table reports pooled, unweighted country–year summary statistics over 2012–2023. Within each trade flow, the raw and FX-neutral statistics for USD, EUR and RMB use exactly the same country–year observations, requiring nonmissing values for all six measures. The export and import samples can differ and are not restricted to the regression sample. Raw shares are measured before the exchange-rate valuation adjustment. FX-neutral shares use 2019 reference exchange rates to reduce mechanical exchange-rate valuation effects, treating the residual invoicing category as the reporting country's home currency. Euro-area countries are included, and their euro invoicing is reported under EUR. SD denotes the sample standard deviation. All share statistics are expressed in percent. Source: Boz et al. (2025) and the author's calculations.

Figure A.1: Import Invoicing Shares and Bilateral Trade Exposure (2023)

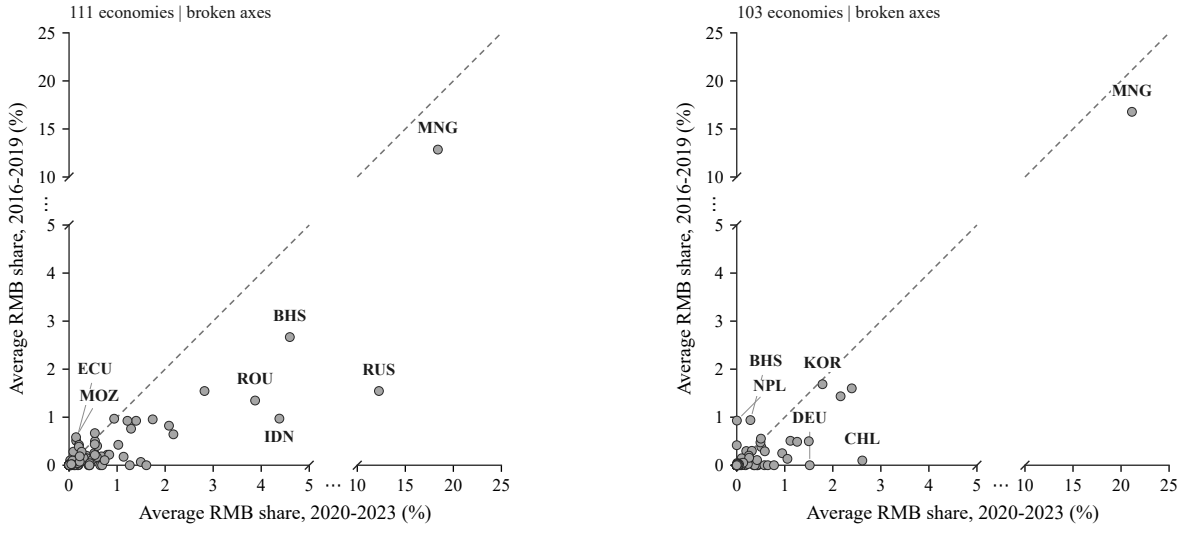


(a) RMB Invoicing vs. Trade Share

(b) USD Invoicing vs. Trade Share

Notes: This figure compares bilateral trade integration and currency invoicing shares for China (Panel a) and the U.S. (Panel b) in 2023. The horizontal axis represents the destination country's import share from China or the U.S., while the vertical axis represents the corresponding FX-neutral RMB or USD invoicing share of total imports from all trading partners, using 2019 exchange rates. The dashed 45-degree line benchmarks a one-to-one proportional relationship where currency usage perfectly matches trade share. In Panel a, RMB usage disproportionately lags behind actual bilateral trade integration. Conversely, in Panel b, the vast majority of countries are positioned well above the 45-degree line. This stark asymmetry visually corroborates the USD's pervasive dominance as a global vehicle currency.

Figure A.2: Changes in FX-neutral RMB invoicing shares: Full Sample



(a) Import invoicing

(b) Export invoicing

Notes: The figure compares country-level average RMB invoicing shares in 2016–2019 and 2020–2023. Panel a reports import invoicing shares, and Panel b reports export invoicing shares. Each point represents an economy with at least one non-missing observation in both periods. The invoicing shares are adjusted using 2019 exchange rates as described in equation (1). The dashed 45-degree line represents no change between the two periods; observations below the line experienced an increase in their average RMB invoicing share. Selected economies are labeled by their ISO country codes. The underlying invoicing data are from Boz et al. (2025).

Table A.3: China-Related Policy Indicators by Country

		FTA	RMB clearing bank	PBOC swap line				
ISO3	Economy	Effective year	Effective year	First month	Initial ceiling	Dec. 2023 ceiling	Ceiling changes (month: bn RMB)	Positive periods
ALB	Albania	–	–	2013-09	2	2	2013-09: 2	2013-09–2024-12
ARE	United Arab Emirates	–	2016	2012-01	35	35	2012-01: 35	2012-01–2024-12
ARG	Argentina	–	2015	2009-04	70	130	2009-04: 70; 2018-12: 130	2009-04–2024-12
ARM	Armenia	–	–	2015-03	1	–	2015-03: 1	2015-03–2018-02
AUS	Australia	2015	2014	2012-03	200	200	2012-03: 200	2012-03–2024-12
AUT	Austria	–	–	2013-10[a]	–	–	–	2013-10–2024-12
BEL	Belgium	–	–	2013-10[a]	–	–	–	2013-10–2024-12
BLR	Belarus	–	–	2009-03	20	–	2009-03: 20; 2015-05: 7	2009-03–2012-02; 2015-05–2021-04
BRA	Brazil	–	2023	2013-03	190	–	2013-03: 190	2013-03–2016-02
BRN	Brunei Darussalam	2005	–	–	–	–	–	–
CAN	Canada	–	2014	2014-11	200	200	2014-11: 200	2014-11–2024-12
CHE	Switzerland	2014	2015	2014-07	150	150	2014-07: 150	2014-07–2024-12
CHL	Chile	2006	2015	2015-05	22	50	2015-05: 22; 2020-07: 50	2015-05–2024-12
CRI	Costa Rica	2011	–	–	–	–	–	–
CYP	Cyprus	–	–	2013-10[a]	–	–	–	2013-10–2024-12
DEU	Germany	–	2014	2013-10[a]	–	–	–	2013-10–2024-12
ECU	Ecuador	2024	–	–	–	–	–	–
EGY	Egypt	–	–	2016-12	18	18	2016-12: 18	2016-12–2024-12
ESP	Spain	–	–	2013-10[a]	–	–	–	2013-10–2024-12
EST	Estonia	–	–	2013-10[a]	–	–	–	2013-10–2024-12
FIN	Finland	–	–	2013-10[a]	–	–	–	2013-10–2024-12
FRA	France	–	2014	2013-10[a]	–	–	–	2013-10–2024-12
GBR	United Kingdom	–	2014	2013-06	200	350	2013-06: 200; 2015-10: 350	2013-06–2024-12
GEO	Georgia	2018	–	–	–	–	–	–
GRC	Greece	–	–	2013-10[a]	–	–	–	2013-10–2024-12
HKG	Hong Kong SAR, China	2004	2003	2009-01	200	800	2009-01: 200; 2011-11: 400; 2020-11: 500; 2022-07: 800	2009-01–2024-12
HRV	Croatia	–	–	2023-01[a]	–	–	–	2023-01–2024-12
HUN	Hungary	–	2015	2013-09	10	40	2013-09: 10; 2019-09: 20; 2020-09: 40	2013-09–2024-12

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Table A.3 continued

ISO3	Economy	FTA	RMB clearing bank	PBOC swap line				
		Effective year	Effective year	First month	Initial ceiling	Dec. 2023 ceiling	Ceiling changes (month: bn RMB)	Positive periods
IDN	Indonesia	2005	–	2009-03	100	250	2009-03: 100; 2015-11: 130; 2018-11: 200; 2021-11: 250	2009-03–2024-12
IRL	Ireland	–	–	2013-10[a]	–	–	–	2013-10–2024-12
ISL	Iceland	2014	–	2010-06	3.5	3.5	2010-06: 3.5	2010-06–2024-12
ITA	Italy	–	–	2013-10[a]	–	–	–	2013-10–2024-12
JPN	Japan	2022	2018	2018-10	200	200	2018-10: 200	2018-10–2024-12
KAZ	Kazakhstan	–	2022	2011-06	7	–	2011-06: 7	2011-06–2021-04
KHM	Cambodia	2005	2024	–	–	–	–	–
KOR	Korea, Rep.	2015	2014	2009-04	180	400	2009-04: 180; 2011-10: 360; 2020-10: 400	2009-04–2024-12
LAO	Lao PDR	2005	2022	2020-05	6	6	2020-05: 6	2020-05–2024-12
LKA	Sri Lanka	–	–	2014-09	10	10	2014-09: 10	2014-09–2024-12
LTU	Lithuania	–	–	2015-01[a]	–	–	–	2015-01–2024-12
LUX	Luxembourg	–	2014	2013-10[a]	–	–	–	2013-10–2024-12
LVA	Latvia	–	–	2014-01[a]	–	–	–	2014-01–2024-12
MAC	Macao SAR, China	2004	2004	2019-12	30	30	2019-12: 30	2019-12–2024-12
MAR	Morocco	–	–	2016-05	10	–	2016-05: 10	2016-05–2019-04
MLT	Malta	–	–	2013-10[a]	–	–	–	2013-10–2024-12
MMR	Myanmar	2005	–	–	–	–	–	–
MNG	Mongolia	–	–	2011-05	5	15	2011-05: 5; 2012-03: 10; 2014-05: 15	2011-05–2024-12
MUS	Mauritius	2021	–	2024-09	2	–	2024-09: 2	2024-09–2024-12
MYS	Malaysia	2005	2015	2009-02	80	180	2009-02: 80; 2012-02: 180	2009-02–2024-12
NGA	Nigeria	–	–	2018-04	15	15	2018-04: 15	2018-04–2024-12
NIC	Nicaragua	2024	–	–	–	–	–	–
NLD	Netherlands	–	–	2013-10[a]	–	–	–	2013-10–2024-12
NZL	New Zealand	2008	–	2011-04	25	25	2011-04: 25	2011-04–2024-12
PAK	Pakistan	2007	2022	2011-12	10	30	2011-12: 10; 2017-12: 20; 2020-07: 30	2011-12–2024-12
PER	Peru	2010	–	–	–	–	–	–
PHL	Philippines	2005	2019	–	–	–	–	–
PRT	Portugal	–	–	2013-10[a]	–	–	–	2013-10–2024-12
QAT	Qatar	–	2014	2014-11	35	35	2014-11: 35	2014-11–2024-12

