Trade-Based Quasi-Fiscal Devaluations*

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Abstract

We study trade-based (quasi-)fiscal devaluations by exploiting changes in China's system of partial value-added tax rebates on exports (VATRX), which act as time-varying export subsidies. Using highly disaggregated VATRX and customs data, we estimate the dynamic effects of rebate changes on export quantities and prices. Permanent changes in VATRX generate a dynamic quantity response—with trade elasticities rising from 1.2 on impact to 17 in the long run—and complete pass-through to export prices. We embed these micro estimates into a two-country New Keynesian model with delayed substitution in trade flows. Applying the model to the 2008 global financial crisis, we find that China's 1.6 percentage point increase in VATRX boosted exports and GDP by 6.33 and 0.48 percentage points, respectively. Trade policy thus offers an effective fiscal-devaluation instrument.

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1 Introduction

Recent years have seen a resurgence of active trade policy around the world and a renewed focus on its interaction with macroeconomic stabilization tools, especially in the context of rising geopolitical tensions. This environment has stimulated a theoretical and empirical literature examining how trade instruments interact with fiscal and monetary policy.[1] Within this broader discussion lies the established literature on fiscal devaluations, which studies policy combinations designed to replicate the allocations of a nominal devaluation when monetary policy is restricted, as in currency unions or under managed exchange rates (De Mooij and Keen, 2013; Farhi et al., 2014; Engler et al., 2017; Giagheddu, 2020; Arachi and Assisi, 2021; Erceg et al., 2023). This literature highlighted two broad classes of policies capable of generating a fiscal devaluation: a fiscal type, involving an increase in the value-added tax (VAT) paired with a uniform reduction in payroll taxes, and a trade type, implemented through higher export subsidies and import tariffs. While the fiscal variant has been examined in episodes such as Germany in 2007 and Spain in 2008, the trade-based variant has received little empirical scrutiny, despite its conceptual symmetry and potential relevance in the current policy landscape.

This paper examines China's system of partial VAT rebates on exports (VATRX) as a real-world instance of the trade-based channel of a *quasi-fiscal devaluation*. Because VAT rebates reduce the effective tax burden on exported goods, changes in rebate rates are economically equivalent to changes in export subsidies. We refer to this policy, in the context of China's managed exchange rate system, as a *quasi-fiscal devaluation* since it uses one of the two elements of the trade-based variant.[2] China's rebate system offers an unusually rich empirical setting: rebate adjustments are frequent, occur at a highly disaggregated product level, and are staggered across industries. We exploit this variation to estimate the dynamic responses of export prices and quantities to rebate changes, and then incorporate those partial-equilibrium estimates into a two-country New Keynesian model. We use the estimated model to quantify the aggregate implications of China's VATRX policy as a quasi-fiscal devaluation tool, focusing on the 2008 global

[1] Itskhoki and Mukhin (2023, 2025); Cuba-Borda et al. (2024); Bianchi and Coulibaly (2025); Monacelli (2025); Guerrieri et al. (2025); Alessandria et al. (2025b); Schmitt-Grohé and Uribe (2025); Ambrosino et al. (2025).

[2] Figure B1 of the appendix shows that average tariffs were relative flat between 2008 and 2010.

financial crisis when export rebates at the aggregate level rose by 1.6 percentage points.

To understand how these policy changes operate in practice, it is useful to describe the structure of China's VAT system. The payable VAT is composed of the difference between the VAT on domestic output and the input VAT, plus any export-related VAT. For exported goods, the effective VAT is reduced by the VATRX, so that the export VAT is calculated as the product of export value times the difference between the statutory VAT rate and the rebate.[3] In practice, this means that higher VATRX rates reduce the effective tax burden on exporters, while lower rates increase it. While most countries tend to rebate back a uniform 100% rate on the VAT for each good, so that the effective VAT on exports is zero, China follows a partial rebate system with significant variation across goods and time. As a result, changes in China's VATRX function economically like adjustments to an export subsidy or tax.

Figure 1, Panel (a), plots the evolution of China's aggregate domestic VAT rate and effective export VAT rate from 2004 to 2019.[4] While the domestic VAT (red dashed line) remained relatively stable, the effective export VAT (blue–o line) exhibited substantial fluctuations driven by changes in the VATRX. These fluctuations reflect a sequence of policy episodes and broader tax reforms.[5] A particularly salient episode is the 1.6 percentage point reduction in the effective export VAT during the 2008 global financial crisis. Panel (b) shows that this increase in VATRX occurred during a period in which China effectively pegged the Yuan to the U.S. dollar, creating conditions closely resembling a textbook fiscal devaluation. Panel (c) further illustrates that goods exposed to VATRX adjustments during the crisis (blue–o line) experienced significantly stronger export growth than goods unaffected by the policy (orange dotted line). Taken together, these patterns suggest that China's use of VATRX around the 2008 recession effectively implemented a fiscal devaluation that was strongly associated with improved export performance.

Although these aggregate patterns suggest that VATRX policy could play a countercyclical role, they do not provide a causal estimate of its effects: the government may adjust rebate rates for goods with unobserved characteristics that also affect export performance. Micro-level evi-

[3]Mathematically, we have that: VAT payable $=$ Dom output VAT$-$Input VAT$+$Export VAT, where Export VAT $=$ exports $\times$ (VAT rate $-$ VATRX), so that the Effective VAT on Exports $=$ (VAT rate $-$ VATRX).

[4]Section 2 provides details on the data construction.

[5]For example, the rise in the effective export VAT (i.e., decline in VATRX) around 2013 stemmed from fiscal reforms replacing business tax revenue with VAT.

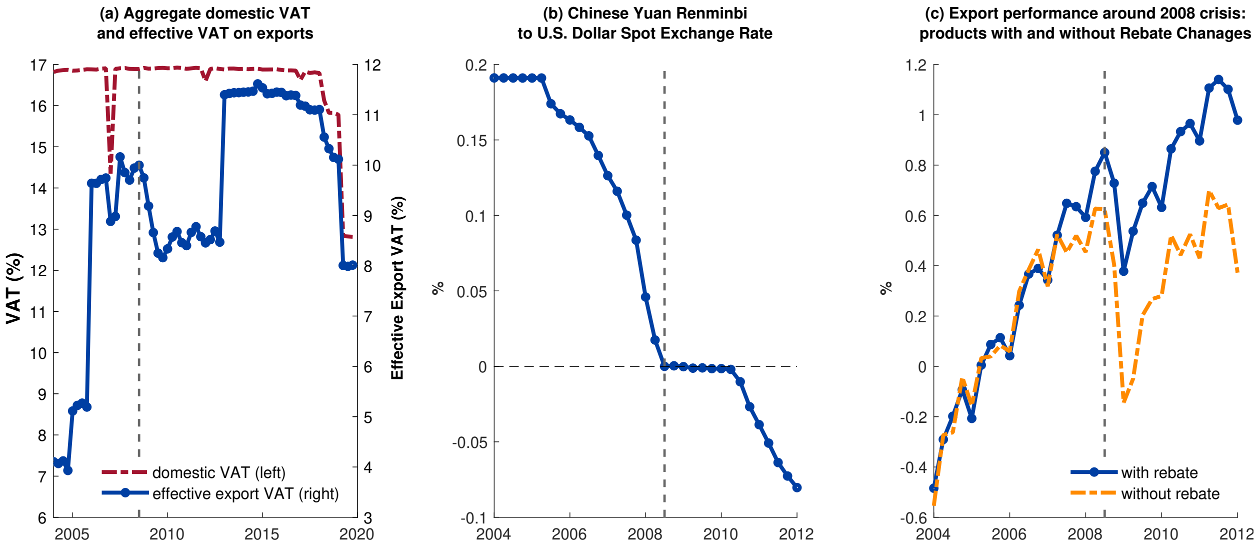


Figure 1: China's VAT System and Export Performance

Note: Panel(a) shows in red-dotted line the domestic VAT in the right y-axis, and in blue-*o* line the effective export VAT in the left y-axis. Panel (b) shows in blue-*o* line the Chinese Yuan Renminbi to U.S. dollars, normalized to be zero in 2008Q3. Panel (c) shows the cumulative aggregate export growth of goods that received an increase in VATRX in the two years window around 2008Q2 (blue-*o* line), and for goods that did not face any change in the VATRX around this window (orange dashed-line).

dence is therefore crucial to identify the transmission of VATRX changes. Moreover, assessing the aggregate impact requires a structural framework, as in the fiscal-devaluation analyses for Germany and Spain (Farhi et al., 2014; Erceg et al., 2023), but one that is explicitly consistent with micro-level dynamics. In particular, VAT-RX adjustments affect both export quantities and export prices, so credible aggregate estimates must be grounded in empirical measures that capture the full dynamic response along both margins. This is where the granular structure of China's VATRX system becomes especially valuable.

We exploit the fact that rebate adjustments are frequent, highly disaggregated, and staggered across industries. Using the difference-in-differences local projection approach developed by Dube et al. (2025), we estimate the dynamic effects of VATRX changes on export quantities and prices. To address potential selection concerns, we propose an identifying restriction based on the presence of a zero lower bound on the VATRX. Our strategy compares goods that experience a sizable decline in VATRX to goods whose rebate rates remain at zero and thus cannot be reduced further. The key idea is that, at a given point in time, control-group goods fail to receive a rebate cut not because of government decisions but because the policy instrument has reached its mechanical lower bound. We further show that treated and control goods exhibit similar export quantity and price dynamics outside treatment windows, providing reassurance that unobserved heterogeneity does not drive our results.

Using this empirical strategy, we find a decline in the VATRX of around 6 percentage points, which is effectively permanent. The policy change's permanence provides an ideal setting for identifying the dynamic elasticities of export quantities and prices to rebate changes. First, we find that the response of export quantities is highly dynamic, with a trade elasticity of 1.2 after one quarter and 17 after six years—magnitudes consistent with recent estimates using tariff variation (Alessandria et al., 2025c). Second, we find evidence of complete pass-through of the VATRX to export prices, both on impact and over time. This result is crucial since the transmission of fiscal devaluation policies relies critically on the degree of pass-through of the VAT (Erceg et al., 2023).

To evaluate the aggregate effects of VATRX changes as a quasi-fiscal devaluation tool, we build a two-country New Keynesian model that is consistent with the partial-equilibrium responses estimated above. We incorporate trade frictions following the delayed-substitution block in Auclert et al. (2024). We show that China's VATRX policy operates as a quasi-fiscal devaluation, closely mirroring the transmission channels identified in the fiscal-devaluation literature. We compare the effects of a VATRX shock with those of standard fiscal-devaluation schemes featuring coordinated export subsidies and import tariffs under a pegged and fixed exchange rate regime. We find that VATRX reproduces many of the same expansionary mechanisms—raising exports, GDP, and consumption—even though it operates solely through export-rebate adjustments. Its effects are slightly weaker than a full fiscal devaluation, reflecting the absence of import-tariff adjustments, but remain sizable. Finally, VATRX policy is weaker under a managed than a fixed exchange rate regime, but still remains an effective quasi-fiscal devaluation tool, especially for net exports and GDP.

We further show that the strength and dynamics of this quasi-fiscal devaluation hinge critically on the trade elasticity. Whereas the standard practice in the literature is to fix the elasticity at its short-run value, doing so distorts the predicted path of GDP—overstating the short-run response and understating the medium-run adjustment. We find that the effectiveness of quasi-fiscal devaluations increases with the average trade elasticity, and that accounting for the dynamic evolution of the elasticity is essential for capturing the full transmission profile of trade-policy-induced fiscal interventions.

To assess the quantitative importance of VATRX as a quasi-fiscal devaluation tool, we estimate the model for China and the rest of the world using Bayesian methods. This approach allows the model to match the observed time series for key variables—including China's VATRX, aggregate exports, GDP, and the terms of trade—over the full sample. We then use the estimated model to quantify the aggregate effects of VATRX changes during the 2008 global financial crisis. Specifically, we construct a counterfactual path in which China keeps the VATRX at its pre-crisis level (the second quarter of 2008), implying a 1.6 percentage point higher rebate rate throughout the crisis period.

The results indicate that the actual 1.6 percentage point reduction in the VATRX during the crisis increased China's exports by 6.33 percentage points (semi-elasticity $\approx$ 4.00), worsened the terms of trade by 0.67 percentage points (semi-elasticity $\approx$ 0.42), and raised GDP by 0.48 percentage points (semi-elasticity $\approx$ 0.30).[6] These results fall between the estimates reported in the existing literature.[7] These findings underscore the sizable macroeconomic effects of trade-based quasi-fiscal devaluations and highlight VATRX adjustments as an important channel through which China stabilized external demand during the crisis.

To summarize, we examine China's system of partial VATRX as a quasi-fiscal devaluation tool. We provide micro-level evidence on how changes in VATRX pass through to export quantities and prices across time, and develop a two-country New Keynesian model consistent with these partial-equilibrium responses to quantify the aggregate effects of VATRX adjustments during the 2008 recession. Overall, our results indicate that trade policy can serve as a effective fiscal-devaluation instrument when monetary policy is constrained.

Literature. We contribute to the literature on the interaction between trade policy and fiscal and monetary policy,[8] and, more specifically, to the fiscal-devaluation literature (De Mooij and

[6] Chinese aggregate exports and GDP fell significantly during the 2008 crisis, so our estimates indicate that in the absence of VATRX changes, exports and GDP would have fallen by an additional 6.33 and 0.48 percentage points, respectively.

[7] While Farhi et al. (2014) report strong effects of fiscal devaluations, Erceg et al. (2023) find weaker ones.

[8] Lerner (1936); Keynes (1937); Boadway et al. (1973); Mendoza and Tesar (1998); Galí and Monacelli (2005); Adao et al. (2009); Barbiero et al. (2018); Itskhoki and Mukhin (2023, 2025); Cuba-Borda et al. (2024); Bianchi and Coulibaly

Keen, 2013; Farhi et al., 2014; Engler et al., 2017; Giagheddu, 2020; Arachi and Assisi, 2021; Erceg et al., 2023). Farhi et al. (2014) argues that fiscal devaluations can be powerful stabilization tools, drawing on Spain in 2008, while Erceg et al. (2023), focusing on Germany in 2007, casts doubt on this result. Both studies theoretically examine the trade-based variant of fiscal devaluations but do not provide direct empirical estimates. We extend this literature by analyzing the trade-based variant in China. A key factor explaining the different results in Farhi et al. (2014) and Erceg et al. (2023) is the combination of VAT pass-through to prices and the persistence of the fiscal adjustment. Erceg et al. (2023) provides indirect evidence of nearly full pass-through and short-lived policy, which limits its impact. We provide direct empirical evidence of complete pass-through of VATRX to prices at the micro level, consistent with Erceg et al. (2023), but document highly persistent changes in the micro and aggregate VATRX, akin to Farhi et al. (2014). In addition, we directly estimate the dynamic response of export quantities to VATRX changes, a key channel for policy transmission. Our aggregate results lie between those of Farhi et al. (2014) and Erceg et al. (2023), indicating that trade policy can function as an fiscal-devaluation instrument when monetary policy is constrained.

We also contribute to the literature studying the impact of export subsidies in general (Desai and Hines Jr, 2008; De Meza, 1986; Collie, 1991; Itoh and Kiyono, 1987; Rotunno and Ruta, 2024) and studying the VATRX in particular. We contribute to this literature in several dimensions. First, we provide a clear estimate of the elasticity of export quantities to export subsidies. We are also the first to study the macroeconomic consequences of export subsidies within a DSGE model, with a dynamic trade elasticity. Most papers have focused on estimating the cross-sectional impact of these subsidies in the context of VATRX on sectoral exports (Gourdon et al., 2022; Liu and Lu, 2015; Chandra and Long, 2013; Braakmann et al., 2020) or have been studied within the context of static trade models (Bond et al., 2023).

Finally, we contribute to the literature estimating the trade elasticity. First, we estimate this elasticity using export subsidies, rather than tariffs which is the usual practice. The use of tariffs has been found to suffer from several caveats, including incomplete reporting and selection bias Teti (2024) and biases due to tariff announcements and anticipation Khan and Khederlarian

(2025); Monacelli (2025); Guerrieri et al. (2025); Alessandria et al. (2025b); Schmitt-Grohé and Uribe (2025); Mac Mullen (2025).

(2021). Our identification has the advantage of being based on a clean trade policy change, applying a methodology that has not been previously used in the literature for this task, the Local Projection difference-in-difference estimator. Estimates of the dynamic trade elasticity in the literature ranges from 0.7 to 3 for the short run, and from 2 to around 14 for the long run (Teti, 2024; Boehm et al., 2023; Alessandria et al., 2025a).[9] Our findings lie within these estimates, being 1.2 in the short run and slightly higher, around 17, in the long-run.

The rest of the paper is organized as follows. Section 2 describes the data, while Section 3 presents the empirical strategy and the estimated dynamic effects of VATRX on export quantities and prices. Section 4 presents the full DSGE model. Section 5 analyzes the transmission of VATRX as a quasi-fiscal devaluation tool. Section 6 discusses the estimation of the full DSGE model. Section 7 quantifies the role of VATRX as a quasi-fiscal devaluation instrument during the 2008 global financial crisis. Finally, Section 8 concludes.

2 Data

Our analysis combines three main sources of data. First, we use highly disaggregated administrative information on China's VAT and export rebate system, recorded at the HS 10-digit level, which provides precise policy changes and implementation dates for all adjustments to VATRX. This data extends from 2004 to 2024, but we restrict our analyses to 2019 to avoid the COVID period. Second, we use quarterly Chinese customs data (HS 6-digit) on export values and quantities, from which we construct unit values and apply consistent-unit filters for clean measurement of prices and quantities.[10] These two micro datasets form the basis of the empirical estimates. Third, for the structural model and Bayesian estimation, we assemble aggregate quarterly macroeconomic time series—including real GDP, consumption, trade flows, the real effective exchange rate, and the terms of trade—from the IMF and the World Bank. All details are presented in Appendix A. Together, these datasets allow us to track policy changes from their product-level implementation through to their aggregate macroeconomic effects.

[9] See Simonovska and Waugh (2014) for a survey on estimates of the trade elasticity in static models.

[10] To ensure our results are not driven by outliers, we implement a 1% winsorization, dropping any product where the log quarterly change in unit values or quantities falls outside the 0.5^{th} and 99.5^{th} percentiles of the change distribution.

3 Dynamic Quantity and Price Responses to Export Subsidies

We now estimate the partial equilibrium effects of China's VAT export rebate changes, which provide the key empirical moments to discipline our DSGE model. We begin by describing the institutional background of China's VAT rebate system, present our identification strategy that exploits institutional constraints, and conclude with empirical results demonstrating strong dynamic responses in quantities but immediate price pass-through.

3.1 Institutional Background

China's VAT system, implemented in 1994, features incomplete refunds on exports that vary significantly across products and time. When the rebate rate falls below the applicable VAT rate, exporters face an effective export tax (Feldstein and Krugman, 1990). As we show in Figure 2, these rebate adjustments are frequent, highly disaggregated, and staggered across products—features that are essential for our identification strategy.

China's system of partial VAT rebates on exports stands out internationally, where most VAT regimes apply a uniform 100% rebate. One of the central reason for China's departure from this standard practice is fiscal. VAT revenues account for a large share of government resources—approximately 38% of total revenue in 2023 (China Briefing, 2024). Although the precise split between domestic and export VAT receipts is not publicly disclosed, the sheer scale of China's export sector implies that moving to a full-rebate system would likely entail a substantial loss of fiscal revenue.[11] Maintaining partial rebates therefore preserves, on average, a significant and stable revenue base, while variations around the average rebate reflect other economic reasons.

Two institutional features of China's VAT rebate system are central for our identification strategy. Specifically rebate rates are subject to strict constraints: they cannot fall below zero and cannot exceed the statutory VAT rate. These constraints limit policymakers' ability to tailor rebates to product-specific export conditions. Second, rebate adjustments take the form of sharp,

[11] Yu and Luo (2018) They estimate that the share of domestic value added in Chinese exports in 2010 was about 77.4%, which, given the exports-to-GDP ratio of 26.7%, implies, using a simple back-of-the-envelope calculation, that a full export value-added rebate could represent almost 8% of total revenues in 2010.

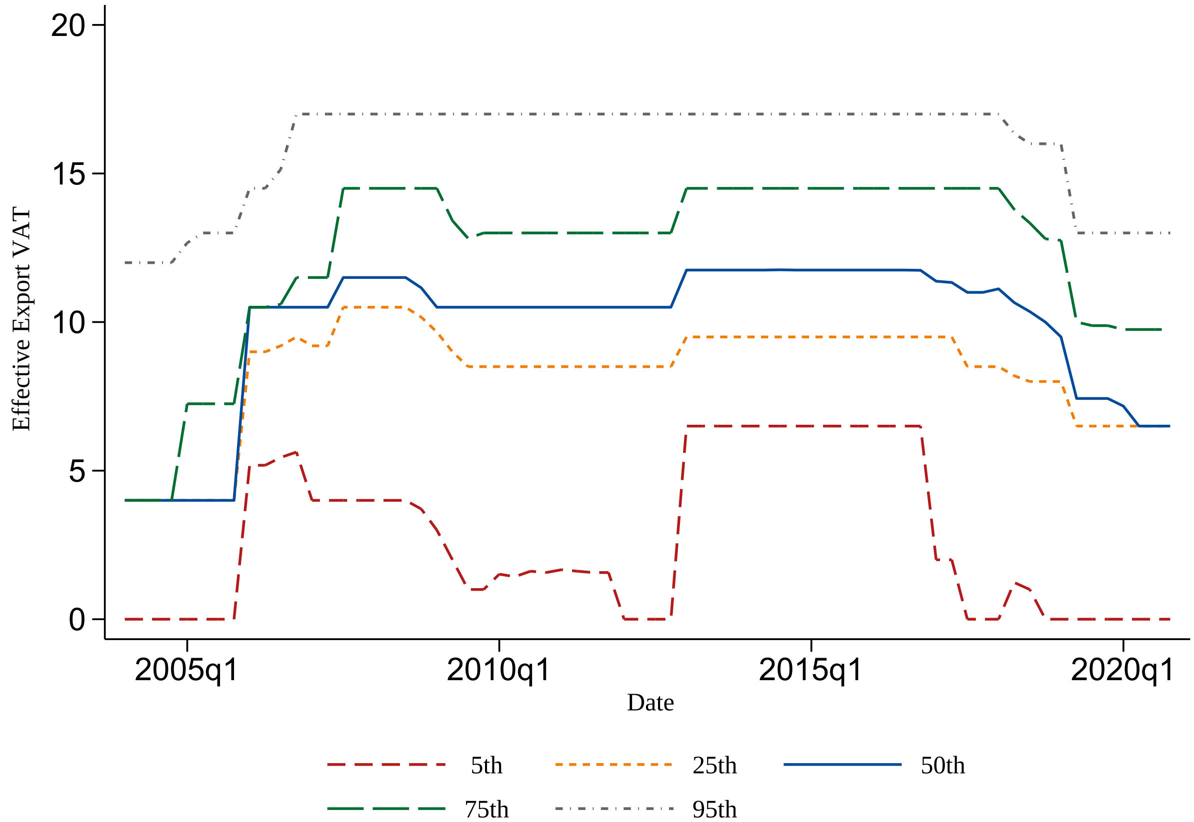


Figure 2: Effective Export VAT Across Products

Note: Effective export VAT equals the domestic VAT rate minus the VATRX. Percentiles are calculated for each time period across all products at the HS6-digit level.

discrete jumps, with different products experiencing these changes at different times. We exploit these institutional constraints, especially episodes in which the bounds become binding, to mitigate endogeneity concerns. As we explain in the next section, products at the zero lower bound provide a natural control group since authorities cannot reduce their rebates further, regardless of their export performance.

3.2 Empirical Strategy

Estimation Approach. We estimate the dynamic effects of VAT rebate changes on export quantities and prices using the local projection difference-in-differences (LP-DiD) method developed by Dube et al. (2025). This approach is particularly suited to our setting for three reasons.

First, as Figure 2 shows, rebate changes occur as sharp, discrete jumps rather than continuous variations, precluding standard panel methods. Second, these changes affect different products at different times, creating the staggered treatment timing that LP-DiD handles well. Third, we are interested in tracing out the full dynamic path of adjustment, not just the long or short-run effect.

Our main specification estimates the cumulative response at different horizons:

$$\Delta_h y_{i,t-1} = \gamma_t + \beta_h \Delta E_{i,t} + \beta \mathbb{X}_{i,t} + \epsilon_{i,t} \quad \text{for h=-8,...,-2,0,...,24} \tag{1}$$

where $\Delta_h y_{i,t-h} = y_{i,t+h} - y_{i,t-1}$ captures the cumulative change in our outcome variable (VATRX, log quantities or prices) h periods after treatment, and $\Delta E_{i,t}$ indicates whether product i enters treatment at time t. We include year fixed effects, γ_t, and control for lagged rebate levels and VAT rates in $\mathbb{X}_{i,t}$. The coefficient of our interest, β_h, identifies the average treatment cumulative effect at horizon h relative to the control group. Given the use of quarterly data, we mitigate potential seasonal bias in the dependent variables (prices or quantities) by incorporating quarter-product level fixed effects.

Identification Strategy. A key identification challenge is that VATRX changes reflect government decisions that likely respond to industry performance. Products receiving rebate reductions might differ systematically from those maintaining constant rebates, biasing our estimates rebates changes are due to specific past performance. We address this endogeneity by exploiting the institutional constraints on the rebate policy.

Our key insight is that the zero lower bound creates a set of products for which further rebate reductions are impossible, regardless of their performance. When the government reduces rebates for over-performing industries, it cannot reduce rebates (increasing export tax) for products already at zero—even if these products are performing equally well. By comparing products receiving large rebate cuts to those constrained at zero, we isolate the effect of the policy change from the selection into treatment, in the control group. Our identification strategy relies on a change in the VATRX that is opposite to the one observed during the global financial crisis (when an increased VATRX effectively lowered the export VAT).[12]

Specifically, we define our treatment group as products experiencing rebate reductions in the quarter below the 15th percentile of all quarterly reductions:

$$T_{i,t} = \begin{cases} 1 & \text{if } \Delta\text{rebate}_{i,t} < \Delta\text{rebate}^{q15} \text{ ; or } T_{i,t-1} = 1 \\ 0 & \text{if } T_{i,t+j} = 1 \text{ for any } j > 0 \end{cases} \tag{2}$$

Our control group consists of products perpetually constrained at zero in our whole sample:

$$C_{i,t} = \begin{cases} 1 & \text{if rebate}_{i,t+h} = 0 \ \forall \ h \in \mathbb{N} \\ 0 & \text{otherwise} \end{cases} \tag{3}$$

We also impose that the set of clean control units is the same across all post-and pre-treatment time horizons, to rule out composition effects across the post-treatment window (Dube et al., 2025).

We then restrict our estimation sample to newly treated products and the products within the clean control group:

$$E_{i,t} = \begin{cases} 1 & \text{if } \Delta T_{i,t} = 1 \text{ (newly treated products)} \\ 0 & \text{if } C_{i,t} = 1 \text{ (clean control products)} \end{cases} \tag{4}$$

This design ensures clean identification: treated units enter treatment once and remain treated (an absorbing state), while control units are those who never receive treatment because the institutional features does not allow it. We focus on large rebate reductions for two reasons. First, this rules out small changes potentially driven by measurement error or compositional effects. Second, since we aim to understand macro-level impacts, large changes are more relevant, particularly if responses exhibit non-linearities.[13]

[12] Consistent with our quantitative model, we assume the effects of VATRX changes on exports are symmetric when incorporating our empirical estimates.

[13] As we will explain later, this strategy provides conservative estimates for the trade response to these policy

While the zero lower bound eliminates government discretion over rebate reductions for our control group, three identifications concerns remain. First, if rebates tend to be counter-cyclical to product performance, treated products might receive rebate cuts (export tax increases) precisely because they are outperforming others. Since rebate reductions should harm exports, this selection would bias our estimates toward zero—making them conservative lower bounds. Second, control group products reached zero rebates through past government decisions that might reflect long-term performance trends. If these products were historically strong exporters, they could continue outperforming treated products, again biasing our estimates upward (in absolute terms). We mitigate this by requiring control products to have maintained zero rebates throughout our entire sample and by conducting pre-treatment balance tests. Third, treatment and control groups might differ in fundamental characteristics like trade elasticities—potentially explaining why certain products were targeted for rebate changes. While this does not invalidate our estimates for treated products, its extrapolation to the broader economy, would generate biases on the aggregate results. We examine all three concerns through the validation tests presented next.

Validation Tests. Before presenting our main results, we verify that treatment and control groups exhibit similar characteristics outside of treatment periods. We focus on periods without rebate changes within a one-year window (past or future) to ensure clean comparisons. The objective is to test for the concerns mentioned above. If products across groups tend to be systematically different, or those in the control group with zero rebates have outperformed other products in terms of export performance, we would expect differential behavior in their price or quantities changes. We do not find this to be the case.

We start by testing for differences in unconditional moments using cross-sectional regressions:

$$Y_i = \beta I_i^{treated} + \epsilon_i$$

where Y_i represents the mean or standard deviation of yearly changes in exports and prices, and $I_i^{treated}$ equals one for ever-treated products and zero for never-treated products. Panel A of Table 1 shows no statistically significant differences between groups in either the average or volatility

relative to other approaches.

of price and quantity changes.

We also test for differential conditional responses by estimating:

$$Y_{i,t} = \beta_1 I_i^{treated} \times \mathbb{X}_{i,t} + \beta_2 \mathbb{X}_{i,t} + \gamma_t + \theta_i + \epsilon_{i,t}$$

where $Y_{i,t}$ represents logged unit values and quantities, $\mathbb{X}_{i,t}$ includes the VAT level (not rebate) and lagged outcomes, and γ_t and θ_i denote time and product fixed effects. Panel B of Table 1 confirms that responses to these variables do not differ significantly between groups, supporting our identification strategy.

Table 1: Pre-treatment analysis

Panel A: Unconditional moments

	(1) $\Delta^y \ln(q_{it})$ std dev	(2) $\Delta^y \ln(q_{it})$ av.	(3) $\Delta^q \ln(p_{it})$ std dev	(4) $\Delta^q \ln(p_{it})$ av.
$I_i^{\text{treatment}}$	-0.06	0.11	0.18	-0.05
	(0.26)	(0.17)	(0.26)	(0.15)
Observations	3495	3517	3495	3517

Panel B: Conditional moments

	(1) $\ln(q_{it})$	(2) $\ln(p_{it})$
$I_i^{\text{treatment}} = 1 \times VAT_{it-1}$	0.00	-0.00
	(0.00)	(0.00)
$I_i^{\text{treatment}} = 1 \times \ln(p_{it-1})$	0.01	
	(0.03)	
$I_i^{\text{treatment}} = 1 \times \ln(q_{it-1})$		0.00
		(0.02)
Observations	48,011	48,011

Notes: Panel 1 reports unconditional moments and Panel 2 reports differential estimates between treated and control groups. The sample is restricted to periods without treatment in the prior or subsequent year. Standard errors in parentheses. *** p < 0.01, ** p < 0.05, * p < 0.1.

3.3 Empirical Results

Rebate Changes. We begin by verifying that our treatment indeed generates substantial variation in export rebates. Figure 3 presents the estimated coefficients from equation (1) using rebates as the dependent variable. The results shows an immediate and permanent reduction in VAT exports rebates of around 6 percentage points for treated products relative to controls. While we expected a reduction by construction, the sharp and permanent nature of this change —with no gradual phase-in or subsequent reversals— was not imposed. This clean policy variation facilitates interpretation of the subsequent quantity and price responses.