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Table A.3 continued

ISO3	Economy	FTA	RMB clearing bank	PBOC swap line				
		Effective year	Effective year	First month	Initial ceiling	Dec. 2023 ceiling	Ceiling changes (month: bn RMB)	Positive periods
RUS	Russian Federation	–	2016	2014-10	150	150	2014-10: 150	2014-10–2024-12
SAU	Saudi Arabia	–	–	2023-11	50	50	2023-11: 50	2023-11–2024-12
SGP	Singapore	2005	2013	2010-07	150	300	2010-07: 150; 2013-03: 300	2010-07–2024-12
SRB	Serbia	2024	2024	2016-06	1.5	–	2016-06: 1.5	2016-06–2019-05
SUR	Suriname	–	–	2015-03	1	–	2015-03: 1	2015-03–2022-02
SVK	Slovakia	–	–	2013-10[a]	–	–	–	2013-10–2024-12
SVN	Slovenia	–	–	2013-10[a]	–	–	–	2013-10–2024-12
THA	Thailand	2005	2015	2011-12	70	70	2011-12: 70	2011-12–2024-12
TJK	Tajikistan	–	–	2015-09	3	–	2015-09: 3	2015-09–2018-08
TUR	Türkiye	–	–	2012-02	10	35	2012-02: 10; 2015-02: 12; 2021-06: 35	2012-02–2024-12
TWN	Taiwan, China	2010	2012	–	–	–	–	–
UKR	Ukraine	–	–	2012-06	15	–	2012-06: 15	2012-06–2021-11
USA	United States	–	2016	–	–	–	–	–
UZB	Uzbekistan	–	–	2011-04	0.7	–	2011-04: 0.7	2011-04–2014-03
VNM	Viet Nam	2005	–	–	–	–	–	–
ZAF	South Africa	–	2015	2015-04	30	30	2015-04: 30	2015-04–2024-12
ZMB	Zambia	–	2015	–	–	–	–	–

Notes: The table combines the three China-related policy variables used in the empirical analysis. FTA entry-into-force dates are obtained from China's Ministry of Commerce, while bilateral swap-line agreements and RMB clearing-bank designations are compiled from PBoC announcements. The reported FTA and RMB clearing-bank effective years reproduce the first year in which the corresponding indicator switches to one. First month represents the earliest month having a positive recorded ceiling, and ceiling amounts are in billions of RMB. Renewals are coded retroactively from the expiration month of the preceding agreement, whereas mid-contract ceiling increases are coded in their announcement month. In the panel, the FTA and clearing-bank indicators equal one from the reported effective year onward, while the annual swap-line indicator equals one when the year-end monthly facility is active. The December 2023 ceiling is the amount observed at the end of the estimation period, not a separate index. Positive Periods report contiguous months with an active facility, so a gap may represent expiration, non-renewal, or missing source coverage. The table reports effective dates and monthly swap observations through 2024; dates after 2024 are omitted. A dash ordinarily denotes no positive or coded observation in the supplied inputs. Countries absent from the table equal zero on all three indicators. [a] The PBoC and ECB first established their bilateral currency swap arrangement on October 8, 2013, with a common maximum size of RMB 350 billion or EUR 45 billion. The same common ceiling remained in force as of December 31, 2023. For each euro-area member, First month is the later of October 2013 and its euro-adoption month; country-specific ceiling and change cells are shown as dashes because the common ceiling cannot be allocated across member countries.

Table A.4: Summary Statistics for the Regression Variables

Variable	N	Mean	SD	Median	Min	Max
Panel A. Invoicing outcomes and their lags						
RMB invoicing share (%)	831	0.322	0.834	0.027	0.000	9.156
Lagged RMB invoicing share (%)	831	0.249	0.660	0.014	0.000	9.156
USD invoicing share (%)	831	50.493	30.022	46.087	4.547	99.764
Lagged USD invoicing share (%)	831	51.275	29.966	46.743	4.547	99.764
EUR invoicing share (%)	640	27.316	28.415	12.479	0.236	95.453
Lagged EUR invoicing share (%)	640	26.995	28.021	12.661	0.236	95.453
Residual invoicing share (%)	831	11.753	18.872	0.000	0.000	81.192
Lagged Residual invoicing share (%)	831	11.416	18.521	0.000	0.000	82.274
Panel B. Time-varying regressors and interactions						
Destination sanctions, $d_{i,t}$	831	0.427	0.839	0.000	0.000	4.745
FTA indicator	831	0.167	0.373	0.000	0.000	1.000
RMB swap-line indicator	831	0.236	0.425	0.000	0.000	1.000
RMB clearing-bank indicator	831	0.164	0.370	0.000	0.000	1.000
Lagged China import share (%)	831	12.552	7.489	10.193	1.420	41.513
Lagged homogeneous-goods import share (%)	831	18.554	7.951	17.925	0.762	52.032
Lagged political-alignment indicator	831	0.735	0.441	1.000	0.000	1.000
China exposure $\times$ sanctions, $E_{i,2010}g_t$	831	30.744	23.582	26.132	0.000	173.017
Low WCI $\times$ sanctions, $D_i^{\mathrm{WC}}g_t$	831	0.048	0.403	0.000	0.000	5.056
Triple interaction, $E_{i,2010}D_i^{\mathrm{WC}}g_t$	831	0.366	3.035	0.000	0.000	35.704
Panel C. Predetermined characteristics: one observation per country						
2010 China import exposure, $E_{i,2010}$ (%)	88	9.573	5.912	8.309	1.326	34.218
Standardized 2010 WCI	88	0.131	0.840	0.175	-1.282	1.867
Panel D. Common sanctions series: one observation per year						
China sanctions, $g_t = \ln(1+S_t^{\mathrm{CN}})$	12	2.967	1.361	3.111	0.000	5.056
China-directed SDN designations, S_t^{CN}	12	36.250	43.124	22.000	0.000	156.000

Notes: This table summarizes the variables used in the fully controlled baseline and working-capital heterogeneity regressions over 2012–2023. Panel A uses each currency's estimation sample: 831 country–year observations from 88 countries for RMB, USD and the residual category, and 640 observations from 69 countries for EUR, which excludes euro-area countries. Panel B uses the common RMB/USD/residual estimation sample. Lagged variables are summarized at the values entering the regressions; FTA, swap-line and clearing-bank indicators enter contemporaneously. Panel C counts each of the 88 countries once, and Panel D counts each of the 12 years once. Predetermined country characteristics are therefore not weighted by the number of observed years, and common sanctions measures are not weighted by the number of countries observed in a year. All statistics are unweighted. Invoicing shares and import shares are expressed in percent; indicators take values zero or one. Sanctions intensity is the natural logarithm of one plus annual SDN designations, including entities and individuals. SD denotes the sample standard deviation. Outcome rows retain the stored invoicing construction used in the reported regressions and the existing conditional zero-filling of missing RMB shares; they differ from the corrected descriptive shares in Table A.2.

Table A.5: Alternative Measures of Industry Working-Capital Requirements

Alternative measure	N	Pearson	Spearman
Industry median, 2010. Median of firms' raw NOWC-to-sales ratios.	28	0.948	0.863
Inventory-to-sales proxy, 2010. Industry median of firms' raw inventories relative to sales (Manova et al., 2015).	28	0.939	0.845
Pre-period median, 2007–2010. For each firm, take the median raw NOWC-to-sales ratio over available years (at least two); then take the industry median of these firm medians.	27	0.942	0.866
Exclude IPO-year observations. Recompute the 2010 revenue-weighted mean after excluding firms listed during 2010; retain the baseline winsorization cutoffs.	27	0.992	0.973
All-industry winsorization cutoffs. Recompute the 2010 revenue-weighted mean using 1st/99th-percentile cutoffs estimated from all mapped industries, including services.	28	0.991	1.000

Notes: Each row compares the alternative with the baseline revenue-weighted mean of firm-level NOWC-to-sales ratios in 2010, winsorized at the pooled 1st and 99th percentiles of eligible goods-sector firms. Correlations are unweighted across 2-digit ISIC Revision 3.1 industries and use pairwise-complete observations. N is the number of paired industries; Spearman correlations use average ranks for ties. The baseline covers 28 goods industries and 1,214 firms. The pre-period median and IPO-year exclusion comparisons each omit ISIC division 37, for which the alternative measure is unavailable. These are industry-level measurement comparisons.

B Additional Panel Results and Sensitivity Checks

Table B.1: US Financial Sanctions on Chinese Entities and Global Invoicing Currencies: LSDVC Estimates

	RMB			**USD**			**EUR**			**Residual**		
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
Lag. Dep. Var.	0.546***	0.542***	0.526***	0.620***	0.621***	0.627***	0.733***	0.728***	0.722***	0.636***	0.636***	0.642***
	(0.043)	(0.044)	(0.043)	(0.040)	(0.040)	(0.041)	(0.047)	(0.047)	(0.046)	(0.041)	(0.041)	(0.042)
ChineseRisk	0.016***	0.015***	0.013***	−0.049	−0.055	−0.057	0.018	0.020	0.023	0.044	0.045	0.041
	(0.004)	(0.004)	(0.004)	(0.036)	(0.037)	(0.037)	(0.023)	(0.024)	(0.025)	(0.040)	(0.041)	(0.042)
Idio. Sanc.	−0.027	−0.025	−0.025	−0.184	−0.177	−0.146	0.268	0.273	0.252	0.175	0.173	0.178
	(0.038)	(0.038)	(0.039)	(0.320)	(0.322)	(0.335)	(0.242)	(0.245)	(0.252)	(0.361)	(0.363)	(0.375)
LR Effects	0.035***	0.034***	0.028***	−0.129	−0.146	−0.152	0.067	0.072	0.081	0.121	0.124	0.115
	(0.009)	(0.010)	(0.009)	(0.094)	(0.097)	(0.102)	(0.089)	(0.088)	(0.090)	(0.110)	(0.114)	(0.118)
Contemp. Ctrl.		✓	✓		✓	✓		✓	✓		✓	✓
Lagged Ctrl.			✓			✓			✓			✓
Year FE	✓	✓	✓	✓	✓	✓	✓	✓	✓	✓	✓	✓
Obs	800	800	785	800	800	785	623	623	608	800	800	785

Notes: This table presents robustness checks addressing potential Nickell bias in dynamic panels by employing the bias-corrected least squares dummy variable (LSDVC) estimator. The bias correction is initialized using the Arellano-Bond estimator. The dependent variables are the FX-neutralized import invoicing shares of the respective currencies (RMB, USD, EUR, and Residual). **ChineseRisk** is the interaction between a destination country's 2010 baseline import share from China and the log-transformed global U.S. sanctions against Chinese entities. **Idio. Sanc.** denotes the log-transformed idiosyncratic U.S. sanctions directly targeting the destination country. **LR Effects** reports the long-run multiplier of the **ChineseRisk** shock, computed via the Delta method based on the lagged dependent variable (**Lag. Dep. Var.**). Contemporaneous controls encompass indicators for bilateral FTAs, currency swap lines, and official RMB clearing banks. Lagged controls include the import share from China, the import share of homogeneous goods, and a political alignment index. All specifications include year fixed effects, and country fixed effects are inherently absorbed by the LSDVC procedure. Standard errors, obtained via bootstrapping with 500 repetitions, are reported in parentheses. * p<0.10, ** p<0.05, *** p<0.01.