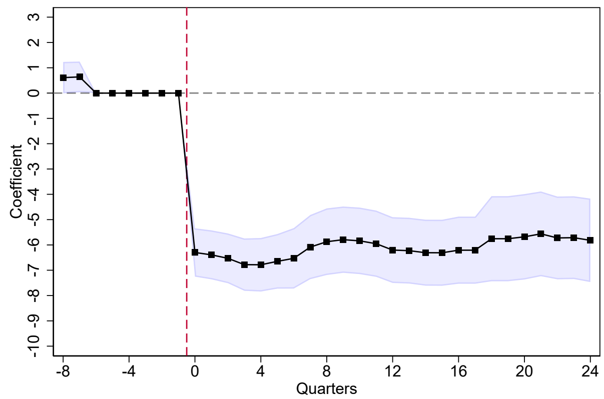


Figure 3: VAT rebate dynamics

Note: Time is measured in quarters relative to the rebate change ($t = 0$). The base period is $t = -1$. Coefficients show the percentage point change in VATRX for treated products relative to non-treated products. The shaded area indicates the 90% confidence interval, clustered by product code.

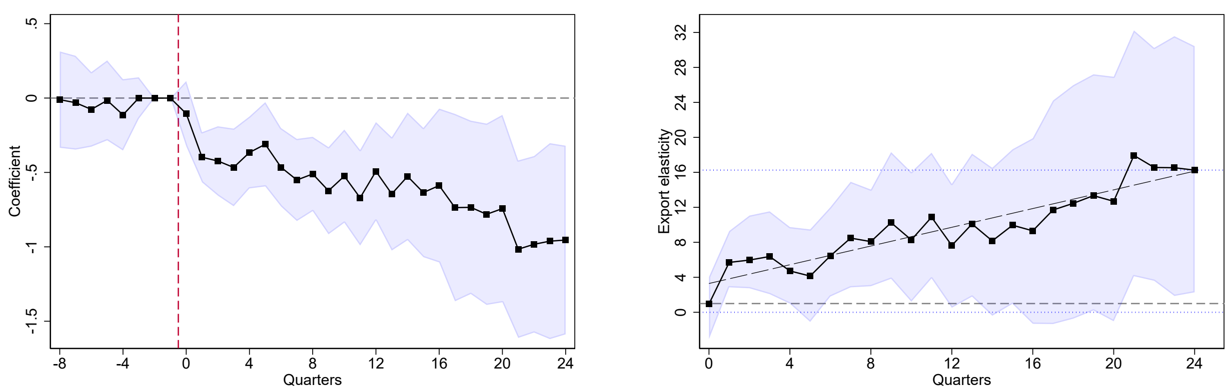
Export Quantity Dynamics. Panel (a) of Figure 4 presents our main result: the dynamic response of export quantities to VATRX reductions. Two features stand out. First, pre-trends are statistically indistinguishable from zero, and if anything suggest treated products slightly outperformed controls before treatment—which would bias our estimates conservatively downward. Second, despite the immediate and permanent rebate change, export quantities adjust gradually. Exports fall by only 6% in the first quarter but decline by approximately 95 log points after 24 quarters.

This gradual adjustment provides clear evidence for dynamic trade responses to permanent policy changes. While a substantial literature has theorized such dynamics (Alessandria et al., 2021a; Fitzgerald et al., 2024; Boehm et al., 2024; Alessandria et al., 2021b, 2025c), clean identification has proven elusive due to policy instability (Alessandria et al., 2025c), phase-out nature of trade policy (Besedes et al., 2020), and data limitations (Teti, 2024). Our setting overcomes these challenges by exploiting sharp, and permanent policy changes.

To quantify the economic magnitude, we compute implied trade elasticities at each horizon by comparing the quantity response to the rebate change:[14]

$$\textit{export-rebate elasticity}_h = \frac{\hat{\beta}_h^{exports} \times 100}{\hat{\beta}^{rebate}}$$

Panel (b) of Figure 4 shows these elasticities rising from around 1.0 in the short run to between 16 after 24 quarters. These estimates exceed unity at all horizons, indicating that rebate increases reduce trade flows both immediately and in the long run. Our short-run elasticity aligns closely with recent estimates by Alessandria et al. (2025a,c), while contrasting with smaller elasticities found in some studies (Boehm et al., 2023; Teti, 2024).[15]



(a) Export response

(b) Trade elasticity

Figure 4: Export Dynamics and Trade Elasticity

Note: Time is measured in quarters relative to the rebate change ($t = 0$). The base period is $t = -1$. Quarter-2 digit products fixed effects are included to control for seasonality. Panel (a) coefficients show the change in export quantity (treated relative to non-treated products), with the 90% confidence interval clustered by product code. Panel (b) plots the estimated export elasticity. The shaded area indicates the 90% pivot confidence interval, derived from 1,250 bootstrap samples.

[14]We multiply by 100, as our VAT data is in percentage points changes while exports response is in log changes.

[15]Since determining the distribution of the export elasticity in closed form is infeasible— as it depends on the ratio of two random variables— we developed a bootstrap algorithm to compute the confidence interval, detailed in Appendix B.2.

Price Pass-Through. Identifying the drivers of gradual quantity adjustment is central to our quantitative analysis. The underlying mechanism is likely to significantly affect the policy impact of fiscal devaluations (Erceg et al., 2023). The literature proposed two channels. The first one emphasizes real frictions—sunk costs, customer accumulation, learning-by-doing, or capacity constraints— where slow export growth does not depend on the price stickiness (Roberts and Tybout, 1997; Alessandria and Choi, 2007; Kohn et al., 2016; Ruhl and Willis, 2017; Steinberg, 2023; Fitzgerald et al., 2024; Merga, 2025), though it can rely on its inherent price dynamics. The second focuses on nominal rigidities—sticky prices or invoicing frictions—where delayed price adjustment governs quantity dynamics (Gopinath and Itskhoki, 2011, 2022; Amiti et al., 2019).

Figure 5 plots the export price dynamics. We cannot reject full pass-through at any horizon; prices seem to adjust within two quarters in this set up. The black line displays the estimated response with 95% confidence intervals, while the dotted line indicates complete pass-through given the estimated rebate change. This rapid price adjustment supports theories where the export dynamics operate through changes in the intercept of the demand directly, such as customer capital accumulation, (Fitzgerald et al., 2024; Steinberg, 2023; Merga, 2025), rather than theories relying on specific price dynamics.

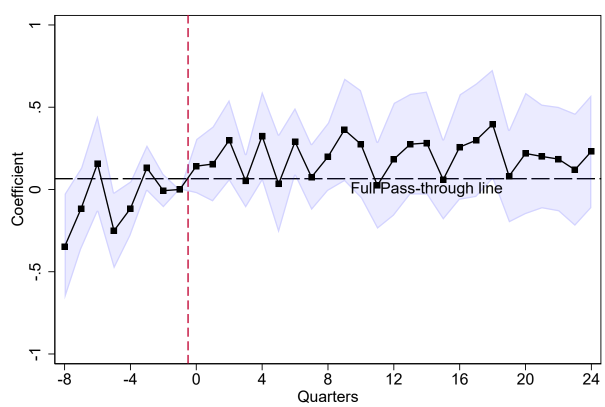


Figure 5: Price Dynamics

Note: Time is measured in quarters relative to the rebate change ($t = 0$). The base period is $t = -1$. Coefficients show the change in export quantity (treated relative to non-treated products). The 90% confidence interval is clustered by product code. The horizontal dashed black line shows the theoretical full pass-through case, assuming 100% of the firm's export cost is affected by the VATRX. Quarter-2 digit products fixed effects are included to control for seasonality.

3.3.1 Robustness

We conduct extensive robustness checks, which we detail in Appendix B.4 to also test some other potential concerns. Our results are remarkably stable across several specifications: (i) restricting treatment products without further change in the rebate after treatment (which addresses concerns about anticipated future changes and post-treatment dynamics); (ii) relaxing our treatment definition to include smaller changes in the rebate; and (iii) constraining the sample to products whose quantities are reported in the main unit (Kilograms). Critically, all specifications yield similar dynamic patterns and elasticity estimates.

4 DSGE Model

We develop a two-country New Keynesian model that incorporates the partial equilibirum dynamic trade elasticity estimated in Section 3, nominal rigidities in wages and prices, and a rich set of government policies, shocks, and frictions. The two countries are China and the ROW, which differ in both size and policy frameworks.

In China, the fiscal authority levies a VAT and provides VATRX rebates on exports. The monetary authority follows a money growth rule that targets inflation while also responding to movements in the nominal exchange rate to stabilize its value. In contrast, the ROW government is represented solely by a monetary authority, which follows a standard Taylor rule aimed at stabilizing inflation.

In each country, intermediate firms produce differentiated varieties using labor as the sole input and are subject to aggregate total factor productivity (TFP) shocks. Firms sell their output to both domestic and foreign wholesalers and face stochastic iceberg trade costs when exporting. Pricing-to-market frictions are introduced by allowing the elasticity of substitution across imported varieties to evolve endogenously with the state of the economy, generating incomplete pass-through of exchange rate changes to prices and deviations from the law of one price.

Wholesalers aggregate domestic and imported intermediate varieties using a CES technology. Sticky quantities are introduced following the delayed substitution mechanism of Auclert et al. (2024).

A continuum of retailers produces differentiated final consumption goods using the composite intermediate bundle supplied by wholesalers. Retailers face Calvo-style price rigidity, so that only a fraction of firms can optimally adjust their prices in each period.

Households in China supply differentiated labor, consume, and derive utility from holding real money balances. Wages are subject to Calvo-style rigidity, so that only a fraction of households can adjust their wages optimally each period. Households save through domestic bonds issued by the Chinese monetary authority and foreign bonds denominated in ROW currency. To capture frictions in international financial markets, and induce stationarity of the model, we include a cost of deviating from the long-run net foreign asset position. Households also receive lump-sum transfers from the fiscal authority and dividends from intermediate firms. Finally, a competitive labor packer combines the differentiated labor services supplied by households into an aggregate labor input sold to intermediate firms.

4.1 Households

Chinese households supply differentiated labor services, $L_t(h)$, to the labor packer at a nominal wage $W_t(h)$. They derive utility from consumption, C_t, and from holding real money balances, M_t/P_t, while facing money demand shocks, $\mathcal{M}_t$. Households discount the future at rate β_t, which follows a stochastic process. Wages are subject to Calvo-style nominal rigidity: in each period, a household can re-optimize its wage with probability $1 - \phi_w$. Following Erceg et al. (2000), we assume a complete set of state-contingent securities that fully insure households against idiosyncratic income risk arising from wage stickiness. The household's period utility is given by

$$u\left(C_t, L_t(h), M_t\right) = \frac{C_t^{1-\sigma}}{1-\sigma} - \eta \frac{L_t(h)^{1+\nu}}{1+\nu} + \chi_m e^{\mathcal{M}_t} \ln\left(\frac{M_t}{P_t}\right)$$

Households hold domestic nominal risk-free bonds, B_{t+1}, and ROW nominal risk-free bonds, B^*_{t+1}, denominated in foreign currency. They receive labor income, $W_t(h)L_t(h)$, earn nominal returns on domestic and foreign bonds, i_t and i^*_t, obtain lump-sum transfers from the government,

T_t, and own profits from intermediate firms, Ω_t. The household's nominal flow budget constraint is therefore given by

$$P_t^c C_t + M_t + B_{t+1} + \frac{\mathcal{E}_t B_{t+1}^*}{e^{\psi_t}} + \frac{\chi}{2}\left(\frac{\mathcal{E}_t B_{t+1}^*}{i_t^*}\right)^2 = W_t(h)L_t(h) + M_{t-1} + B_t(1+i_t) + \mathcal{E}_t B_t^*(1+i_t^*) + \Omega_t + T_t$$

where P_t^c is the price index of the final consumption good, $\mathcal{E}_t$ is the nominal exchange rate, χ is an adjustment cost on internationally traded bonds that ensures stationarity and introduces frictions in international asset markets, and ψ_t is a financial shock faced only by Chinese households, generating a wedge relative to the return faced by ROW households and inducing deviations from uncovered interest parity.

The household's problem is to maximize the present discounted value of flow utility subject to the budget constraint and the labor demand schedule from the labor packer, $L_t(h) = \left(\frac{W_t(h)}{W_t}\right)^{-\epsilon_w} L_t$, where ϵ_w is the elasticity of substitution across labor varieties. The solution to the household problem is characterized by a set of Euler Equations, displayed in Appendix C.

The wage setting problem is the standard dynamic Calvo problem, with the optimality conditions shown in Appendix C.

4.2 Intermediate Firms

We consider a representative firm in China and omit the firm subscript f since all firms are identical. The production of the firm is given by

$$X_t = e^{z_t} L_t$$

where z_t is a total factor productivity (TFP) shock. Firms sell their output in both domestic and foreign markets, subject to the resource constraint

$$X_t = X_{ch,ch,t} + X_{ch,row,t} e^{\xi_{ch,row,t}}$$

where $X_{ch,ch,t}$ denotes quantities sold domestically, $X_{ch,row,t}$ quantities exported to the ROW, and $\xi_{ch,row,t}$ is a stochastic iceberg trade cost on exports, as in Alessandria and Choi (2019); Mac Mullen and Woo (2025).

Chinese firms are subject to domestic and export value added taxes, so their profits are given by,

$$\Omega_t = P_{ch,ch,t}X_{ch,ch,t}\left[1-\tau_t\right] + \mathcal{E}_t P_{ch,row,t}X_{ch,row,t}\left[1-(\tau_t-\zeta_t)\right] - W_t L_t$$

where $P_{ch,ch,t}$ is the price of goods sold domestically, $P_{ch,row,t}$ the price of goods sold in the foreign market, τ_t the domestic value added tax, and ζ_t the rebate on the export VAT. The effective value added tax on exports is given by $\tau_t - \zeta_t$.

The solution to the intermediate firm's problem yields a labor demand,

$$W_t L_t = X_t MC_t$$

where the marginal cost of production is

$$MC_t = e^{-z_t} W_t$$

Optimal pricing for domestic and export markets is given by

$$P_{ch,ch,t} = \frac{\theta}{\theta-1} \times \frac{MC_t}{[1-\tau_t]}$$

$$\mathcal{E}_t P_{ch,row,t} = e^{\xi_{ch,row,t}} \frac{\theta_t}{\theta_t - 1} \times \frac{MC_t}{[1-(\tau_t-\zeta_t)]}$$

where the elasticity of substitution for exported Chinese varieties, $\theta_t = \theta Q_t^{-\lambda}$, depends on the real exchange rate Q_t and a positive scalar $\lambda > 0$. This formulation captures reduced-form pricing-to-market frictions and generates incomplete pass-through of exchange rates to export prices (Alessandria and Choi, 2021). A symmetric specification applies to the CES aggregator of foreign varieties in China.

The CES aggregators of domestic and exported Chinese varieties are

$$X_{ch,ch,t} = \left(\int_0^1 X_{ch,ch,t}(f)^{\frac{\theta-1}{\theta}} df\right)^{\frac{\theta}{\theta-1}} \quad \text{, and} \quad X_{ch,row,t} = \left(\int_0^1 X_{ch,row,t}(f)^{\frac{\theta_t-1}{\theta_t}} df\right)^{\frac{\theta_t}{\theta_t-1}}$$

so that $X_{ch,ch,t}$ and $X_{ch,row,t}$ correspond to the bundles of Chinese intermediate goods sold domestically and exported to the ROW, respectively.

4.3 Wholesalers

In each country, there is a continuum of wholesalers indexed by $j \in [0,1]$ combining bundles of domestic and imported intermediates using a CES aggregator. Following Auclert et al. (2024), we introduce sticky input choices à la Calvo—i.e., delayed substitution in intermediate inputs. Each period, a wholesaler can re-optimize its input bundle with probability $1-\delta$, while with probability δ it must retain the input composition chosen in period $t-1$. Because input choices may remain fixed for several periods, the wholesaler's optimization problem is inherently dynamic. We provide the full description of the wholesaler block in Appendix C.

A wholesaler that is allowed to optimally choose its inputs maximizes the expected discounted flow of profits. In each period, with probability δ, the wholesaler cannot adjust its inputs. The CES aggregator of wholesaler j in the ROW is

$$\hat{D}_t^*(j) = \left[(1-\omega)^{1/\gamma}\left(\hat{X}_{row,row,t}(j)\right)^{\frac{\gamma-1}{\gamma}} + \omega^{1/\gamma}\left(\hat{X}_{ch,row,t}(j)\right)^{\frac{\gamma-1}{\gamma}}\right]^{\frac{\gamma}{\gamma-1}}$$

The dynamic problem of the wholesaler can be written as:

$$\max_{\{\hat{X}_{row,row,t}(j),\hat{X}_{ch,row,t}(j)\}} \mathbb{E}_t \sum_{s=0}^{\infty} \delta^s \beta^s \Bigg\{ P_{row,t+s}\left[(1-\omega)^{1/\gamma}\left(\hat{X}_{row,row,t}(j)\right)^{\frac{\gamma-1}{\gamma}} + \omega^{1/\gamma}\left(\hat{X}_{ch,row,t}(j)\right)^{\frac{\gamma-1}{\gamma}}\right]^{\frac{\gamma}{\gamma-1}} - P_{row,row,t+s}\hat{X}_{row,row,t}(j) - P_{ch,row,t+s}\hat{X}_{ch,row,t}(j)\Bigg\}$$

The solution to the ROW wholesaler is characterized by a pair of demand functions:

$$\hat{X}_{row,row,t}(j) = (1-\omega)\,\hat{D}^{*}_{t}(j)\left(\frac{\mathcal{P}_{row,row,t}}{\mathcal{P}_{row,t}}\right)^{-\gamma}$$

$$\hat{X}_{ch,row,t}(j) = \omega\hat{D}^{*}_{t}(j)\left(\frac{\mathcal{P}_{ch,row,t}}{\mathcal{P}_{row,t}}\right)^{-\gamma}$$

where

$$\mathcal{P}_{row,row,t} = P_{row,row,t} + \delta\beta\mathbb{E}_{t}\mathcal{P}_{row,row,t+1}$$

$$\mathcal{P}_{row,t} = P_{row,t} + \delta\beta\mathbb{E}_{t}\mathcal{P}_{ch,t+1}$$

$$\mathcal{P}_{ch,row,t} = P_{ch,row,t} + \delta\beta\mathbb{E}_{t}\mathcal{P}_{ch,row,t+1}$$

are the price recursions.

Combining the demands of optimizing and non-optimizing wholesalers from the ROW yields the aggregate demands for domestic ROW and imported Chinese intermediate bundles, as well as expenditure in the ROW:

$$X_{row,row,t} = (1-\delta)\hat{X}_{row,row,t} + \delta X_{row,row,t-1}$$

$$X_{ch,row,t} = (1-\delta)\hat{X}_{ch,row,t} + \delta X_{ch,row,t-1}$$

$$D^{*}_{t} = (1-\delta)\hat{D}^{*}_{t} + \delta D^{*}_{t-1}$$

and the ideal price index is,

$$\mathcal{P}_{row,t} = \left[(1-\omega)\,\mathcal{P}^{1-\gamma}_{row,row,t} + \omega\mathcal{P}^{1-\gamma}_{ch,row,t}\right]^{\frac{1}{1-\gamma}}$$

The presence of sticky quantities at the wholesaler level manifests as *habit formation* at the aggregate level, generating a dynamic trade elasticity. This feature is formalized in the following proposition.

Proposition 1 (Dynamic Trade Elasticity)**.** *Consider an unexpected permanent increase in the VA-TRX in China, ζ (i.e. export subsidy shock). The time-t elasticity of Chinese exports to a permanent* *change in the VATRX is,*

$$\frac{\partial \tilde{x}_{ch,row,t}}{\partial \tilde{\zeta}} = \gamma(1-\delta)\left(\sum_{i=0}^{t} \delta^i\right)$$

where $\tilde{x}^*_{ch,row,t}$ and $\tilde{\zeta}$ are log-deviations from the steady state Chinese exports and VATRX, respectively. The habit, δ, governs the dynamic trade elasticity, which converges to γ.

See Appendix C.4 for the proof. Proposition 1 is particularly useful because it allows us to pin down the trade elasticity, γ, and the degree of habits, δ, using the estimated trade elasticity from Section 3. We make use of this in the estimation of the model in Section 6.

4.4 Retailers

There is a continuum of retailers, indexed by $i \in (0, 1)$, producing varieties of the final good using the output of the wholesalers, $D_t(i)$. The production function is

$$Y_t(i) = D_t(i).$$

Retailers face Calvo-style price stickiness: in each period, only a fraction $1 - \phi_p$ of firms can adjust their price. They face a demand for each variety: $Y_t(i) = \left(\frac{P_t(i)}{P_t^c}\right)^{-\varepsilon_p} Y_t$, where ε_p is the elasticity of substitution across final good varieties. The problem yields the standard solution to Calvo price stickiness, described in Appendix C.

4.5 Fiscal Authority in China

In China, the fiscal authority collects the VAT and provides VATRX. The revenue collected each period is redistributed lump-sum to households, T_t, ensuring a balanced budget. The government budget constraint is

$$T_t = P_{ch,ch,t} X_{ch,ch,t} \tau_t + \mathcal{E}_t P_{ch,row,t} X_{ch,row,t} (\tau_t - \zeta_t),$$

where τ_t is the domestic VAT and ζ_t is the VATRX.

We model VAT and VATRX as stochastic processes:

$$\tau_t = \mu_\tau(1-\rho_\tau) + \rho_\tau \tau_{t-1} + \sigma_\tau \varepsilon_{\tau,t}, \tag{5}$$

$$\zeta_t = \mu_\zeta(1-\rho_\zeta) + \rho_\zeta \zeta_{t-1} + \sigma_\zeta \varepsilon_{\zeta,t}. \tag{6}$$

where $\varepsilon_{\tau,t} \sim \mathcal{N}(0,1)$ and $\varepsilon_{\zeta,t} \sim \mathcal{N}(0,1)$, μ_τ and μ_ζ are the unconditional mean of the VAT and VATRX processes and ρ_τ, ρ_ζ, σ_τ and σ_ζ ther persistence and variance.

For simplicity, we assume that there is no fiscal authority in the ROW.

4.6 Monetary Authority

We assume that China's monetary authority follows a money growth rule, consistent with the institutional framework of China's monetary policy (Li-gang and Zhang, 2010; Li and Liu, 2017; Chen et al., 2018; Cho et al., 2025). To capture China's managed exchange rate regime, we assume that the monetary authority stabilizes the nominal exchange rate by adjusting money growth. In practice, China also intervenes in foreign exchange markets, inducing deviations from uncovered interest parity. While we do not model these interventions directly, we interpret part of the exogenous uncovered interest parity wedge, ψ_t, as capturing these interventions. We have experimented directly modeling foreign exchange interventions following Cavallino (2019), but found similar results.[16] For this reason, we opt for capturing foreign exchange interventions through the exogenous wedge, ψ_t.

The money growth rule in China is

$$\Delta M_t = \rho_m \Delta M_{t-1} + \phi_{m,\pi}\pi_{t-1} + \phi_{m,gdp}\Delta gdp_t + \phi_{m,e}\Delta e_t + \sigma_m \varepsilon_{m,t} \tag{7}$$

where $\pi_{t-1} = \ln(\Pi_{t-1})$ is the inflation rate, $\Delta gdp_t \equiv \ln(X_t) - \ln(X_{t-1})$ the change in log real GDP, $e_t \equiv \ln(\mathcal{E}_t)$ the log of the nominal exchange rate, and $\varepsilon_{m,t} \sim \mathcal{N}(0,1)$ are monetary shocks.

[16]Specifically, we allow the monetary authority to hold domestic bonds $B_{cb,t+1}$ and foreign bonds $B^*_{cb,t+1}$ under a zero-capital strategy $B_{cb,t+1} = -\mathcal{E}_t B^*_{cb,t+1}$, with a rule for interventions $B^*_{cb,t+1} = \rho_{fxi} B^*_{cb,t} - \phi_{fxi}\Delta e_t + \sigma_{fxi}\varepsilon_{fxi,t}$, $\varepsilon_{fxi,t} \sim \mathcal{N}(0,1)$. We found results similar to the benchmark model under this specification, even under a strong degree of frictions in international financial markets, $\chi >> 0$.

In contrast, the ROW monetary authority follows a standard Taylor rule that stabilizes inflation:

$$i_t^* = \ln\left(\frac{1}{\beta}\right)(1-\rho_{i^*}) + \rho_{i^*} i_{t-1}^* + (1-\rho_{i^*})\,\phi^* \pi_t^* + \sigma_{i^*}\varepsilon_{i^*,t}$$

where $\phi^* > 0$ is the Taylor rule coefficient, ρ_{i^*} and σ_{i^*} are the persistence and standard deviation of the rule, and $\varepsilon_{i^*,t} \sim \mathcal{N}(0,1)$ are shocks.

4.7 Shock Processes

In addition to VAT and VATRX shocks in China, the model features ten additional stochastic processes. Each country has processes for total factor productivity (TFP), z_t and z_t^*, iceberg trade costs, $\xi_{ch,row,t}$ and $\xi_{row,ch,t}$, and discount factors, β_t and β_t^*. Furthermore, there are stochastic processes for China's money growth, ΔM_t, and the ROW interest rate, R_t^*. Chinese households also face shocks to money demand, $\mathcal{M}_t$, and to the return on foreign bonds, ψ_t.

Let $\Theta \in \{z, z^*, \xi_{ch,row}, \xi_{row,ch}, \psi, \beta, \beta^*, m, i^*, \mathcal{M}\}$ denote the vector of all these stochastic processes. Each process is assumed to follow an AR(1) specification:

$$\Theta_t = \mu_\Theta (1-\rho_\Theta) + \rho_\Theta \Theta_{t-1} + \sigma_\Theta \epsilon_{\Theta,t}$$

where μ_Θ, ρ_Θ and σ_Θ are the unconditional mean, persistence and standard deviation of the processes, and $\varepsilon_{\Theta,t} \sim \mathcal{N}(0,1)$ are independent innovations.

4.8 Market Clearing

Market clearing for final goods in each country requires

$$Y_t = C_t, \quad \text{, and} \quad Y_t^* = C_t^*.$$

while market clearing in international bond markets requires

$$B_{t+1}^* + B_{t+1}^{**} = 0,$$

where B_t^* and B_t^{**} denote the holdings of foreign bonds by Chinese and ROW households, respectively.

Finally, China's consolidated budget constraint implies market clearing in international goods markets:

$$B_{t+1}^* Q_t - B_t^* Q_t \frac{R_t^*}{\Pi_t^*} = Q_t P_{ch,row,t} X_{ch,row,t} - P_{row,ch,t} X_{row,ch,t},$$

The ROW budget constraint is redundant by Walras' Law.

5 VATRX as a Quasi-Fiscal Devaluation Tool

We now study the transmission of VATRX as a quasi-fiscal devaluation tool. The fiscal-devaluation literature—developed primarily in the context of currency unions—shows that coordinated changes in import tariffs and export subsidies can replicate the real effects of a nominal devaluation (Farhi et al., 2014; Erceg et al., 2023). To interpret China's VATRX policy through this lens, we compare the responses generated by our benchmark model with those implied by standard fiscal-devaluation schemes, highlighting the distinctive role played by export-rebate adjustments. We then examine how the trade elasticity shapes the transmission mechanism. In particular, we show that the strength of a quasi-fiscal devaluation increases with the average trade elasticity. Furthermore, abstracting from the dynamic nature of the trade elasticity results in overestimating the response of GDP in the short-run and underestimating the response in the medium run, while the effects on consumption are similar between dynamic and static trade elasticity models.

5.1 Quasi-Fiscal Devaluations

We parametrize the quasi-fiscal devaluation using the estimated structural parameters of the benchmark model from Section 6. We then contrast the benchmark VATRX experiment with two alternative scenarios. First, we consider a full fiscal devaluation—combining an increase in export subsidies an import tariffs adjusted to keep the government budget-neutral—while keeping all other parameters unchanged. Second, we examine this full fiscal devaluation under a fixed nominal exchange rate.[17] This comparison allows us to isolate the mechanisms through which VATRX operates relative to canonical fiscal-devaluation designs.

We study, across all cases, the impulse responses of key variables to a shock to the VATRX. Figure 6 plots the dynamic responses. Panel (a) displays the responses of the VATRX and the effective export VAT, shown by the black solid line and black-dashed-dotted line respectively. These are identical across all scenarios by construction.

We begin with the full fiscal devaluation under a fixed nominal exchange rate—shown in green dashed lines in Panels (b) through (l)—as this case is most comparable to the settings analyzed in Farhi et al. (2014) and Erceg et al. (2023). Panels (b) and (c) show that the shock lowers China's export price and raises export quantities. In turn, Panels (d)–(e) illustrate that GDP, net exports, and consumption all increase. This pattern captures the core mechanism emphasized by the fiscal-devaluation literature: when nominal depreciation is unavailable, coordinated export subsidies and import tariffs adjustments can deliver expansionary effects similar to a devaluation.

A fiscal devaluation raises GDP through both consumption and net exports. Consumption increases because the policy lowers the expected path of the real interest rate in China— the standard New Keynesian transmission channel (Gali, 2008)—driven by the inflationary effects of the policy shown in Panel (i). Net exports rise because export subsidies directly stimulate exports, while higher import tariffs discourage imports. As a result, the price of imported goods in China increases and the foreign price of Chinese exports falls, leading to a deterioration in China's terms of trade, shown in Panel (j). Finally, the inflationary nature of the policy causes an appreciation of China's real exchange rate, as displayed in Panel (k).

The red dash-dotted lines in Figure 6 show the case of a full devaluation under China's managed exchange rate regime. The main differences relative to the fixed-exchange-rate case appear in the behavior of consumption, nominal and real interest rates, and inflation. When the exchange rate is managed, the monetary authority reacts to inflation, which leads to a weaker fall in the nominal interest rate. Consequently, inflation rises by less than in the fixed nominal exchange rate case, and the real interest rate falls less in the short run. This more muted decline

[17] In the fixed nominal exchange rate case we replace the money growth rule equation 7 by $\Delta e_t = 0$.

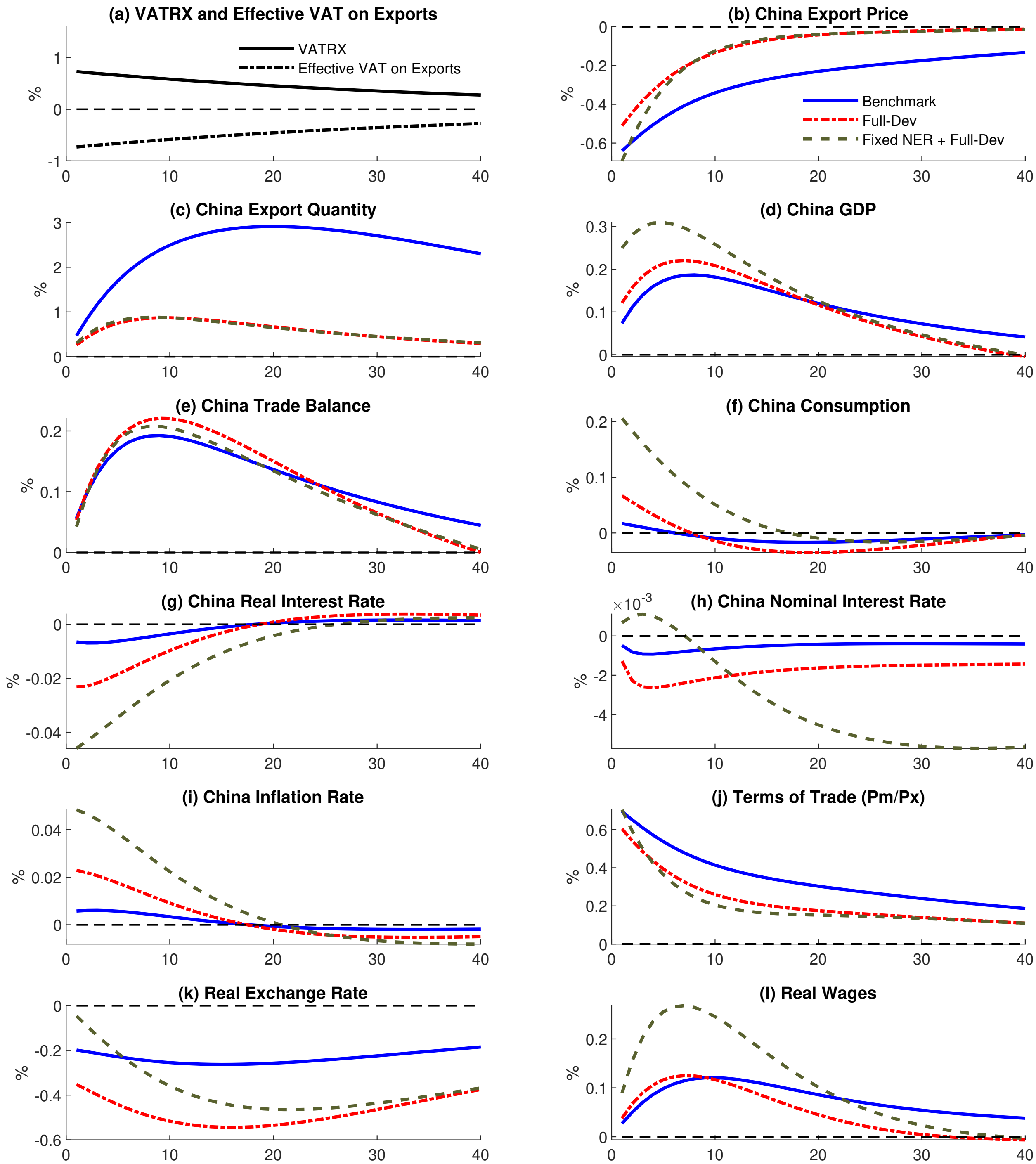


Figure 6: VATRX as a Quasi-Fiscal Devaluation Tool

in the real interest rate translates into a weaker consumption response and thus a more moderate increase in GDP under a managed exchange rate regime. Nonetheless, the overall GDP effect remains sizable, driven primarily by the expansion of net exports.