Table B.2: Baseline Estimates under Alternative Country Exclusions: All Currencies

	RMB			USD			EUR			Residual		
	Original	No Russia	Excl. five	Original	No Russia	Excl. five	Original	No Russia	Excl. five	Original	No Russia	Excl. five
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
ChineseRisk	0.014***	0.014***	0.015***	-0.062**	-0.062**	-0.052*	0.018	0.021	0.024	0.048	0.045	0.028
	(0.005)	(0.005)	(0.006)	(0.027)	(0.027)	(0.026)	(0.022)	(0.021)	(0.019)	(0.038)	(0.037)	(0.034)
Idio. Sanc.	-0.024	-0.025	-0.026	-0.152	-0.112	-0.261	0.270	0.404**	0.466**	0.183	0.045	0.150
	(0.043)	(0.045)	(0.048)	(0.264)	(0.276)	(0.282)	(0.213)	(0.197)	(0.204)	(0.278)	(0.261)	(0.270)
Lag. Dep. Var.	0.438***	0.433***	0.428***	0.514***	0.515***	0.513***	0.565***	0.561***	0.550***	0.501***	0.499***	0.487***
	(0.121)	(0.120)	(0.119)	(0.064)	(0.065)	(0.070)	(0.076)	(0.077)	(0.080)	(0.105)	(0.108)	(0.119)
p: $\theta = 0$	0.0090	0.0097	0.0094	0.0256	0.0245	0.0509	0.4232	0.3319	0.2292	0.2170	0.2277	0.4043
Observations	831	821	784	831	821	784	640	630	593	831	821	784
Countries	88	87	83	88	87	83	69	68	64	88	87	83
Within R^2	0.194	0.188	0.189	0.281	0.281	0.279	0.328	0.327	0.322	0.252	0.250	0.235

Notes: Within each currency, the three columns report the fully controlled baseline specification in Equation 7 using the original estimation sample ("Original"), excluding Russia only ("No Russia"), and jointly excluding Russia, Brazil, India, Turkey, and Kazakhstan ("Excl. five"). The estimation period is 2012–2023; Russian observations cover 2012–2021. The EUR regressions exclude euro-area countries under the baseline sample rule. The dependent variables are FX-neutral import invoicing shares, expressed in percentage points. ChineseRisk$_{i,t} = E_i g_t$, where E_i is the fixed 2010 China import share, in percent, and g_t is the natural logarithm of one plus annual China-directed U.S. SDN designations, including entities and individuals. **Idio. Sanc.** is the corresponding log-plus-one measure of destination-directed designations. All columns include the lagged dependent variable, destination-directed sanctions, contemporaneous FTA, RMB swap-line and RMB clearing-bank indicators, lagged China import share, homogeneous-goods import share, political alignment, and country and year fixed effects. Standard errors in parentheses are based on country-clustered covariance matrices. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

Table B.3: Working-Capital Heterogeneity: Excluding Russia

	RMB		USD		EUR		Residual	
	Full	Excl. Russia	Full	Excl. Russia	Full	Excl. Russia	Full	Excl. Russia
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
θ: $E_i g_t$	0.014***	0.014***	-0.062**	-0.062**	0.018	0.021	0.047	0.045
	(0.005)	(0.005)	(0.027)	(0.027)	(0.022)	(0.021)	(0.038)	(0.038)
κ: $E_i g_t D_i$	-0.014*	-0.016*	0.326***	0.328***	-0.219***	-0.212***	-0.047	-0.055
	(0.008)	(0.008)	(0.091)	(0.095)	(0.077)	(0.076)	(0.103)	(0.101)
λ: $g_t D_i$	0.143**	0.165***	-2.066**	-2.081**	1.727**	1.633**	-0.060	0.031
	(0.065)	(0.062)	(0.879)	(0.930)	(0.825)	(0.807)	(0.985)	(0.972)
Observations	831	821	831	821	640	630	831	821
Countries	88	87	88	87	69	68	88	87

Notes: Each pair compares the fully controlled WCI specification in Equation 9 using the original sample and a sample excluding Russia only. The estimation period is 2012–2023; Russia contributes observations through 2021. The dependent variables are FX-neutral import invoicing shares, in percentage points. E_i is the fixed 2010 China import share, measured in percent; g_t is the natural logarithm of one plus annual China-directed U.S. SDN designations, including entities and individuals. D_i identifies the bottom decile of the original WCI reference distribution. Exposure and group assignments are fixed. All columns retain country and year fixed effects, the lagged dependent variable, destination-directed sanctions, contemporaneous FTA, RMB swap-line and RMB clearing-bank indicators, and lagged China import share, homogeneous-goods import share, and political alignment. Euro-area countries are excluded from the EUR regressions. $\Delta R(10) = \lambda + 10\kappa$ is the low-WCI-minus-remaining-group contrast at a 10 percent China import share, in percentage points per one-unit increase in g_t. Standard errors are clustered by country. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

Table B.4: Robustness to the IHS Transformation of Sanctions Measures

	RMB			USD			EUR			Residual		
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
ChineseRisk (IHS)	0.015***	0.014***	0.012***	-0.047**	-0.052**	-0.055**	0.008	0.011	0.014	0.047	0.047	0.044
	(0.005)	(0.005)	(0.005)	(0.021)	(0.023)	(0.024)	(0.023)	(0.021)	(0.020)	(0.031)	(0.034)	(0.035)
Idio. Sanc. (IHS)	-0.020	-0.018	-0.018	-0.150	-0.144	-0.123	0.239	0.241	0.225	0.140	0.141	0.142
	(0.034)	(0.034)	(0.036)	(0.212)	(0.210)	(0.217)	(0.159)	(0.163)	(0.170)	(0.214)	(0.216)	(0.224)
Lag. Dep. Var.	0.460***	0.456***	0.440***	0.509***	0.510***	0.514***	0.579***	0.572***	0.565***	0.496***	0.496***	0.500***
	(0.140)	(0.136)	(0.121)	(0.061)	(0.062)	(0.064)	(0.078)	(0.079)	(0.076)	(0.098)	(0.101)	(0.105)
LR Effects	0.027***	0.026***	0.022***	-0.095**	-0.107**	-0.112***	0.020	0.025	0.032	0.093*	0.093	0.089
	(0.009)	(0.009)	(0.007)	(0.039)	(0.042)	(0.043)	(0.056)	(0.050)	(0.048)	(0.056)	(0.062)	(0.063)
Contemp. Ctrl.		✓	✓		✓	✓		✓	✓		✓	✓
Lagged Ctrl.			✓			✓			✓			✓
Year and Cty FE	✓	✓	✓	✓	✓	✓	✓	✓	✓	✓	✓	✓
Observations	849	849	831	849	849	831	658	658	640	849	849	831
Countries	91	91	88	91	91	88	72	72	69	91	91	88
Within R^2	0.173	0.174	0.192	0.278	0.279	0.281	0.327	0.328	0.328	0.252	0.252	0.252

Notes: This table replaces the log-plus-one transformation with the inverse hyperbolic sine (IHS) transformation for both sanctions measures in the baseline dynamic panel regressions. For a designation count S, $\text{asinh}(S) = \ln(S+\sqrt{S^2+1})$. ChineseRisk (IHS) is $E_{i,2010}\,\text{asinh}(S_t^{\text{CN}})$, where $E_{i,2010}$ is the fixed 2010 China import share, measured in percent. Idio. Sanc. (IHS) is $\text{asinh}(S_{i,t}^{\text{D}})$. Counts include annual new designations of entities and individuals. Outcomes are FX-neutral import invoicing shares, in percentage points. Samples and controls match the corresponding baseline columns over 2012–2023; euro-area countries are excluded from the EUR regressions. Contemporaneous controls are FTA, RMB swap-line and RMB clearing-bank indicators. Lagged controls are the China import share, commodity-import share and political alignment. All columns include the lagged outcome and country and year fixed effects. LR Effects is the ChineseRisk coefficient divided by one minus the lagged-outcome coefficient, with delta-method standard errors. Coefficient standard errors in parentheses are clustered by country. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

C Complementary Dynamic Evidence

This appendix complements the panel regressions with two exercises describing the timing of invoicing changes. The first reports year-specific differences in RMB invoicing between countries with high and low baseline exposure to China. The second uses local projections to examine horizon-specific relationships between China-directed sanctions pressure, baseline exposure, and invoicing changes. These exercises organize the temporal patterns in the data.

C.1 Year-Specific RMB Invoicing Differences by China Exposure

I examine how the RMB import invoicing gap between high- and low-exposure countries evolves over 2012–2023. Baseline exposure is measured by the share of imports from China in 2010. I retain countries at or above the 75th-percentile cutoff and at or below the 25th-percentile cutoff, excluding countries between the cutoffs. Let H_i equal one for high-exposure countries and zero for low-exposure countries. Exposure is fixed at its 2010 value, and the group assignment is held constant over time. I interact year indicators with an indicator for the high-exposure group and normalize the 2021 coefficient to zero. The resulting estimates describe the adjusted high–low invoicing-share gap in each year relative to the gap in 2021.

I estimate

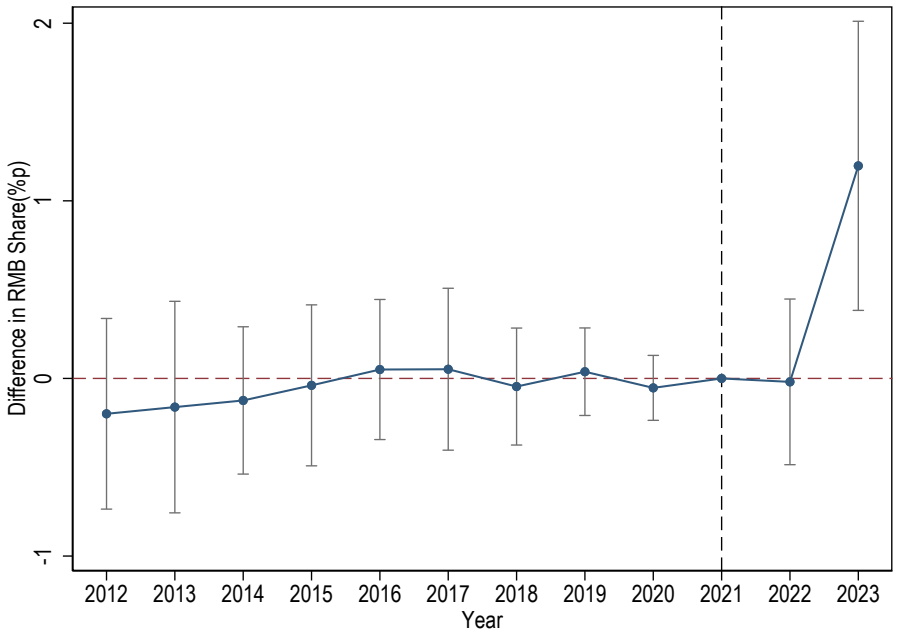
$$y_{it}^{\text{RMB}} = \alpha_i + \beta_t + \sum_{\substack{\tau \in \mathcal{T} \\ \tau \neq 2021}} b_\tau H_i \mathbf{1}\{t = \tau\} + \gamma' X_{i,t-1} + \psi' Z_{it} + \delta d_{it} + \varepsilon_{it},$$

where $\mathcal{T} = \{2012, \ldots, 2023\}$ and y_{it}^{RMB} is the FX-neutral RMB import invoicing share, expressed in percentage points. The covariates $X_{i,t-1}$ and Z_{it} and the destination-directed sanctions measure d_{it} are defined as in Equation 7. The specification includes country and year fixed effects but no lagged dependent variable. Standard errors are clustered by country.