The case of a quasi-fiscal devaluation appears in the solid blue lines. Its dynamics closely resemble those of the full fiscal devaluation under the managed exchange rate regime. This similarity reflects the inflation-stabilization role of monetary policy discussed above. However, the quasi-fiscal devaluation is slightly less expansionary because it generates a smaller increase in inflation. Inflation rises less in this scenario because, unlike the full fiscal devaluation, the quasi-fiscal variant does not include changes in import tariffs. Despite this difference in magnitude, the figure shows that quasi-fiscal devaluations transmit through the economy in much the same way as full fiscal devaluations.

Overall, these results demonstrate that China's VATRX policy replicates many of the key transmission channels emphasized in the fiscal-devaluation literature, even though it operates solely through the export-subsidy component. While the absence of import-tariff adjustments makes the quasi-fiscal devaluation somewhat less expansionary than a full fiscal devaluation, its macroeconomic effects remain sizable—especially through net exports—and closely resemble those under more comprehensive policy packages. This highlights that, in environments where monetary policy is constrained or the exchange rate is managed, export-rebate adjustments can serve as a practical and effective, though not all-encompassing, tool for stabilization.

5.2 The Role of the Trade Elasticity

The standard practice in the business-cycle literature is to use a static trade-elasticity parameter calibrated to match the short-run response of trade flows. For example, Farhi et al. (2014) assume a value of 1.3, while Erceg et al. (2023) use 1.25. We now study how the trade elasticity governs the transmission of quasi-fiscal devaluations—very close to the short-run elasticity of 1.19 that we estimate in Section 3. We now study how the trade elasticity governs the transmission of quasi-fiscal devaluations.

Figure 7 compares impulse responses in our Benchmark model with two counterfactual cases in which we shut down the delayed substitution mechanism and impose a static trade elasticity. The first counterfactual follows the conventional approach in the macro–open-economy literature by fixing the elasticity at its short-run value—1.19 in our case as estimated in Section 3 and similar to values in Farhi et al. (2014) and Erceg et al. (2023)—shown by the red dash–dot line. The second sets the elasticity to the one-year substitution value of 4.31, shown in green dashed lines, which is closer to the average trade elasticity estimated in Section 3. The comparison makes clear that the responses of GDP, the trade balance, and consumption are all increasing in the trade elasticity: models with higher static elasticities generate stronger effects of quasi-fiscal devaluations.

Building on the comparison across different static trade elasticity specifications, we explain the mechanisms behind these results. A VATRX shock lowers the export price of Chinese intermediates (Panel b). The higher the trade elasticity, the stronger the substitution of ROW agents toward Chinese intermediates. As a result, a higher elasticity amplifies the demand shift toward Chinese goods, which raises Chinese production and labor demand and, in turn, pushes nominal wages upward. The resulting increase in nominal wages generates higher domestic inflation, and therefore a larger decline in the real interest rate. This mechanism makes the shock more expansionary for consumption. Consequently, the larger the average trade elasticity, the stronger the effects on domestic production, exports and net exports (Panels (c) and (e)), and the larger the expansion in consumption (Panel f) and GDP (Panel d).

While the average trade elasticity governs the overall strength of the transmission mechanism, the dynamic path of the elasticity shapes the time profile of export quantities, net exports and GDP. The static model produces a stronger expansion in exports and net exports on impact but a noticeably weaker response in the medium run. This pattern translates into a larger initial GDP response under static trade elasticity but a slightly weaker effect at longer horizons.

Overall, the results show that the standard macro practice of using a static trade elasticity overstates the initial response of GDP and net exports, and understates the effects on the medium-run. Capturing both the average level and the dynamic path of the trade elasticity is therefore essential for understanding the full transmission of quasi-fiscal devaluations to external adjustment, and overall economic activity.

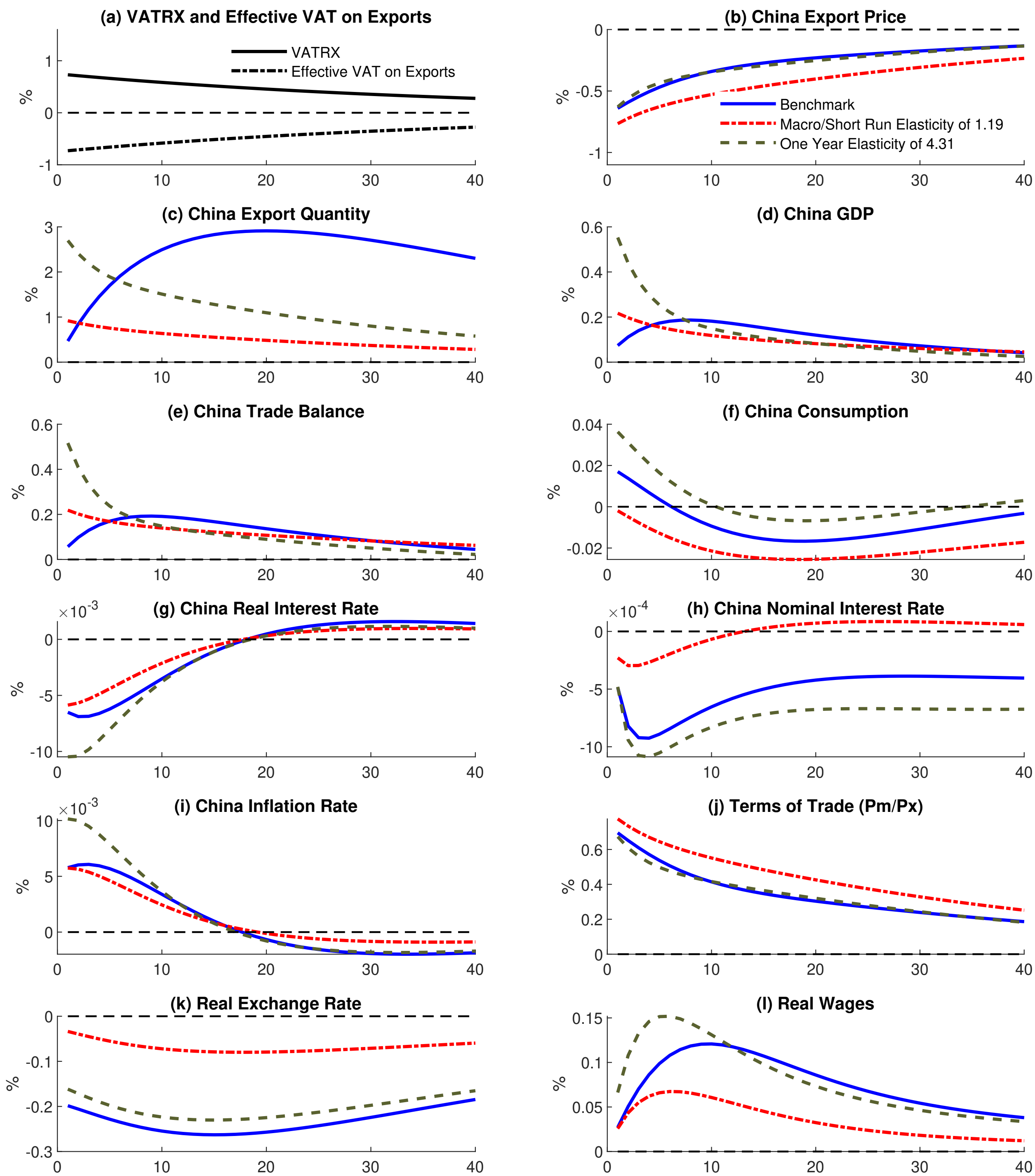


Figure 7: Trade Elasticities and Quasi-Fiscal Devaluations

6 Estimation of Full Model

In this section we estimate the full structural model that links China's VATRX policy to macroeconomic dynamics. The estimation proceeds in two steps. First, we fix a set of parameters that can be disciplined using external empirical evidence or calibration targets. Second, we estimate the remaining parameters with Bayesian methods, using Chinese and ROW aggregate data. The Bayesian framework not only incorporates prior information and quantify uncertainty, but also ensures that the model replicates the time series of key variables used for the counterfactual analysis in Section 7.

6.1 Parameters Externally Set

A subset of model parameters is disciplined outside the Bayesian estimation block using either direct empirical evidence, steady-state targets, or values standard in the international macroeconomics literature. Table 2 summarizes all externally set parameters, together with the moments or sources that guide their calibration.

Preferences and household-side parameters are chosen to match standard macroeconomic statistics. The mean discount factors in China and the ROW, μ_β and μ_{β^*}, are calibrated to imply an average annual real interest rate of 4 percent. Risk aversion is set to $\sigma = 2$, consistent with an intertemporal elasticity of substitution of 0.5. Hours worked in both regions guide the calibration of consumption weights (η_C and η_R), while real money balances in China determine the normalization of χ_m. The Frisch elasticity is set to unity through the parameter ν.

On the production and trade side, the elasticity of substitution across varieties, $\theta = 4$, is calibrated to generate a producer markup of roughly 33 percent. The pricing-to-market friction $\lambda = 0.75$ is chosen to match the empirical pass-through of exchange rate movements into export prices. Home-bias parameters (γ^C and γ^R) discipline steady-state trade volumes in China and the ROW.

The trade elasticity γ and the quantity-adjustment parameter δ are set to match the long and short-run partial equilibrium trade elasticity estimated in Section 3 following proposition 1.

We estimate the elasticity in the model via the simulated method of moments (SMM), targeting horizons 0 to 25 of the estimated trade elasticity coefficients from the data. Each moment is weighted by the inverse of its variance, constructed from the corresponding confidence intervals. Let $\mathbf{\Psi}(\delta, \gamma)$ denote the mapping from (δ, γ) to the model trade elasticity moments, and $\hat{\mathbf{\Psi}}$ the corresponding empirical estimates. Then the SMM estimators $\hat{\delta}$ and $\hat{\gamma}$ solve

$$J = \min_{\delta,\gamma} [\hat{\mathbf{\Psi}} - \mathbf{\Psi}(\delta, \gamma)]' \mathbf{V}^{-1} [\hat{\mathbf{\Psi}} - \mathbf{\Psi}(\delta, \gamma)]$$

where $\mathbf{V}$ is a diagonal matrix with the sample variances of $\hat{\mathbf{\Psi}}$ on the diagonal. We estimate a degree of habits $\hat{\delta} = 0.930$ and a long-run trade elasticity $\hat{\gamma} = 17.16$. These values imply a short-run partial equilibrium trade elasticity of 1.19, very close to the values of 1.3 and 1.25 in Farhi et al. (2014) and Erceg et al. (2023), respectively. Figure 8 compares the dynamic trade elasticity in the model and in the data, showing that the model closely tracks the empirical counterpart. In the next section, we embed this mechanism in a full quantitative model that preserves the same partial-equilibrium dynamic trade elasticity.

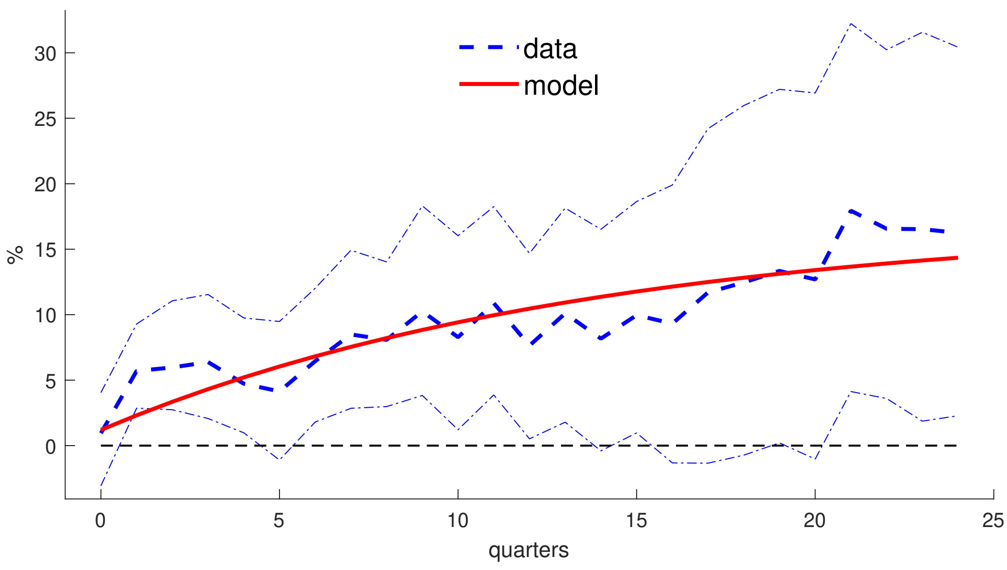


Figure 8: Partial Equilibrium Dynamic Trade Elasticity: Data and Model

Policy and institutional parameters are set to match observed features of the Chinese tax system and the conduct of monetary policy in the ROW, which we take as driven by the US. The mean VAT and VATRX processes, μ_τ and μ_ζ, are pinned down by their averages in the data.

Monetary policy in the ROW follows a Taylor rule with an inflation coefficient $\phi_\pi = 1.5$. Wage and price stickiness parameters (ϕ_w and ϕ_p) are set to 0.6 and 0.9, respectively, following the estimates by Li and Liu (2017).

Finally, several steady-state normalizations close the calibration. The mean TFP levels in China and the ROW, μ_z and μ_z^*, match China's share of world GDP. Bond-adjustment costs χ ensure stationarity of net foreign assets, and steady state net foreign assets is normalized to zero ($\bar{B} = 0$).

Collectively, these calibrated parameters anchor the model in well-established empirical targets while allowing the Bayesian estimation to focus on the structural elasticities most directly tied to the macroeconomic transmission of VATRX policy and other shocks.

Table 2: Calibrated Parameters

Parameter		Value	Target Moment or Source
Mean discount factor process in China and ROW	μ_β & μ_{β^*}	0.99	mean annual interest rate of 4%
Risk aversion	σ	2	intertemporal elasticity of substitution of 0.5
Preference of real balances in China	χ_m	0.1	normalization steady state real balances
Weight on consumption in China	η_C	1	hours worked of 1/3
Weight on consumption in ROW	η_R	0.2513	hours worked of 1/3
Macro Frisch Elasticity	ν	1	Frisch elasticity
Elasticity of substitution across varieties	θ	4	producer markup of 33%
Pricing-to-market friction	λ	0.75	75% exchange rate pass-though to prices (Alessandria and Choi, 2021)
Trade Elasticity	ρ	17.16	PE trade elasticity estimated in Section 3
Quantity stickiness	δ	0.93	PE trade elasticity estimated in Section 3
Home bias in China	γ^C	0.25	trade-to-GDP ratio of 50% in China
Home bias in ROW	γ^R	0.0628	zero net exports in steady state
Mean VAT process	μ_τ	0.168	mean VAT value in data
Mean VATRX process	μ_ζ	0.062	mean VATRX value in data
Mean TFP process in ROW	μ_z^*	1	normalization
Mean TFP process in China	μ_z	0.067	China's GDP in world GDP
Wage stickiness	ϕ_w	0.6	Li and Liu (2017)
Price stickiness	ϕ_p	0.9	Li and Liu (2017)
Taylor Rule inflation targeting in ROW	ϕ_π	1.5	standard value in literature
Bond adjustment cost	χ	0.001	stationarity of NFA
Steady state NFA	$\bar{B}$	0	normalize zero steady state NFA

6.2 Bayesian Estimation

The remaining parameters are estimated using a Bayesian approach. These include parameters related to the 12 shock processes, for which we feed the time series of twelve variables.