The interaction for 2021 is omitted, normalizing $b_{2021} = 0$. Country fixed effects absorb permanent differences across countries, including the main effect of H_i, while year fixed effects absorb changes common to both exposure groups. Thus, b_τ measures the change in the covariate-adjusted high-minus-low invoicing gap between 2021 and year τ. A positive coefficient indicates that this gap is larger in year τ than in 2021. This specification allows the exposure-group gap to vary freely across years rather than linking its evolution to measured China-directed sanctions pressure. The coefficients therefore describe an annual exposure profile.

Figure C.1 shows a pronounced positive difference in 2023 relative to the 2021 reference year. This pattern documents the timing of the divergence between the exposure groups, but does not separately identify sanctions as its cause. Other contemporaneous developments may have affected the groups differently. As a sensitivity check, I also include the lagged RMB invoicing share. On a common estimation sample, the year-specific profile and the positive differential in 2023 remain qualitatively similar.

Figure C.1: Year-Specific RMB Invoicing Differences by China Exposure



Notes: This figure plots the year specific estimates of the difference in the RMB invoicing share across two groups: high– and low–exposure to China. Points are coefficients on interactions between year indicators and the high-exposure-group indicator, expressed in percentage points. The omitted reference year is 2021. The regression includes country and year fixed effects, destination-directed sanctions, FTA, swap-line and RMB clearing-bank indicators, and lagged measures of the China import share, homogeneous-goods import share, and political alignment. The lagged dependent variable is not included. Vertical bars show pointwise 90% confidence intervals based on standard errors clustered by country.

C.2 Local Projections of Invoicing Changes

Figure C.1 describes when invoicing gaps emerge, without linking those differences directly to measured sanctions pressure. I complement that exercise with local projections (LPs) following Jordà (2005). These projections examine how the relationship between China-directed sanctions pressure and invoicing changes varies across horizons and with baseline exposure to China. Unlike the long-run multiplier calculated from the dynamic panel model, each horizon is estimated separately rather than obtained by recursively applying a common adjustment coefficient.[26]

Baseline specification. For each currency $c \in \{\text{RMB}, \text{USD}, \text{EUR}, \text{Residual}\}$ and horizon $h \in \{0,1,2,3\}$, I estimate

$$y^{(c)}_{i,t+h} - y^{(c)}_{i,t-1} = \alpha^{(c)}_{i,h} + \beta^{(c)}_{t,h} + \xi^{(c)}_{h} y^{(c)}_{i,t-1} + \gamma^{(c)\prime}_{h} X_{i,t-1} + \psi^{(c)\prime}_{h} Z_{it} + \delta^{(c)}_{h} d_{it} + \theta^{(c)}_{h} E_i g_t + u^{(c)}_{i,t+h} \tag{C.1}$$

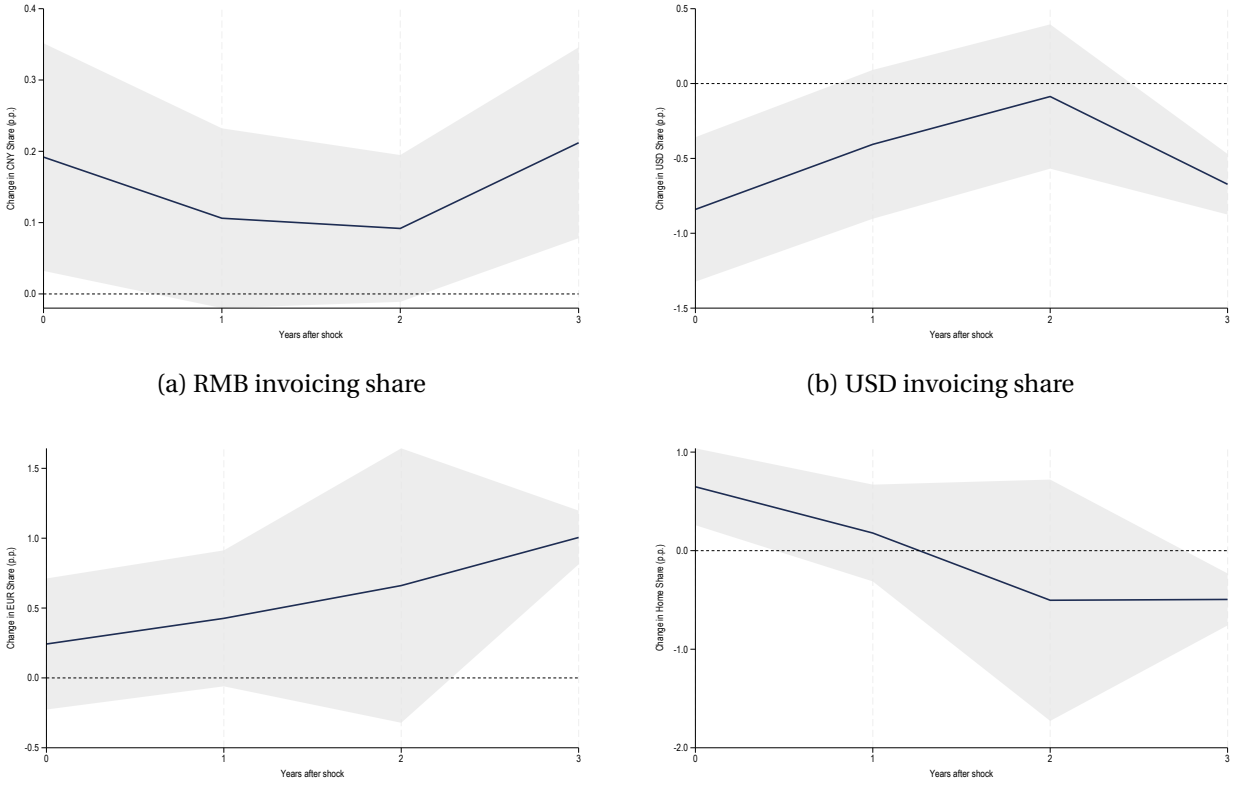
[26] Unlike traditional Vector Autoregression (VAR) models, the LP method directly estimates the impact of a shock at time t on the outcome at horizon $t+h$ through a series of separate horizon-specific regressions, making it significantly more robust to dynamic misspecification.Jordà (2005); Jordà and Taylor (2025)

where $y_{it}^{(c)}$ is the FX-neutral import invoicing share, expressed in percentage points, and E_i is the country's import share from China in 2010, expressed in percent. The sanctions measure g_t is the logarithm of one plus annual China-directed SDN designations. The destination-directed sanctions measure d_{it} and the covariates $X_{i,t-1}$ and Z_{it} follow the definitions in Equation 7.

The dependent variable $(y_{i,t+h}^{(c)} - y_{i,t-1}^{(c)})$ is the cumulative change in the invoicing share between $t-1$ and $t+h$. At $h=0$, it measures the change from $t-1$ to t; at $h=3$, it measures the change from $t-1$ to $t+3$. Each projection includes its own country and year fixed effects and conditions on the initial invoicing share $y_{i,t-1}^{(c)}$. Since g_t is common across countries, its main effect is absorbed by year fixed effects. The coefficient $\theta_h^{(c)}$ captures how the conditional relationship between current sanctions pressure and the subsequent invoicing-share change varies with baseline China exposure.

Figure C.2 reports $10 \times 1.36 \times \widehat{\theta}_h^{(c)}$, in percentage points. This scaling evaluates the exposure-related component at a benchmark China import share of 10 percent and a 1.36-unit increase in g_t, approximately one standard deviation in the reference sample. It does not recover the common effect of sanctions absorbed by year fixed effects or the total predicted invoicing change for an individual country.