6.2.1 Observables and Choice of Priors

We estimate the model using twelve quarterly macroeconomic and policy time series spanning 2004Q1–2018Q4, presented in Figure D1 in Appendix E. For China, we include the VAT and VATRX series, real export and import growth, nominal GDP, CPI inflation, M2 growth, the real exchange rate, and the terms of trade. For the ROW, we use nominal GDP, CPI inflation, and the nominal interest rate. The ROW nominal interest rate and inflation rate are taken from the United States (federal funds rate and CPI inflation), reflecting the central role of U.S. monetary conditions in global financial markets. ROW GDP, in turn, is constructed as a trade-weighted average of output across China's major trading partners.[18]

The set of twelve observables provide information to identify the shock processes driving both domestic and external dynamics. For China and the ROW, movements in nominal GDP, inflation, and trade flows discipline the TFP and iceberg trade-cost shocks: persistent co-movements in output and export/import quantities pin down the autoregressive components of both productivity and trade frictions, while higher-frequency deviations in bilateral trade volumes inform their innovation variances. Discount-factor shocks in China and the ROW are identified through the joint behavior of GDP and the trade balance. Because discount-factor disturbances generate intertemporal wedges rather than static trade distortions, the model leverages fluctuations in net exports to identify the persistence and volatility of these shocks.

The China-specific shocks—VAT, VATRX, UIP deviations, money-growth, and money-demand disturbances—are identified through observables that are particularly sensitive to nominal and policy wedges. VAT and VATRX series directly anchor the respective policy processes, with export and import quantities and prices providing additional discipline by linking tax-policy dis-

[18]See Section 2 for details.

turbances to trade volumes and relative price movements. High-frequency exchange-rate movements that cannot be reconciled with structural interest-rate differentials are absorbed by the UIP shock process, and this reflect both external shocks and foreign direct interventions by China. China's inflation, M2 growth, and the behavior of the real exchange rate are informative for identifying money-growth and money-demand shocks, which affect nominal variables and external adjustment. Finally, the ROW Taylor-rule shock is identified through the joint dynamics of U.S. inflation and the U.S. nominal interest rate.

The parameters of the model are estimated using Bayesian techniques, which require the specification of prior distributions for each structural parameter. Table 3 summarizes the priors and resulting posterior distributions. For autoregressive parameters, such as the persistence of TFP, trade-cost, VAT, VATRX, money-growth, money-demand, and discount-factor shocks, we employ Beta distributions with means reflecting the conventional persistence found in the literature and moderate standard deviations to allow for data-driven adjustments. Standard deviations of shocks are assigned inverse-gamma priors, which are standard in Bayesian DSGE estimation. We select relatively tight priors for the VAT and VATRX shocks to reflect their observed dynamics. For VATRX, we set a prior on the persistence that captures a highly persistent process, consistent with the adjustments observed in the data. In contrast, the VAT series exhibits fluctuations around a mean value of approximately 0.16 with relatively short-lived deviations, motivating a prior with lower persistence. This choice ensures that the priors are informative yet flexible enough to accommodate the empirical behavior of both policy instruments.

Priors on monetary policy parameters, including the coefficients on inflation, exchange rate, and GDP growth in the Chinese money-growth rule, are chosen based on estimates in previous studies. In particular, we set priors for the autocorrelation of money growth, and the responses of money growth to changes in inflation and GDP according to Chen et al. (2018).[19] We set the prior mean and standard deviation for the response of money growth to changes in the nominal exchange rate to the be the same as for the response to inflation, hence assuming ex-ante the same weight on stabilizing inflation and the nominal exchange rate. Nevertheless, we allow for sufficient variance to let the data refine estimates.

[19]See their Table A1.

Table 3: Prior and Posterior Distributions of Estimated Parameters

Parameter	Prior Mean	Posterior Mean	90% HPD Interval		Prior Dist.	Prior Std. Dev.
			Lower	Upper		
ρ_z	0.900	0.9263	0.8975	0.9550	Beta	0.075
ρ_{z^*}	0.900	0.8657	0.8159	0.9166	Beta	0.075
σ_z	0.100	0.0941	0.0773	0.1101	Inv-Gamma	0.050
σ_{z^*}	0.100	0.0964	0.0747	0.1180	Inv-Gamma	0.050
$\rho_{\xi_{ch,row}}$	0.900	0.9657	0.9347	0.9900	Beta	0.075
$\sigma_{\xi_{ch,row}}$	0.150	0.0877	0.0712	0.1034	Inv-Gamma	0.050
$\rho_{\xi_{row,ch}}$	0.900	0.9375	0.9021	0.9740	Beta	0.075
$\sigma_{\xi_{row,ch}}$	0.150	0.0802	0.0655	0.0945	Inv-Gamma	0.050
ρ_ψ	0.800	0.5614	0.4957	0.6287	Beta	0.050
σ_ψ	0.010	0.0464	0.0374	0.0552	Inv-Gamma	0.025
ρ_τ	0.050	0.1126	0.0308	0.1905	Beta	0.025
σ_τ	0.001	0.0023	0.0019	0.0026	Inv-Gamma	0.010
ρ_ζ	0.970	0.9754	0.9533	0.9987	Beta	0.025
σ_ζ	0.010	0.0073	0.0062	0.0083	Inv-Gamma	0.005
ρ_{i^*}	0.800	0.8190	0.7901	0.8466	Beta	0.025
σ_{i^*}	0.001	0.0041	0.0034	0.0047	Inv-Gamma	0.010
ρ_m	0.500	0.4316	0.3360	0.5356	Beta	0.075
$\phi_{m,\pi}$	0.600	0.5676	0.4555	0.6806	Beta	0.075
$\phi_{m,e}$	0.600	0.4820	0.3867	0.5757	Beta	0.075
σ_m	0.050	0.0200	0.0168	0.0231	Inv-Gamma	0.050
$\phi_{m,gdp}$	0.100	0.0594	0.0010	0.1184	Beta	0.075
$\rho_{\mathcal{M}}$	0.900	0.9862	0.9830	0.9894	Beta	0.050
$\sigma_{\mathcal{M}}$	0.500	0.4693	0.3856	0.5459	Inv-Gamma	0.100
ρ_β	0.900	0.7968	0.7459	0.8514	Beta	0.050
σ_β	0.100	0.0328	0.0257	0.0392	Inv-Gamma	0.050
ρ_{β^*}	0.900	0.6716	0.6237	0.7235	Beta	0.050
σ_{β^*}	0.100	0.0278	0.0227	0.0319	Inv-Gamma	0.050

We set the prior for the persistence of the trade shocks and UIP processes following estimates in the literature (Alessandria and Choi, 2021; Mac Mullen and Woo, 2025; Bodenstein et al., 2024). The discount-factor shocks for China and the ROW are given Beta priors with mean 0.9, while their innovation variances are set to small inverse-gamma priors, reflecting moderate uncertainty about the speed of consumption-smoothing behavior. Overall, the priors are informative enough to incorporate institutional knowledge and stylized facts, yet sufficiently flexible to allow the rich set of observables to meaningfully update beliefs about persistence and volatility across all shock processes.

6.2.2 Results

Posterior Distributions. Table 3 reports the posterior means and 90% highest posterior density (HPD) intervals for all estimated parameters. Overall, the posterior distributions indicate that the data strongly inform the key structural parameters. The TFP processes for both China and the ROW are tightly estimated, with high posterior persistence (0.926 and 0.865, respectively). This suggests that low-frequency co-movements in output and trade are strongly influenced by persistent productivity fluctuations. Similarly, the bilateral trade-cost shocks exhibit high posterior persistence, reflecting the slow-moving nature of trade wedges needed to match observed trade dynamics, consistent with findings in the literature (Alessandria and Choi, 2019; Mac Mullen and Woo, 2025).

Shocks specific to China are also well-identified. The UIP wedge exhibits a reduction in persistence relative to its prior (from 0.80 to 0.56) and a higher innovation variance, implying that high-frequency exchange-rate movements are largely absorbed by this channel, consistent with the role of active foreign exchange management. The VAT process has low persistence (0.11) while the VATRX process is highly posterior persistent (0.97), and their variances are tightly estimated. Monetary policy parameters in China, including the responses of money growth to inflation, exchange rates, and GDP growth, are tightly identified, with the money-demand shock showing high persistence and substantial volatility, which helps to disentangle nominal liquidity effects from other channels.

The discount-factor process is found to be more persistence in China than in the ROW (0.79 vs 0.67), helping the model match differential behavior of relative GDP, inflation and net exports across countries and separate demand-driven trade balance fluctuations from those arising from trade-cost or VATRX shocks.

Convergence and Identification Diagnostics. We assess convergence and identification to ensure the validity of our Bayesian estimation. Diagnostic tests show that the MCMC sampler converges satisfactorily and that the model's observables provide sufficient information to identify all structural parameters. These results confirm that the estimated parameters are both well-identified and supported by the data. Detailed convergence statistics, sensitivity analyses, and complementary identification checks are reported in Appendix D.1.

7 Quantifying the role of VATRX as a Quasi-Fiscal Devaluation Tool

We use the global financial crisis (2008–2012) as a case study to evaluate the countercyclical role of the VATRX policy in China. To this end, we perform a counterfactual exercise using our estimated model: we fix the VATRX at its pre-crisis level (second quarter of 2008) and compare the resulting dynamics with the actual data, where the VATRX increased by 1.6 percentage points. This approach allows us to isolate the effects of VATRX adjustments on China's external and aggregate economy.

Figure 9 presents the dynamics of key variables in the benchmark model and the counterfactual. Panel (a) shows the evolution of the effective VAT on exports, while Panel (b) illustrates that the VATRX policy effectively depreciated the real exchange rate on exports relative to the counterfactual. This depreciation stimulated real exports, as seen in Panel (g), which increased on average by 6.33 percentage points (semi-elasticity ≈ 4.00) during the crisis.[20] Panel (c) shows that GDP rose on average by 0.48 percentage points (semi-elasticity ≈ 0.30). Finally, the terms of trade deteriorated by 1.15 percentage points (semi-elasticity ≈ 0.72), and the real exchange rate appreciated by 0.67 percentage points (semi-elasticity ≈ 0.42).[21]

[20] This value is obtained from comparing the change between 2008Q2 and 2011Q1.

[21] The semi-elasticities are presented in Table E1 under the column corresponding to the Benchmark model.

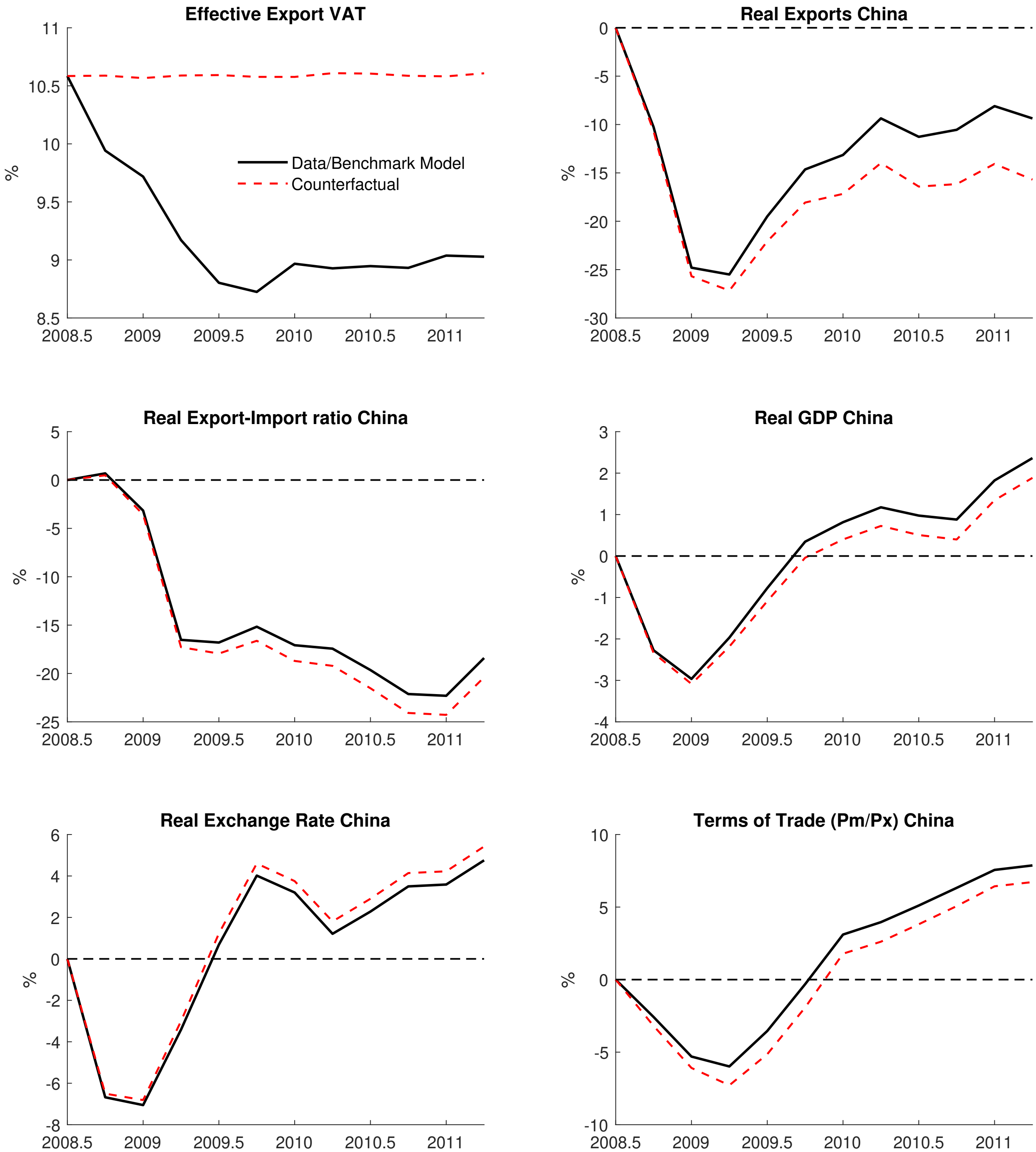


Figure 9: Global Financial Crisis: The Importance of VATRX Policy

These results are quantitatively significant and fall between the estimates reported in the literature: Farhi et al. (2014) report stronger effects of fiscal devaluations, while Erceg et al. (2023) find weaker impacts. The intermediate magnitude of our results reflects a combination of features: the VATRX shock is highly persistent, as in Farhi et al. (2014), but price pass-through is also high, similar to Erceg et al. (2023). Overall, these findings highlight VATRX adjustments as an effective quasi-fiscal devaluation instrument, through which China stabilized external demand and supported exports during the global financial crisis.

To further illustrate the role of trade elasticities in amplifying VATRX effects, we perform an additional counterfactual in which the trade elasticity is held fixed at its short-run value of 1.19, reflecting the smaller average elasticity typically assumed in the literature.[22] Figure E1 compares the dynamics of key variables under this static-elasticity scenario with the benchmark model, and Table E1 presents a comparison of the semi-elasticities in the two cases. While the short-run effects on exports are similar across models, the static trade elasticity model significantly unpredicted the medium-run effects. Consistent with the analysis in Section 5.2, fixing the trade elasticity at the short-run overestimates the short-run effects on net exports and GDP, while it underestimates its medium-run effects.

These results emphasize the importance of capturing both the level and dynamic path of the trade elasticity when assessing the effects of quasi-fiscal devaluations. The benchmark model, with a dynamic and higher average elasticity, predicts stronger medium-run effects on exports, net exports, and GDP. This reinforces the broader finding that trade-policy-driven quasi-fiscal devaluations, such as VATRX adjustments, can be a powerful countercyclical tool, especially when their transmission is amplified by realistic trade elasticities.

8 Concluding Remarks

This paper investigates the role of trade-policy-driven fiscal devaluations interventions in shaping macroeconomic outcomes, with a particular focus on China's VATRX policy. Using highly

[22] This model assumes all the parameters are the same as the estimated ones in the Benchmark model except that there is no delayed substitution, i.e. $\delta = 0$, and that the trade elasticity is 1.19. We re-estimate the path of shocks in this version so that we exactly match the same observables as in the Benchmark model.

disaggregated VATRX data, we estimate the dynamic effects of rebate changes on export quantities and prices. Permanent reductions generate a dynamic quantity response—with trade elasticities rising from 3.8 on impact to 16.8 in the long run—and complete pass-through to export prices. We develop a two-country DSGE model capturing the interaction of fiscal, monetary, and trade policies, and estimate it using a Bayesian framework. Our estimation highlights the importance of persistent shocks, trade frictions, and policy instruments in determining the dynamics of key macroeconomic variables.

Using the estimated model, we quantify the effects of VATRX adjustments during the global financial crisis (2008–2012). Counterfactual experiments show that the VATRX increase raised China's exports, GDP, and terms of trade significantly, effectively acting as a quasi-fiscal devaluation. These results underscore the macroeconomic impact of VATRX adjustments and illustrate how trade policies can serve as countercyclical instruments in response to external shocks.

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Appendix

A Data Appendix

This appendix provides additional details on the data sources and construction procedures used in the empirical analysis. Our main dataset combines product-level export information from Chinese Customs Statistics, VAT rebate schedules from official Chinese government records, and customs Data. We further complement these sources with aggregate macroeconomic time series.

VAT and Rebate Data. As discussed in Section 3, our key policy variation comes from China's incomplete VAT rebates on exports. The underlying data, sourced from official Chinese government records, specify both domestic VAT rates and export rebate rates at the HS 10-digit level, including exact implementation dates for all policy changes. Since our trade data operates at the HS 6-digit level, we must reconcile this granularity difference. We aggregate the finer VAT data by taking simple averages across HS 10-digit products within each HS 6-digit category, preserving most product-level heterogeneity while ensuring compatibility with available export data.[23] For our aggregate analysis, we then construct trade-weighted measures using lagged export shares from Chinese customs—a timing choice that mitigates potential simultaneity between policy changes and export performance.[24]

Trade Flows and Prices. Our trade data comes from Chinese customs via Trade Data Monitor, providing quarterly export values and quantities at the HS 6-digit level. We construct unit values by dividing nominal values by quantities within each product-quarter cell. For our empirical analysis, we impose an important sample restriction: we include only products with consistent quantity units throughout our sample period. This ensures that apparent changes in prices or quantities reflect true economic responses rather than changes in measurement units.

[23] Since the official data has some data gaps on the total VAT, we extrapolate them using the reported average for those products at the 4-digit level.

[24] We use all reported exports for each HS 6-digit product in constructing these weights, regardless of destination or export type.

Together, these data sources allow us to trace policy changes from their implementation at the product level through to their aggregate macroeconomic effects, providing both the micro-level variation for identification and the macro-level moments for structural estimation.

Aggregate Macroeconomic Data. For our DSGE model estimation, we construct quarterly macroeconomic series from multiple sources. From the International Monetary Fund (IMF), we obtain real GDP, consumption, and both real and nominal trade flows for China and our rest-of-world (ROW) aggregate. We complement these with real effective exchange rate data from the IMF and terms of trade indices from the World Bank. The ROW aggregate combines data from China's major trading partners using trade-weighted averages, with detailed aggregation procedures provided in Appendix A.

A.1 Construction of Aggregate Variables

This appendix describes the methodology used to construct aggregate variables for the Rest of the World (RoW) and global indicators. The data are sourced from the IMF's Global Data Statistics and contain quarterly, seasonally adjusted series for 44 countries. All series are expressed in domestic currency units and include: real and nominal GDP, real consumption, real investment, policy interest rates, and exchange rates.

Real Consumption and Investment for the Rest of the World. To obtain real consumption and investment aggregates for the RoW, we apply the following procedure:

1. **Compute the Implicit Price Deflator.** For each country i and quarter t, we compute the implicit deflator as the ratio of nominal to real values:

$$\text{Deflator}_{i,t} = \frac{\text{Nominal}_{i,t}}{\text{Real}_{i,t}}$$

2. **Rebase Deflators to a Common Period.** All country-level deflators are rebased to the

first quarter of 2002:

$$\widetilde{\text{Deflator}}_{i,t} = \frac{\text{Deflator}_{i,t}}{\text{Deflator}_{i,2002Q1}}$$

3. **Recover Nominal Values Using the Rebased Deflator.** We convert the real series into nominal terms:

$$\widetilde{\text{Nominal}}_{i,t} = \text{Real}_{i,t} \cdot \widetilde{\text{Deflator}}_{i,t}$$

4. **Compute Time-Varying Nominal Weights.** For each period, we calculate each country's share in global nominal consumption (or investment):

$$w^C_{i,t} = \frac{\widetilde{\text{Nominal}}^C_{i,t}}{\sum_{j=1}^N \widetilde{\text{Nominal}}^C_{j,t}}, \quad w^I_{i,t} = \frac{\widetilde{\text{Nominal}}^I_{i,t}}{\sum_{j=1}^N \widetilde{\text{Nominal}}^I_{j,t}}$$

5. **Construct Global Deflators.** We compute weighted-average deflators using the shares from step 4:

$$\text{Global Deflator}^C_t = \sum_{i=1}^N w^C_{i,t} \cdot \widetilde{\text{Deflator}}^C_{i,t}, \quad \text{Global Deflator}^I_t = \sum_{i=1}^N w^I_{i,t} \cdot \widetilde{\text{Deflator}}^I_{i,t}$$

6. **Compute Real Series in Common Units.** Each country's nominal values are deflated using the global deflator to obtain real values on a common basis:

$$\widehat{\text{Real}}^C_{i,t} = \frac{\widetilde{\text{Nominal}}^C_{i,t}}{\text{Global Deflator}^C_t}, \quad \widehat{\text{Real}}^I_{i,t} = \frac{\widetilde{\text{Nominal}}^I_{i,t}}{\text{Global Deflator}^I_t}$$

The RoW real aggregates are then computed by summing over the set of countries not included in the focus group $\mathcal{F}$:

$$\text{RoW Real Consumption}_t = \sum_{i \notin \mathcal{F}} \widehat{\text{Real}}^C_{i,t}, \quad \text{RoW Real Investment}_t = \sum_{i \notin \mathcal{F}} \widehat{\text{Real}}^I_{i,t}$$

Global Policy Interest Rate. To construct a global policy interest rate, we use nominal GDP weights:

$$\text{Global Policy Rate}_t = \sum_{i=1}^{N} \left(\frac{\text{Nominal GDP}_{i,t}}{\sum_{j=1}^{N} \text{Nominal GDP}_{j,t}} \cdot \text{Policy Rate}_{i,t} \right)$$

This aggregation method ensures that countries with larger economic size have proportionally greater influence on the global average.

B Appendix for empirical results.

B.1 Tariffs in China

We construct three complementary measures of Chinese import tariffs to examine tariffs policy changes during our sample period. Figure B1 presents these measures from 2004 to 2020.

We use two different tariffs measures. For MFN and preferential tariffs schedules we use Feodora Teti's Global Tariff Database (pair-level beta1-2024-12) simple average bilateral tariff data (Teti, 2024). From this source, we calculate: (i) the simple average of all tariffs faced by China's trading partners (red dashed line), and (ii) the minimum tariff rate across all bilateral pairs in each year (green dash-dot line). These measures reflect China's scheduled tariff rates as reported to international organizations.

Our second measure of tariffs uses actual customs collections. We construct trade-weighted effective tariffs by dividing government customs revenue by total import values (blue long-dashed line). The customs revenue data comes from the Government Revenue Dataset maintained by UNU-WIDER, while import values come from IMF trade statistics.

Figure B1 shows that Chinese import tariffs remained remarkably stable throughout between the 2008-2010 period (shaded area). All three measures show minimal variation—scheduled tariffs drift downward gradually consistent with WTO commitments, while effective collection rates remain essentially flat around 2-3%. This stability contrasts sharply with the substantial VAT rebate adjustments we document in our main analysis. The absence of import tariff changes during the crisis period, when authorities actively adjusted export rebates, reinforces the underlying idea

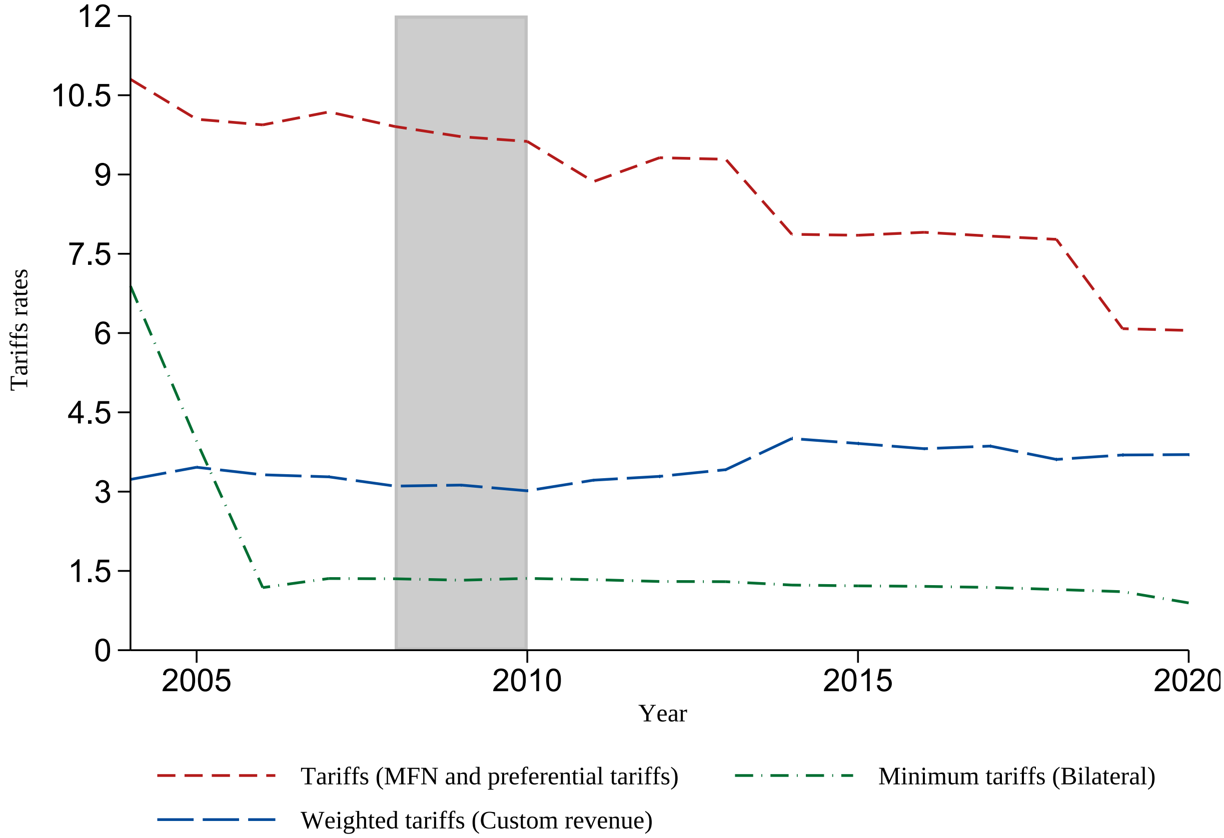


Figure B1: China Tariffs

Note: The first two variables (simple mean and minimum rates) are calculated from MFN and preferential tariff schedules in Teti (2024). The third measure is constructed by dividing government customs revenue by total import values.

of the quasi-fiscal devaluation we carry on the paper.

B.2 Bootstrap Procedure for Elasticity Estimates

Bootstrap Algorithm

We implement a cluster bootstrap procedure to obtain standard errors and confidence intervals for the export elasticity estimates reported in Section 3.2. The procedure accounts for the panel structure of our data and potential correlation within production units over time.

Step 1: Sample Selection

For each bootstrap iteration $b = 1, \ldots, B$ where $B = 1250$:

1.1. From the universe of N unique production units (codes), randomly sample with replacement. This subsampling approach ensures robustness to outlier clusters while maintaining sufficient sample size for identification.

1.2. For each selected unit i, include all time periods t in the bootstrap sample, preserving the within-unit temporal structure.

Step 2: Treatment Assignment

2.1. Calculate the treatment threshold for the new sample define as Δ_b.

2.2. Assign treatment status and controls status following the main specification in the text but using Δ_b.

Step 3: Estimation

3.1. Estimate the quantity response using the local projections difference-in-differences (LP-DiD) estimator:

$$\Delta^h \log(Q_{i,t+h}) = \gamma_t^h + \beta_Q^h \Delta E_{i,t} + \beta_{\mathbf{X}}^{\mathbf{h}} \mathbb{X}_{i,t} + \epsilon_{i,t} \text{ for h=-8,...,-2,0,...,24}$$

3.2. Estimate the rebate changes:

$$\Delta^h \text{rebate}_{i,t+h} = \gamma_t^h + \beta_Q^h \Delta E_{i,t} + \beta_{\mathbf{X}}^{\mathbf{h}} \mathbb{X}_{i,t} + \epsilon_{i,t}^h \text{ for h=-8,...,-2,0,...,24} \tag{8}$$

3.3. Calculate the elasticity at each horizon:

$$\varepsilon_h^{(b)} = \frac{\beta_Q^h}{\beta_R^h} \times 100 \tag{9}$$

Step 4: Aggregation

Store the vector of elasticities $\{\varepsilon_0^{(b)}, \varepsilon_1^{(b)}, \dots, \varepsilon_H^{(b)}\}$ from iteration b.

B.3 Inference

After completing all B bootstrap iterations:

Pivot Confidence Intervals

We construct 90% confidence intervals using the bootstrap pivot confidence interval method.

For each horizon h:

1. Obtain the $\alpha/2$ and $1-\alpha/2$ quantiles of the bootstrap distribution of each ε_h:

$$q^*_{h,\alpha/2} = Q_{\alpha/2}(\{\varepsilon_h^{(1)}, \dots, \varepsilon_h^{(B)}\}) \tag{10}$$

$$q^*_{h,1-\alpha/2} = Q_{1-\alpha/2}(\{\varepsilon_h^{(1)}, \dots, \varepsilon_h^{(B)}\}) \tag{11}$$

2. Construct the $(1-\alpha)$% pivot confidence interval:

$$CI^{\text{pivot}}_{1-\alpha}(\varepsilon_h) = \left[2\hat{\varepsilon}_h - q^*_{h,1-\alpha/2}, 2\hat{\varepsilon}_h - q^*_{h,\alpha/2}\right] \tag{12}$$

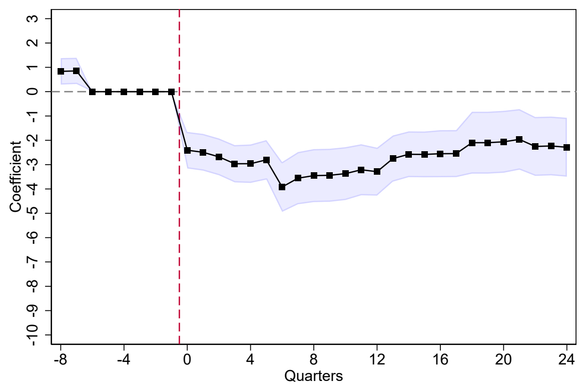
Note that the quantiles are reversed relative to standard intervals due to the pivoting operation.