Figure C.2: Exposure-Related Dynamics in Import Invoicing Shares



(a) RMB invoicing share

(b) USD invoicing share

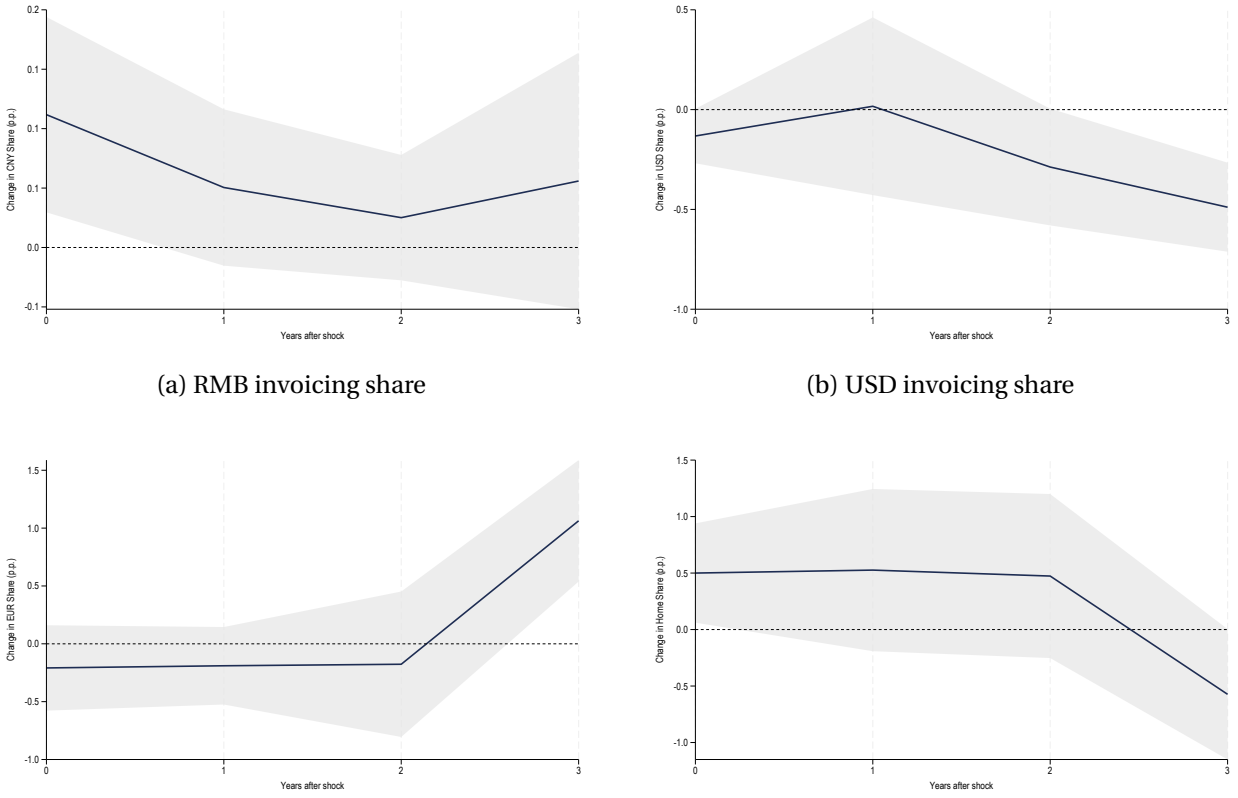
(c) EUR invoicing share

(d) Residual invoicing share

Notes: Each panel reports $10 \times 1.36 \times \widehat{\theta}_h^{(c)}$ from Equation C.1, in percentage points. Each horizon is estimated separately with country and year fixed effects, the lagged dependent variable, destination-directed sanctions, and the baseline covariates. Shaded areas denote 90 percent pointwise normal confidence intervals constructed using Driscoll–Kraay standard errors with bandwidth $h+1$. Samples vary across currencies and horizons. Euro-area countries are excluded from the EUR specification.

The RMB point estimates are positive at all reported horizons. The displayed confidence intervals exclude zero at $h = 0$ and $h = 3$, where the scaled estimates are approximately 0.19 and 0.21 percentage points, respectively. The USD estimates are negative at all horizons, with confidence intervals excluding zero at $h = 0$ and $h = 3$ and corresponding estimates of approximately -0.84 and -0.67 percentage points. The EUR estimates are positive, with the confidence interval excluding zero only at $h = 3$. The residual-currency estimate is positive at $h = 0$ and negative at $h = 3$, with both confidence intervals excluding zero. Several intermediate-horizon estimates are imprecise. These results describe currency-specific conditional patterns. Because the estimation samples can differ across currencies and horizons, the coefficients should not be interpreted as an exact accounting decomposition of invoicing reallocation. As a sensitivity check, Figure C.3) reports projections using reference years 2012–2019. This exercise assesses whether the estimated relationships also appear before the most recent sanctions escalation. Even in this truncated sample, the immediate surge in RMB usage remains robustly identified.

Figure C.3: Exposure-Related Dynamics in Import Invoicing Shares (2012-2019)



(a) RMB invoicing share

(b) USD invoicing share

(c) EUR invoicing share

(d) Residual invoicing share

Notes: Each panel reports $10 \times 1.36 \times \widehat{\theta}_h^{(c)}$ from Equation C.1 with subsample (2012-2019), in percentage points. Each horizon is estimated separately with country and year fixed effects, the lagged dependent variable, destination-directed sanctions, and the baseline covariates. Shaded areas denote 90 percent pointwise normal confidence intervals constructed using Driscoll–Kraay standard errors with bandwidth $h+1$. Samples vary across currencies and horizons. Euro-area countries are excluded from the EUR specification.

Heterogeneity by working-capital intensity. I next examine whether the relationship between sanctions pressure and invoicing changes differs with the working-capital intensity of imports from China ($\text{WC}_{i,2010}$). This comparison uses a different grouping from the high- and low-China-exposure comparison. Let D_i equal one for countries in the upper quartile of baseline working-capital intensity and zero for countries in the lower quartile. Countries in the middle two quartiles are excluded. The comparison uses the precomputed quartile indicators for the 2010 index.

To allow the coefficients on the controls to differ across groups, define

$$Q_{it}^{(c)} = \left(y_{i,t-1}^{(c)}, X_{i,t-1}', Z_{it}', d_{it}\right)'.$$

For each currency c and horizon $h \in \{0,1,2,3\}$, I estimate

$$\begin{aligned} y_{i,t+h}^{(c)} - y_{i,t-1}^{(c)} = \alpha_{i,h}^{(c)} + \beta_{t,h}^{(c)} + \eta_{L,h}^{(c)\prime} Q_{it}^{(c)} + \eta_{\Delta,h}^{(c)\prime} \left(D_i Q_{it}^{(c)}\right) \\ + \theta_{L,h}^{(c)} E_i g_t + \theta_{\Delta,h}^{(c)} D_i E_i g_t + \lambda_h^{(c)} D_i g_t + v_{i,t+h}^{(c)}. \end{aligned} \tag{C.2}$$

The coefficients on the lagged dependent variable, destination-directed sanctions, and all other included controls may differ between the two groups. Country fixed effects absorb the time-invariant terms E_i, D_i, and $D_i E_i$, while common year fixed effects absorb the main effect of g_t. The interaction $D_i g_t$ is not absorbed by these fixed effects and is included explicitly.

At a given baseline exposure E, the sanctions-related slope is $E\theta_{L,h}^{(c)}$ for the low-WCI group and $E(\theta_{L,h}^{(c)} + \theta_{\Delta,h}^{(c)}) + \lambda_h^{(c)}$ for the high-WCI group. Thus, $\theta_{\Delta,h}^{(c)}$ captures the difference in the exposure gradient, while $\lambda_h^{(c)}$ allows the two slopes to differ independently of exposure. The regressions use each country's observed 2010 China import exposure. I evaluate both groups at $E = 10$ percent and use the same 1.36-unit increase in g_t as in the baseline LPs. The plotted quantities in Figure C.4 are

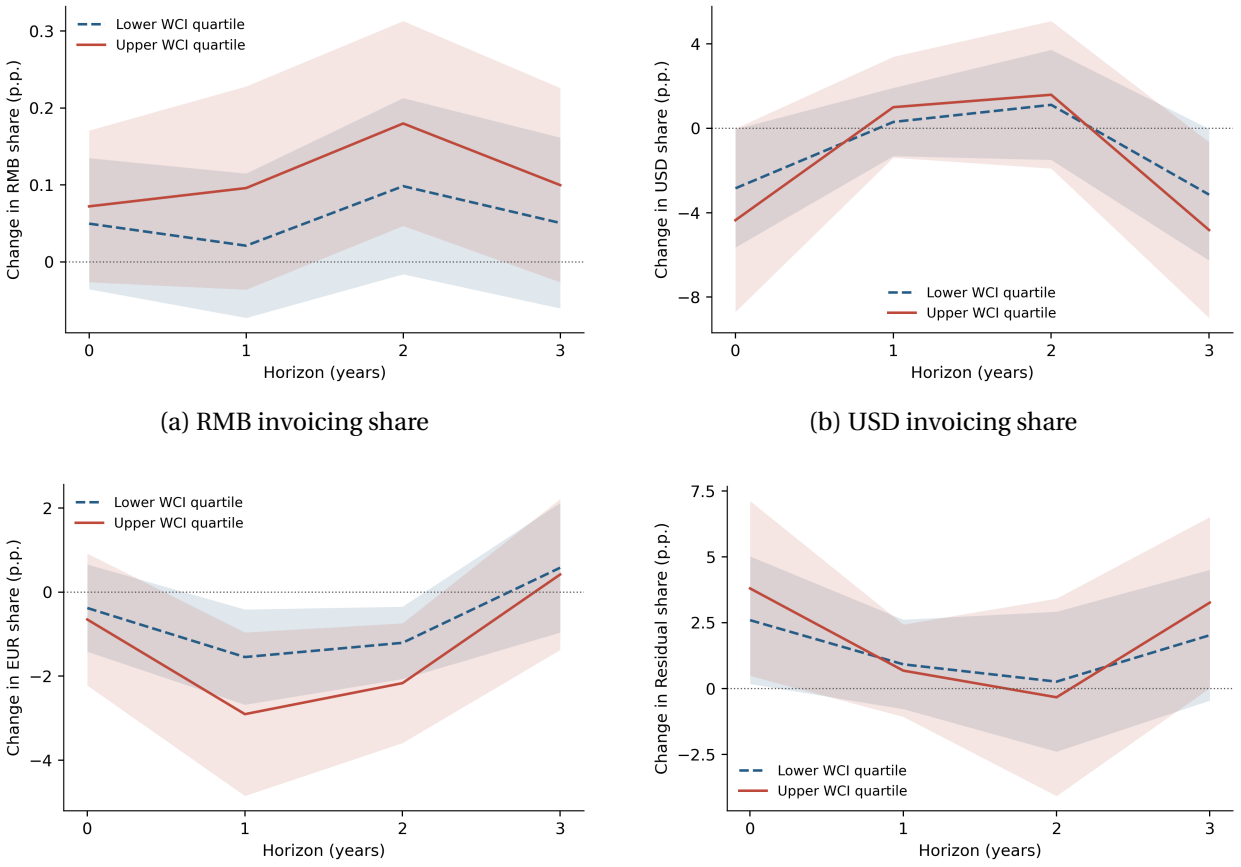
$$\widehat{m}_{L,h}^{(c)} = 1.36\left(10\hat{\theta}_{L,h}^{(c)}\right), \tag{C.3}$$

$$\widehat{m}_{H,h}^{(c)} = 1.36\left[10\left(\hat{\theta}_{L,h}^{(c)} + \hat{\theta}_{\Delta,h}^{(c)}\right) + \hat{\lambda}_h^{(c)}\right]. \tag{C.4}$$

These quantities are expressed in percentage points. Using a common exposure benchmark avoids attributing differences in average China exposure to differences in WCI. The term involving $\lambda_h^{(c)}$ is scaled by 1.36. The plotted quantities summarize the estimated sanctions-related components.

The high-WCI group has positive RMB point estimates throughout the reported horizons, exceeding those of the low-WCI group. However, the high-WCI RMB confidence interval excludes zero only at $h = 2$. At $h = 0$, the high-WCI USD estimate is negative and the residual-currency estimate is positive, with their displayed confidence intervals excluding zero. The EUR estimate is negative at $h = 0$ and positive at $h = 3$, but both intervals include zero. These patterns provide a horizon-specific description of invoicing relationships across WCI groups.

Figure C.4: Invoicing Dynamics by Baseline Working-Capital Intensity



(a) RMB invoicing share

(b) USD invoicing share

(c) EUR invoicing share

(d) Residual invoicing share

Notes: The solid red and dashed blue lines report $\widehat{m}^{(c)}_{H,h}$ and $\widehat{m}^{(c)}_{L,h}$ from Equation C.4 and Equation C.3, respectively. Both are evaluated at a 10 percent baseline China import share and a 1.36-unit increase in g_t, and are expressed in percentage points. The high-WCI estimate includes $1.36\widehat{\lambda}^{(c)}_{h}$, the component associated with $D_i g_t$. The sample retains the upper and lower quartiles of baseline WCI. All included control coefficients, including that on the lagged dependent variable, may differ by group. The specification includes country fixed effects and common year fixed effects. Shaded areas denote 90 percent pointwise normal confidence intervals based on Driscoll–Kraay covariance estimates with bandwidth $h+1$. Each interval incorporates the covariances among the coefficients in the corresponding linear combination. Euro-area countries are excluded from the EUR specification. Samples vary across currencies and horizons.

D Model and Quantitative Details

D.1 Normal Financing Costs and Working-Capital Requirements

This appendix derives the normal-financing benchmarks in Equation 15, Equation 16, and Equation 17. Fix firm j, sector s, macroeconomic state $\mathbf{X}$, and required output $Y > 0$. In the application, $Y = Y_i^c(P^c, \mathbf{X})$ is determined by the preset price fixed in the chosen invoicing currency c and the realized macroeconomic state. Throughout this derivation, ordinary credit is sufficient to finance the normal input choice in the good state. All prices are positive, and state arguments are suppressed where this does not create ambiguity.

For an input quantity X_j, the purchase bill before financing charges is $\bar{P}_X^e X_j$. The firm borrows the fraction ν_s of this bill, so the required principal is

$$B_j(X_j) = \nu_s \bar{P}_X^e X_j. \tag{D.1}$$

Under the maintained quoted borrowing-rate convention i_j, interest is $i_j B_j(X_j)$. The internally funded input expenditure and repayment of principal plus interest therefore sum to

$$\begin{aligned}(1-\nu_s)\bar{P}_X^e X_j + (1+i_j)B_j(X_j) &= \bar{P}_X^e X_j + i_j \nu_s \bar{P}_X^e X_j \\ &= \underbrace{\bar{P}_X^e (1+\nu_s i_j)}_{\tilde{P}_X} X_j.\end{aligned} \tag{D.2}$$

Thus input purchases are counted once, and borrowing adds an interest charge. This accounting follows the model's convention that the internally funded portion carries no additional financing charge. Wages are paid at $t = 2$ and require no advance financing. The derivation takes the input bundle and its quoted financing terms as specified in the main text.

Normal total cost is obtained from

$$C_j^0(Y, \mathbf{X}) = \min_{X_j, L_j \geq 0} \left\{ \tilde{P}_X X_j + W L_j : X_j^\alpha L_j^{1-\alpha} \geq Y \right\}. \tag{D.3}$$

For $Y > 0$, both inputs must be positive. Because both input prices are positive, the output constraint binds at the optimum. The Lagrangian is

$$\mathcal{L} = \tilde{P}_X X_j + W L_j + \lambda_j \left(Y - X_j^\alpha L_j^{1-\alpha} \right).$$

The first-order conditions are

$$\tilde{P}_X = \lambda_j \alpha X_j^{\alpha-1} L_j^{1-\alpha} = \lambda_j \alpha \frac{Y}{X_j}, \tag{D.4}$$

$$W = \lambda_j (1-\alpha) X_j^\alpha L_j^{-\alpha} = \lambda_j (1-\alpha) \frac{Y}{L_j}. \tag{D.5}$$

Rearranging gives

$$X_j^0 = \frac{\lambda_j \alpha}{\tilde{P}_X} Y, \qquad L_j^0 = \frac{\lambda_j (1-\alpha)}{W} Y. \tag{D.6}$$

Substituting these expressions into the binding production constraint yields

$$\begin{aligned} Y &= \left(\frac{\lambda_j \alpha}{\tilde{P}_X} Y \right)^{\alpha} \left(\frac{\lambda_j (1-\alpha)}{W} Y \right)^{1-\alpha} \\ &= \lambda_j Y \left(\frac{\alpha}{\tilde{P}_X} \right)^{\alpha} \left(\frac{1-\alpha}{W} \right)^{1-\alpha}. \end{aligned}$$

Dividing by Y and solving for the multiplier gives

$$\lambda_j = \left(\frac{\tilde{P}_X}{\alpha} \right)^{\alpha} \left(\frac{W}{1-\alpha} \right)^{1-\alpha} \equiv MC_j^0(\mathbf{X}). \tag{D.7}$$

The first-order conditions also imply $\tilde{P}_X X_j^0 = \alpha \lambda_j Y$ and $W L_j^0 = (1-\alpha)\lambda_j Y$. Adding them gives

$$C_j^0(Y, \mathbf{X}) = MC_j^0(\mathbf{X}) Y. \tag{D.8}$$

Because technology has constant returns to scale, unit cost and marginal cost coincide. The normal factor demands are consequently

$$X_j^0(Y, \mathbf{X}) = \frac{\alpha MC_j^0}{\tilde{P}_X} Y, \tag{D.9}$$

$$L_j^0(Y, \mathbf{X}) = \frac{(1-\alpha) MC_j^0}{W} Y. \tag{D.10}$$

To obtain the advance-financing requirement, evaluate the principal definition at this input choice:

$$\begin{aligned} \mathcal{W}_j^0(Y, \mathbf{X}) &\equiv B_j\big(X_j^0(Y, \mathbf{X})\big) \\ &= \nu_s \bar{P}_X^e X_j^0(Y, \mathbf{X}) \\ &= \nu_s \bar{P}_X^e \frac{\alpha MC_j^0}{\tilde{P}_X} Y \\ &= \frac{\nu_s \alpha}{1 + \nu_s i_j} MC_j^0 Y \equiv \kappa_j^0(\mathbf{X}) Y. \end{aligned} \tag{D.11}$$

Hence κ_j^0 is the required working-capital principal per unit of output, measured in RMB. The factor $1/(1+\nu_s i_j)$ converts financing-inclusive input expenditure into the corresponding purchase expenditure before financing charges. Principal is distinct from interest $i_j \mathcal{W}_j^0$ and gross loan repayment $(1+i_j)\mathcal{W}_j^0$. These results describe the unconstrained normal benchmark; the usable-credit problem subsequently allows input choices to respond to a financing shortfall.

D.2 Production and Financing after the Credit State Is Observed

This appendix derives the cost function used in Section 4.3. Fix required output $Y > 0$, macroeconomic state $\mathbf{X}$, and usable ordinary credit $\mathcal{A}$. Write $p_X = \bar{P}_X^e$, $\tilde{P}_X = p_X(1+\nu_s i_j)$, and $r = \alpha/(1-\alpha) > 0$. All monetary costs and credit amounts are measured in RMB.

Optimal inputs after credit is observed. Because costs increase with either input, the production constraint binds. Solving it for labor gives $L = Y^{1/(1-\alpha)}X^{-r}$, so the cost problem becomes

$$C_j(Y,\mathbf{X};\mathcal{A}) = \min_{X>0}\left\{\tilde{P}_X X + WY^{1/(1-\alpha)}X^{-r} + \Psi_s([\nu_s p_X X - \mathcal{A}]_+)\right\}. \tag{D.12}$$

The objective tends to infinity as X approaches zero or infinity. The labor-cost term is strictly convex in X, and the emergency-finance term is convex. Hence there is a unique, positive, finite optimal input quantity X^*. Since $\Psi_s'(0) = 0$, the objective is differentiable even where the shortfall first becomes positive. Its first-order condition is

$$\tilde{P}_X + \nu_s p_X \Psi_s'(G^*) = \frac{\alpha}{1-\alpha}\frac{WL^*}{X^*}. \tag{D.13}$$

Here $L^* = Y^{1/(1-\alpha)}(X^*)^{-r}$ and $G^* = [\nu_s p_X X^* - \mathcal{A}]_+$. The left side is the marginal cost of intermediates, including the extra financing cost they create. The right side is the labor cost saved by using more intermediates. When ordinary credit covers the normal principal requirement $\mathcal{W}_j^0(Y,\mathbf{X})$, the solution is (X^0, L^0) and cost is $MC_j^0 Y$. When it falls short, the normal input choice has a positive additional marginal financing cost; the optimum has $X^* < X^0$, $L^* > L^0$, and $G^* > 0$.

Cost bounds and the value of usable credit. Emergency finance adds a nonnegative cost to the normal production problem, giving $C_j \geq MC_j^0 Y$. Evaluating the full objective at (X^0, L^0) gives the upper bound in Equation 23. Optimizing inputs can lower this upper bound but cannot eliminate the cost of a binding shortfall.

Increasing $\mathcal{A}$ weakly lowers the cost of every feasible input pair. Minimization therefore preserves the ordering. The cost envelope gives

$$\frac{\partial C_j(Y,\mathbf{X};\mathcal{A})}{\partial \mathcal{A}} = -\Psi_s'(G^*) \leq 0. \tag{D.14}$$

The inequality is strict when emergency finance is used. This proves the good-state/bad-state cost ordering in Equation 24. It compares minimized total costs, including input substitution; a difference in emergency-finance charges alone does not generally equal the difference in total cost.

The ordinary credit draw is $\min\{\nu_s p_X X^*, \mathcal{A}\}$ and emergency principal is G^*. Their sum is the external principal requirement. Normal financing charges apply to that requirement through $\tilde{P}_X$, while Ψ_s is an additional charge. Input purchases, principal, and normal interest are consequently counted once.

D.3 Currency Values, Exact Cutoffs, and the Local Approximation

This appendix follows the sequence in Section 4.4: optimize over a common preset price, establish the switching cutoff, and then relate it to the fixed-price approximation. The macroeconomic distribution, credit limits, and remaining primitives are fixed as π changes.

The pricing problem and regularity conditions. For each invoicing currency c, let the feasible price set be $K_c = [\underline{P}_c, \overline{P}_c]$, where $0 < \underline{P}_c < \overline{P}_c < \infty$. Assume that expected profits in each credit state are finite and continuous in the preset price.

For DCP, expected profit at a given preset price is

$$F_j(P,\pi) = J^v_{j,g}(P) - \pi \mathcal{L}_j(P), \qquad V^v_j(\pi) = \max_{P \in K_v} F_j(P,\pi).$$

For each $\pi \in [0,1]$, continuity of the objective and compactness of the price set ensure that at least one feasible price $P^{v,*}_j(\pi)$ achieves the maximum.

The expected bad-state loss $\mathcal{L}_j(P)$ is also continuous in price, since it is the difference between the two state-specific expected-profit functions. It therefore has a finite upper bound over feasible prices:

$$\overline{\mathcal{L}}_j = \max_{P \in K_v} \mathcal{L}_j(P) < \infty.$$

This bound applies regardless of which price the firm selects.

Proof of Proposition 1. For each fixed price, $F_j(P,\pi)$ is affine and weakly decreasing in π, because $\mathcal{L}_j(P) \geq 0$. Maximization over the same price set preserves this ordering. If $\pi_2 > \pi_1$ and P_1, P_2 are optimal at the two probabilities, evaluating each objective at the other optimal price gives

$$(\pi_2 - \pi_1)\mathcal{L}_j(P_2) \leq V^v_j(\pi_1) - V^v_j(\pi_2) \leq (\pi_2 - \pi_1)\mathcal{L}_j(P_1).$$

where

$$\begin{aligned} V^v_j(\pi_1) &\geq J^v_{j,g}(P_2) - \pi_1 \mathcal{L}_j(P_2) = V^v_j(\pi_2) + (\pi_2 - \pi_1)\mathcal{L}_j(P_2), \\ V^v_j(\pi_2) &\geq J^v_{j,g}(P_1) - \pi_2 \mathcal{L}_j(P_1) = V^v_j(\pi_1) - (\pi_2 - \pi_1)\mathcal{L}_j(P_1). \end{aligned}$$

The value change is therefore bounded in absolute value by $\overline{\mathcal{L}}_j|\pi_2 - \pi_1|$, establishing continuity.

For $\lambda \in [0,1]$, affinity at every fixed price implies

$$\begin{aligned} V^v_j(\lambda\pi_1 + (1-\lambda)\pi_2) &= \max_P \{\lambda F_j(P,\pi_1) + (1-\lambda)F_j(P,\pi_2)\} \\ &\leq \lambda V^v_j(\pi_1) + (1-\lambda)V^v_j(\pi_2). \end{aligned}$$

Thus V^v_j is convex. The right side permits a separate price in each risk environment, while the left side uses one price for the weighted objective.

Since V_j^e is constant in π, Δ_j inherits continuity, convexity, and weak monotonicity. The endpoint signs

$$\Delta_j(0) > 0, \qquad \Delta_j(1) < 0.$$

ensure an interior zero. To establish uniqueness, suppose there were two zeros $u < w < 1$. Convexity would require

$$0 = \frac{\Delta_j(w) - \Delta_j(u)}{w-u} \leq \frac{\Delta_j(1) - \Delta_j(w)}{1-w} < 0,$$

a contradiction. Continuity and the endpoint signs then give the preference ordering on either side of the unique cutoff.

The marginal and cumulative costs of risk. When expected profits are differentiable in price and the interior optimal-price rule is differentiable in π, the chain rule gives

$$\begin{aligned}\frac{dV_j^v}{d\pi} &= \left[J_{j,g}^{v\prime}(P_j^{v,*}) - \pi \mathcal{L}_j'(P_j^{v,*})\right] \frac{dP_j^{v,*}}{d\pi} - \mathcal{L}_j(P_j^{v,*}) \\ &= -\mathcal{L}_j(P_j^{v,*}).\end{aligned}$$

The bracket is zero by the price first-order condition. More generally, the value envelope gives the same derivative at a unique optimum without requiring a differentiable price rule.

To translate the marginal effect into a total value loss, define $H_j(\pi) \equiv V_j^v(0) - V_j^v(\pi)$. It measures the reduction in maximized DCP profit as the bad-state probability rises from zero to π. The value bound above implies absolute continuity. Integrating the envelope derivative therefore gives

$$H_j(\pi) = \int_0^\pi \mathcal{L}_j(P_j^{v,*}(u))\, du. \tag{D.15}$$

At almost every u, the integrand equals the marginal value loss evaluated at an optimal preset price. Here, u indexes alternative risk environments: in each environment, the firm chooses its preset price before the credit state is realized.

To relate this value loss to currency choice, recall that $N_j \equiv V_j^v(0) - V_j^e$ is the initial DCP advantage defined in the main text. At any probability π, the remaining DCP advantage is therefore $\Delta_j(\pi) = N_j - H_j(\pi)$. At the exact switching cutoff, the firm is indifferent between DCP and PCP, so

$$H_j(\pi_j^*) = N_j.$$

Thus, the cutoff is the probability at which the total reduction in DCP value exactly offsets its initial advantage over PCP.

The fixed-price cutoff and its bounds. Let $P_{j,0}^v$ maximize good-state DCP profits and recall

$$L_{j,0} = \mathcal{L}_j(P_{j,0}^v), \qquad N_j = V_j^v(0) - V_j^e > 0,$$

where N_j is the normal-time DCP premium. Since $J^v_{j,g}(P^v_{j,0}) = V^v_j(0)$ and the normal-time price remains feasible at every probability, optimality implies

$$\begin{aligned}\Delta_j(\pi) &= \max_P \left\{ J^v_{j,g}(P) - \pi \mathcal{L}_j(P) \right\} - V^e_j \\ &\geq J^v_{j,g}(P^v_{j,0}) - \pi \mathcal{L}_j(P^v_{j,0}) - V^e_j \\ &= N_j - \pi L_{j,0}. \end{aligned} \tag{D.16}$$

The right-hand side is the value difference when the firm retains its normal-time DCP price.

At $\pi = 1$, the endpoint restriction gives $N_j - \mathcal{L}_{j,0} \leq \Delta_j(1) < 0$. Since $N_j > 0$, it follows that $\mathcal{L}_{j,0} > N_j > 0$. Evaluating the same bound at the exact cutoff gives

$$0 < \pi^{\text{lin}}_j = \frac{N_j}{\mathcal{L}_{j,0}} \leq \pi^*_j < 1.$$

Thus the conditions for an exact interior cutoff also put the fixed-price cutoff inside the probability interval. The expansion concerns the value near zero; it does not supply a global error bound for the approximate cutoff. The fixed-price interpretation and lower bound remain exact.

D.4 Working Capital and Dollar-Input Comparative Statics

This appendix proves the two conditional results in Section 4.5. All comparisons allow input quantities and preset prices to adjust optimally.

Differentiating the exact cutoff. At a regular interior cutoff, $\Delta_j(\pi, z) = N_j(z) - H_j(\pi, z)$ is differentiable and $H_{j,\pi} = \mathcal{L}_j(P^{v,*}_j(\pi), z) > 0$. Differentiating the indifference condition gives

$$0 = \partial_z N_j - \partial_z H_j - H_{j,\pi} \frac{\partial \pi^*_j}{\partial z}.$$

Rearranging proves Equation 32. Substituting $z = \nu_s$ proves Proposition 2; substituting $z = \omega_v$ proves Proposition 3. Both signs compare changes in optimized value levels at the same probability.

Working-capital dependence and financing needs. A larger ν_s increases the share of the input bill that must be financed before revenues arrive. In the normal-financing benchmark, this raises required principal even after inputs adjust to financing costs. At fixed input prices, wage, and borrowing rate, κ^0_j is proportional to $\nu_s(1+\nu_s i_j)^{\alpha-1}$, so

$$\frac{\partial \log \kappa^0_j}{\partial \nu_s} = \frac{1 + \alpha \nu_s i_j}{\nu_s(1+\nu_s i_j)} > 0. \tag{D.17}$$

Since $\mathcal{W}^0_j = \kappa^0_j Y$, total required principal also increases at fixed output. The effect on the switching cutoff depends on how financing needs affect both the normal-time premium N_j and the risk cost H_j.

The following sufficient case establishes that comparison using the full cost and pricing problems.

A sufficient case for working-capital amplification. Set $i_e = i_v = 0$ and hold the macroeconomic distribution, input prices, wage, technology, all usable-credit limits, and Ψ_s fixed. Require ordinary-credit slackness locally at the PCP optimum and in the good state at both the normal-time and cutoff DCP prices. Finally, suppose $G^{v,*}_{j,b} > 0$ with positive probability at the cutoff allocation.

The state-cost envelope, at fixed output, is

$$\partial_{\nu_s} C_j = p_X X^* \Psi_s'(G^*),$$

where $\tilde{P}_X = p_X$ and ν_s enters cost only through G. The derivatives are zero at the specified slack allocations. They are positive in the bad state whenever $G^{v,*}_{j,b} > 0$. Applying the price envelope therefore gives $\partial_{\nu_s} N_j = 0$ and

$$\partial_{\nu_s} \Delta_j(\pi_j^*, \nu_s) = -\pi_j^* \mathbb{E}_{\mathbf{X}} \left[p_X X^{v,*}_{j,b} \Psi_s'(G^{v,*}_{j,b}) \right] < 0. \tag{D.18}$$

Thus $\partial_{\nu_s} H_j > 0$ and $\partial \pi_j^* / \partial \nu_s < 0$. The input and price envelopes include optimal adjustment. With positive normal borrowing rates, the state-cost derivative also contains $p_X i_j X^*$, and the relative effects on normal and stressed values must be compared.

Dollar-priced inputs, hedging, and financing costs. A larger ω_v links a greater share of input costs to the dollar's RMB value, creating a motive to match those costs with dollar revenues. The normal-time value comparison in Appendix D.6 derives this covariance channel. To assess the switching cutoff, its contribution to N_j must be compared with its effect on H_j. Even with $\mathcal{A}$ fixed, changing ω_v changes the input price and the weighted borrowing rate. For example,

$$\partial_{\omega_v} \log \bar{P}^e_X = \log\left(\frac{\mathcal{E}_{ev} P^v_x}{P^e_x} \right), \qquad \partial_{\omega_v} i_j = i_v - i_e.$$

The first expression varies with the exchange rate, while the second depends on currency-specific financing terms. At a fixed output, the cost envelope is

$$\partial_{\omega_v} C_j = X^* \partial_{\omega_v} \tilde{P}_X + \nu_s X^* \Psi_s'(G^*) \partial_{\omega_v} \bar{P}^e_X. \tag{D.19}$$

Thus holding the loss of credit capacity fixed does not hold its profit cost fixed. Proposition 3 requires that the increase in the normal-time level premium dominate the change in the risk cost. The next appendix isolates the normal-time hedging channel.

D.5 The Normal-Time Hedging Motive

This appendix isolates the covariance channel discussed before Proposition 3. It first derives normal-time optimized profits and then distinguishes the relative-value comparison from the level premium used in the exact cutoff.

Normal-time prices and values. Suppose ordinary credit is slack at the relevant unconstrained normal-time optima. Write $D_i = (\mathcal{E}_{ed} Q_i)^{\epsilon}$ for the demand shifter and define

$$U_c = \mathbb{E}_{\mathbf{X}}[D_i a_c^{1-\epsilon}], \qquad T_{j,c} = \mathbb{E}_{\mathbf{X}}[MC_j^0 D_i a_c^{-\epsilon}].$$

At preset price P^c, expected normal profit is $U_c (P^c)^{1-\epsilon} - T_{j,c}(P^c)^{-\epsilon}$. Differentiating and solving the first-order condition gives

$$P_{j,c}^0 = \frac{\epsilon}{\epsilon - 1} \frac{T_{j,c}}{U_c}, \qquad V_{j,c}^0 = \frac{(\epsilon-1)^{\epsilon-1}}{\epsilon^{\epsilon}} \frac{U_c^{\epsilon}}{T_{j,c}^{\epsilon-1}}. \tag{D.20}$$

Assume these prices lie inside the feasible intervals. Slackness at these optima is sufficient for the values to equal the full normal-state values: $V_{j,v}^0 = V_j^v(0)$ and $V_{j,e}^0 = V_j^e$. Emergency-finance costs cannot raise profits at any alternative price.

Incomplete insulation under PCP. If the PCP credit arrangement also deteriorates in the bad state, define its expected loss $\mathcal{L}^e(P^e)$ analogously. The relative value derivative becomes

$$\Delta_j'(\pi) = -\mathcal{L}_j^v(P_j^{v,*}(\pi)) + \mathcal{L}_j^e(P_j^{e,*}(\pi)). \tag{D.21}$$

The relevant condition is a larger optimized incremental loss under DCP. In this extension, the difference of the two optimized values need not be convex, so uniqueness requires a separate single-crossing condition. Complete PCP insulation is sufficient for the baseline result, but relative impairment is the underlying economic comparison.

D.6 The Normal-Time Hedging Motive

Suppose ordinary credit is slack at the relevant unconstrained PCP and good-state DCP optima, almost surely over the macroeconomic distribution. Write $D_i = (\mathcal{E}_{ed} Q_i)^{\epsilon}$ for the demand shifter and define

$$U_c = \mathbb{E}[D_i a_c^{1-\epsilon}], \qquad T_{j,c} = \mathbb{E}[MC_j^0 D_i a_c^{-\epsilon}], \qquad a_e = 1, \quad a_v = \mathcal{E}_{ev}.$$

Expected normal profits equal $U_c (P^c)^{1-\epsilon} - T_{j,c}(P^c)^{-\epsilon}$. For $\epsilon > 1$, the interior optimum and its value are

$$P_{j,c}^0 = \frac{\epsilon}{\epsilon - 1} \frac{T_{j,c}}{U_c}, \qquad V_{j,c}^0 = \frac{(\epsilon-1)^{\epsilon-1}}{\epsilon^{\epsilon}} \frac{U_c^{\epsilon}}{T_{j,c}^{\epsilon-1}}. \tag{D.22}$$

Slackness at these optima implies $V_{j,v}^0 = V_j^v(0)$ and $V_{j,e}^0 = V_j^e$: emergency financing adds no cost there and cannot increase profits at another price.

The two currency values coincide at a common deterministic benchmark. Consider a small-risk expansion around this benchmark, maintaining normal-credit slackness locally. Let $v = \log(\mathcal{E}_{ev}/\overline{\mathcal{E}}_{ev})$ and $d_i = \log(D_i/\overline{D}_i)$, and write $\tilde{i}_j = i_j - \bar{i}_j$. Scale these shocks by σ, holding their joint distribution after division by σ fixed. Assume positive normal financing factors and moment conditions that support the

second-order expansion below.

With fixed wages and upstream invoice prices, define

$$\lambda_j = \frac{\nu_s}{1+\nu_s \bar{i}_j}, \qquad m_j = \alpha\omega_v v + \alpha\lambda_j \tilde{i}_j.$$

Here m_j is the first-order part of the log marginal-cost deviation:

$$\begin{aligned} \log \frac{MC_j^0}{\overline{MC}_j^0} &= \alpha\omega_v v + \alpha \log(1+\lambda_j \tilde{i}_j) \\ &= m_j - \frac{\alpha}{2}\lambda_j^2 \tilde{i}_j^2 + o(\sigma^2). \end{aligned} \tag{D.23}$$

The quadratic financing term is present in log costs. Its mean enters both log currency values with the same coefficient and cancels in their difference. Its covariance with v is of order σ^3, so it does not enter the second-order value comparison. Joint Gaussianity is not required.

Applying $\log \mathbb{E}[e^z] = \mathbb{E}[z] + \frac{1}{2}\mathrm{Var}(z) + o(\sigma^2)$ to the moments in Equation D.22 gives

$$\begin{aligned} &\log V_{j,v}^0 - \log V_{j,e}^0 \\ &= \frac{\epsilon}{2}\left[\mathrm{Var}(d_i + (1-\epsilon)v) - \mathrm{Var}(d_i)\right] \\ &\quad - \frac{\epsilon-1}{2}\left[\mathrm{Var}(d_i + m_j - \epsilon v) - \mathrm{Var}(d_i + m_j)\right] + o(\sigma^2). \end{aligned} \tag{D.24}$$

The mean terms cancel. Expanding the variances also cancels the demand covariance: its coefficient is $\epsilon(1-\epsilon) + \epsilon(\epsilon-1) = 0$. The remaining terms yield

$$\log V_{j,v}^0 - \log V_{j,e}^0 = \epsilon(\epsilon-1)\left[\mathrm{Cov}(m_j, v) - \tfrac{1}{2}\mathrm{Var}(v)\right] + o(\sigma^2). \tag{D.25}$$

Since $\mathrm{Cov}(m_j, v) = \alpha\omega_v \mathrm{Var}(v) + \alpha\lambda_j \mathrm{Cov}(\tilde{i}_j, v)$, this is equivalently

$$\log V_{j,v}^0 - \log V_{j,e}^0 = \epsilon(\epsilon-1)\left[\left(\alpha\omega_v - \tfrac{1}{2}\right)\mathrm{Var}(v) + \alpha\frac{\nu_s}{1+\nu_s \bar{i}_j}\mathrm{Cov}(\tilde{i}_j, v)\right] + o(\sigma^2). \tag{D.26}$$

Inside the brackets, the exchange-rate variance term combines the input-currency hedge contribution $\alpha\omega_v \mathrm{Var}(v)$ with the remaining currency-pricing contribution $-\frac{1}{2}\mathrm{Var}(v)$. The covariance term $\alpha\lambda_j \mathrm{Cov}(\tilde{i}_j, v)$ captures how normal financing costs move with the dollar's RMB value.

To make the risk scale explicit, write $v = \sigma z_v$ and $\tilde{i}_j = \sigma z_j$, with a fixed joint distribution of (z_v, z_j). Then $\mathrm{Var}(v) = \sigma^2 \mathrm{Var}(z_v)$ and $\mathrm{Cov}(\tilde{i}_j, v) = \sigma^2 \mathrm{Cov}(z_j, z_v)$, so

$$\log V_{j,v}^0 - \log V_{j,e}^0 = \sigma^2 \Gamma_j + o(\sigma^2), \tag{D.27}$$

where $\Gamma_j \equiv \epsilon(\epsilon-1)\left[\left(\alpha\omega_v - \frac{1}{2}\right)\mathrm{Var}(z_v) + \alpha\frac{\nu_s}{1+\nu_s \bar{i}_j}\mathrm{Cov}(z_j, z_v)\right]$. Thus Γ_j can have either sign. If $\Gamma_j > 0$, then $V_{j,v}^0 > V_{j,e}^0$, and hence $N_j > 0$, for sufficiently small nonzero risks. This is a local sufficient

condition; the expansion does not determine the sign at arbitrary volatility.

Holding the financing-covariance term and exchange-rate distribution fixed, a larger ω_v raises the leading log advantage. However, $i_j = (1-\omega_v)i_e + \omega_v i_v$ generally changes with ω_v, affecting both λ_j and the financing covariance. Moreover, the cutoff concerns the level premium N_j, whereas this expansion describes the log ratio. With $R_j = V_j^v(0)/V_j^e$, the identity $N_j = (R_j - 1)V_j^e$ shows why changes in the PCP value also matter. A higher cutoff therefore requires $\partial_{\omega_v} N_j > \partial_{\omega_v} H_j$ at the cutoff, holding π fixed in the partial derivatives, as stated in Proposition 3.

D.7 Value-Weighted Aggregation and Local Empirical Predictions

Let $m_{ijs,0} \geq 0$ denote the baseline value of country i's imports from Chinese exporter j in industry s, where 0 denotes 2010. For destination–industry pairs with positive baseline trade, the value-weighted distribution is

$$F_{is}(\pi) = \frac{\sum_{j\in(i,s)} m_{ijs,0}\mathbf{1}\{\pi_j^* \leq \pi\}}{\sum_{j\in(i,s)} m_{ijs,0}}. \tag{D.28}$$

This is the share of baseline trade accounted for by exporters choosing RMB at risk level π, with RMB chosen at indifference. Exporters that choose RMB even at $\pi = 0$ can be assigned threshold zero; exporters that retain USD throughout $\pi \in [0,1]$ can be assigned threshold $+\infty$. Thus $F_{is}(\pi)$ includes all RMB users at that probability, while $F_{is}(\pi_t) - F_{is}(\pi_0)$ measures the additional share switching when $\pi_t > \pi_0$, with $\pi_t = \pi(\tau_t)$.

The local formulas apply where the threshold distribution admits a density, as in a smooth continuum approximation. At a differentiability point u, let $f_{is}(u) = F'_{is}(u)$. Then, as $\delta \to 0$,

$$F_{is}(u+\delta) - F_{is}(u) = f_{is}(u)\delta,$$

where $f_{is}(u)$ measures the concentration of baseline trade value per unit increase in the bad-state probability. A large exporter contributes more to this distribution than a small exporter. Substituting $u = \pi(\tau_t)$ and using the chain rule gives

$$\frac{\partial F_{is}(\pi(\tau_t))}{\partial \tau_t} = f_{is}(\pi(\tau_t))\pi'(\tau_t).$$

Multiplying by $E_{i,2010}$ and $\omega_{is,2010}$, and summing across industries yields Equation 35.

If a sector has lower thresholds in the sense that $F_H(\pi) \geq F_L(\pi)$ at all feasible probabilities, it has a weakly higher RMB adoption level at each probability under otherwise identical weights. This does not order $F_H(\pi_1) - F_H(\pi_0)$ and $F_L(\pi_1) - F_L(\pi_0)$. For example, thresholds uniformly distributed on $[0, 0.4]$ are lower than thresholds on $[0.4, 0.8]$. Between probabilities 0.5 and 0.6, the first group has already fully switched and makes no additional adjustment; the second group's RMB share rises by 0.25. Earlier adoption and a larger current response are therefore distinct predictions.