B.4 Robustness for empirical results

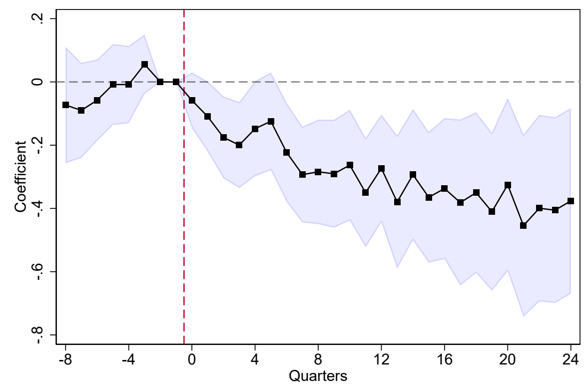
Here we detailed the different robustness we performed highlighted in Section 3.

B.5 Fixed VAT rebate after treatment.

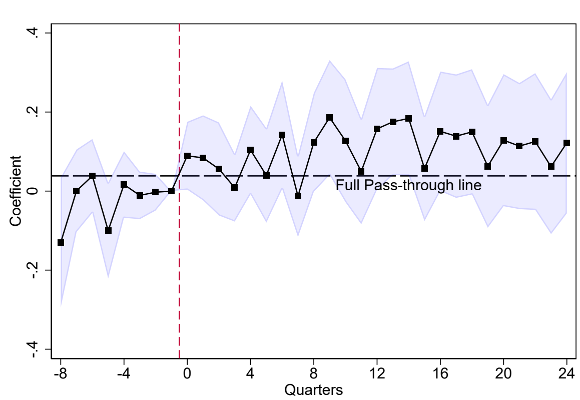
B.6 Smaller VAT rebate treatment.



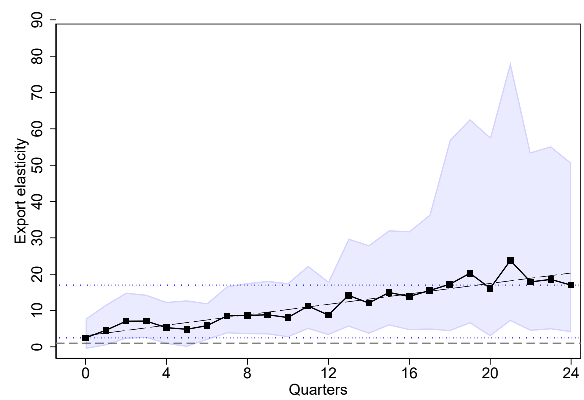
(a) VAT Rebate to Exporters



(b) Log Quantities



(c) Log Prices

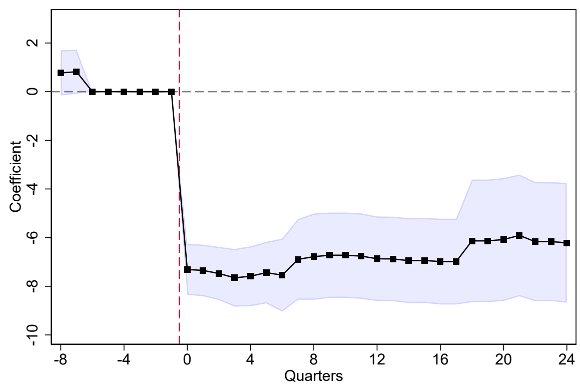


(d) Export elasticity to VAT rebate changes.

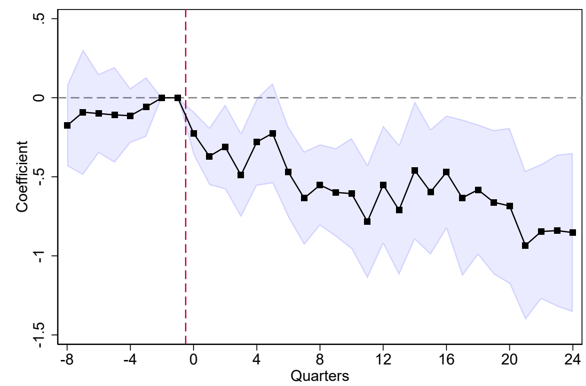
Robustness B.3: Smaller rebate treatment.

Note: Quarters are relative to the rebate change ($t = 0$). The base period for comparison is $t = -1$. Shaded area represents 95% confidence intervals clustered by product code. Prices are computed using unit values. Quarter-2 digit products fixed effects are included to control for seasonality. Treated products are those that had a rebate change below the 20th percentile.

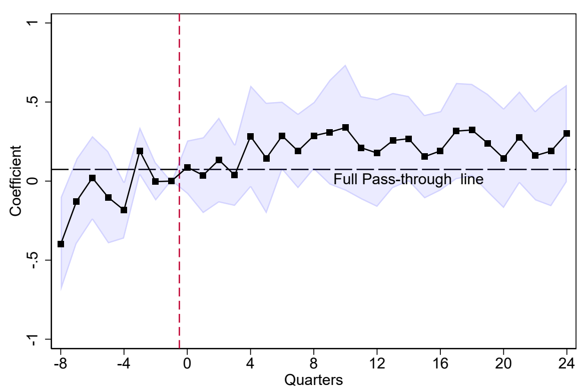
B.7 Sample based on exports reported in one common unit across products.



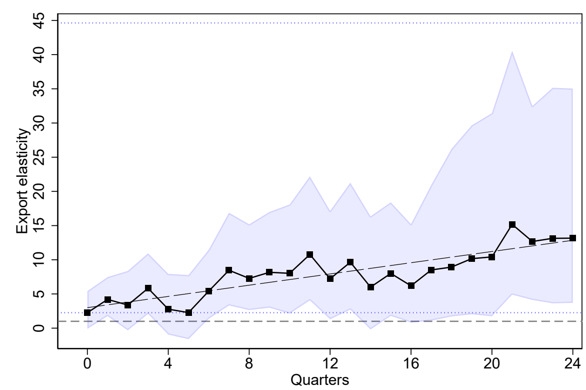
(a) VAT Rebate to Exporters



(b) Log Quantities



(c) Log Prices



(d) Export elasticity to VAT rebate changes.

Robustness B.4: Only products reported in Kilograms over sample.

Note: Quarters are relative to the rebate change ($t = 0$). The base period for comparison is $t = -1$. Shaded area represents 95% confidence intervals clustered by product code. Prices are computed using unit values. Quarter-2 digit products fixed effects are included to control for seasonality. Sample restricted to products whose units are reported in Kilograms over the whole sample.

C Full Model Appendix

C.1 Households

The household's problem is to maximize the present discounted value of flow utility subject to the budget constraint and the labor demand schedule from the labor packer, $L_t(h) = \left(\frac{W_t(h)}{W_t}\right)^{-\epsilon_w} L_t$, where ϵ_w is the elasticity of substitution across labor varieties.

$$1 = \beta_t \mathbb{E}_t \left(\frac{C_{t+1}}{C_t}\right)^{-\sigma} \frac{(1+i_t)}{\Pi_{t+1}}$$

$$1 + \chi B^*_{t+1} = \beta_t \mathbb{E}_t \left(\frac{C_{t+1}}{C_t}\right)^{-\sigma} \frac{\mathcal{Q}_{t+1}}{\mathcal{Q}_t} \frac{(1+i^*_t)}{\Pi^*_{t+1}}$$

$$1 = \frac{\chi_m e^{\mathcal{M}_t} C_t^\sigma}{m_t} + \beta_t \left(\frac{C_{t+1}}{C_t}\right)^{-\sigma} \frac{1}{\Pi_{t+1}}$$

where $\Pi_t \equiv \frac{P^c_t}{P^c_{t-1}}$ is the gross inflation rate, $\mathcal{Q}_t = \mathcal{E}_t P^{c*}_t / P^c_t$ the CPI real exchange rate, and $m_t \equiv M_t/P_t$ the real money balances.

The wage setting problem is the standard dynamic Calvo problem

$$\max_{W_t(i)} \mathbb{E}_t \sum_{s=0}^{\infty} \phi_w^s \beta_t \left[-\eta \frac{L_{t+s}(h)^{1+\nu}}{1+\nu} - \frac{W_{ch,t+s} L_{t+s}(h)}{P^c_{t+s}} \right]$$

subject to $L_t(h) = \left(\frac{W_t(h)}{W_t}\right)^{-\epsilon_w} L_t$.

The solution is given by,

$$\left(W_t^\#\right)^{1+\epsilon_w \nu} = \frac{\epsilon_w}{\epsilon_w - 1} \frac{\mathcal{H}_{1,t}}{\mathcal{H}_{2,t}}$$

$$\mathcal{H}_{1,t} = \eta L_t^{1+\nu} W_t^{\epsilon_w(1+\nu)} + \phi_w \beta_t \mathbb{E}_t \Pi_{t+1}^{\epsilon_w(1+\nu)} \mathcal{H}_{1,t+1}$$

$$\mathcal{H}_{2,t} = C_t^{-\sigma} w_t^{\epsilon_w} L_t + \phi_w \beta_t \mathbb{E}_t \Pi_{t+1}^{\epsilon_w - 1} \mathcal{H}_{2,t+1}$$

$$W_t^{1-\epsilon_w} = (1-\phi_w)\left(W_t^\#\right)^{1-\epsilon_w} + \phi_w \Pi_t^{\epsilon_w - 1} W_{t-1}^{1-\epsilon_w}$$

where $W_t^\#$ is the wage chosen by households that re-optimize.

C.2 Wholesalers

The wholesaler that re-optimizes faces the CES technology,

$$\hat{D}_t^*(j) = \left[(1-\omega)^{1/\gamma} \left(\hat{X}_{row,row,t}(j) \right)^{\frac{\gamma-1}{\gamma}} + \omega^{1/\gamma} \left(\hat{X}_{ch,row,t}(j) \right)^{\frac{\gamma-1}{\gamma}} \right]^{\frac{\gamma}{\gamma-1}}$$

The dynamic problem of the wholesaler can be written as:

$$\max_{\{\hat{X}_{row,row,t}(j), \hat{X}_{ch,row,t}(j)\}} \mathbb{E}_t \sum_{s=0}^{\infty} \delta^s \beta^s \Bigg\{ P_{row,t+s} \left[(1-\omega)^{1/\gamma} \left(\hat{X}_{row,row,t}(j) \right)^{\frac{\gamma-1}{\gamma}} + \omega^{1/\gamma} \left(\hat{X}_{ch,row,t}(j) \right)^{\frac{\gamma-1}{\gamma}} \right]^{\frac{\gamma}{\gamma-1}} \\ - P_{row,row,t+s} \hat{X}_{row,row,t}(j) - P_{ch,row,t+s} \hat{X}_{ch,row,t}(j) \Bigg\}$$

The first order condition with respect to $X_{row,row,t}(j)$ is

$$\mathbb{E}_t \sum_{s=0}^{\infty} \left[\delta^s \beta^s P_{row,t+s} \hat{D}_t^*(j) (1-\omega)^{1/\gamma} (X_{row,row,t}(j))^{-1/\gamma} - P_{row,row,t+s} \right] = 0$$

$$\hat{D}_t^*(j)(1-\omega)^{1/\gamma} (X_{row,row,t}(j))^{-1/\gamma} \mathbb{E}_t \sum_{s=0}^{\infty} \delta^s \beta^s P_{row,t+s} = \mathbb{E}_t \sum_{s=0}^{\infty} P_{row,row,t+s}$$

$$X_{row,row,t}(j) = (1-\omega) \hat{D}_t^*(j) \frac{\mathbb{E}_t \sum_{s=0}^{\infty} \delta^s \beta^s P_{row,t+s}}{\mathbb{E}_t \sum_{s=0}^{\infty} P_{row,row,t+s}}$$

$$\hat{X}_{row,row,t}(j) = (1-\omega)\, \hat{D}_t^*(j) \left(\frac{\mathcal{P}_{row,row,t}}{\mathcal{P}_{row,t}} \right)^{-\gamma}$$

And similarly for the choice of $X_{ch,row}(j)$ we obtain

$$\hat{X}_{ch,row,t}(j) = \omega \hat{D}_t^*(j) \left(\frac{\mathcal{P}_{ch,row,t}}{\mathcal{P}_{row,t}} \right)^{-\gamma}$$

The aggregate bundles of inputs of optimizing firms are given by,

$$\int_0^{1-\delta} \hat{X}_{row,row,t}(j) dj = (1-\delta) \int_0^1 \hat{X}_{row,row,t}(j) dj$$

and

$$\int_0^{1-\delta} \hat{X}_{ch,row,t}(j)dj = (1-\delta)\int_0^1 \hat{X}_{ch,row,t}(j)dj$$

and the aggregate quantities sold by these firms are

$$\int_0^{1-\delta} \hat{D}_t^*(j)dj = (1-\delta)\int_0^1 \hat{D}_t^*(j)dj$$

Since all optimizing firms have the same state variables then they all pick the same bundles, hence

$$\hat{X}_{row,row,t} \equiv \hat{X}_{row,row,t}(j) = \int_0^1 \hat{X}_{row,row,t}(j)dj$$

$$\hat{X}_{ch,row,t} \equiv \hat{X}_{ch,row,t}(j) = \int_0^1 \hat{X}_{ch,row,t}(j)dj$$

$$\hat{D}_t^* \equiv \hat{D}_t^*(j) = \int_0^1 \hat{D}_t^*(j)dj$$

Then it follows that the aggregate bundles are

$$\int_0^{1-\delta} \hat{X}_{row,row,t}(j)dj = (1-\omega)\left(\int_0^{1-\delta} \hat{D}_t^*(j)dj\right)\left(\frac{\mathcal{P}_{row,row,t}}{\mathcal{P}_{row,t}}\right)^{-\gamma}$$

$$(1-\delta)\int_0^1 \hat{X}_{row,row,t}(j)dj = (1-\omega)\left((1-\delta)\int_0^1 \hat{D}_t^*(j)dj\right)\left(\frac{\mathcal{P}_{row,row,t}}{\mathcal{P}_{row,t}}\right)^{-\gamma}$$

$$(1-\delta)\hat{X}_{row,row,t} = (1-\omega)\,(1-\delta)\hat{D}_t^*\left(\frac{\mathcal{P}_{row,row,t}}{\mathcal{P}_{row,t}}\right)^{-\gamma}$$

$$\Rightarrow \hat{X}_{row,row,t} = (1-\omega)\,\hat{D}_t^*\left(\frac{\mathcal{P}_{row,row,t}}{\mathcal{P}_{row,t}}\right)^{-\gamma}$$

and similarly for the demand of imported intermediates we have

$$\hat{X}_{ch,row,t} = \omega\hat{D}_t^*\left(\frac{\mathcal{P}_{ch,row,t}}{\mathcal{P}_{row,t}}\right)^{-\gamma}$$

We can then derive the law of motion for inputs and quantities produced,

$$X_{row,row,t} = \int_0^1 X_{row,row,t}(j)dj = \int_0^{1-\delta} \hat{X}_{row,row,t}(j)dj + \int_{1-\delta}^1 X_{row,row,t}(j)dj$$

$$X_{row,row,t} = (1-\delta)\int_0^1 \hat{X}_{row,row,t}(j)dj + \delta\int_0^1 X_{row,row,t}(j)dj$$

$$X_{row,row,t} = (1-\delta)\hat{X}_{row,row,t} + \delta\int_0^1 X_{row,row,t}(j)dj$$

since non-optimizing firms will pick the same inputs as in $t-1$, then

$$X_{row,row,t} = (1-\delta)\hat{X}_{row,row,t} + \delta\int_0^1 X_{row,row,t-1}(j)dj$$

then,

$$X_{row,row,t} = (1-\delta)\hat{X}_{row,row,t} + \delta X_{row,row,t-1}$$

Similarly,

$$X_{ch,row,t} = (1-\delta)\hat{X}_{ch,row,t} + \delta X_{ch,row,t-1}$$

and

$$D_t^* = (1-\delta)\hat{D}_t^* + \delta D_{t-1}^*$$

Finally, the ideal price index can be solved from the CES of the optimizing wholesalers under their optimal demands,

$$\int_0^1 \hat{D}_t^*(j)dj = \int_0^1 \left\{(1-\omega)^{1/\gamma}\left[(1-\omega)\,\hat{D}_t^*(j)\left(\frac{\mathcal{P}_{row,row,t}}{\mathcal{P}_{row,t}}\right)^{-\gamma}\right]^{\frac{\gamma-1}{\gamma}} + \omega^{1/\gamma}\left[(1-\omega)\,\hat{D}_t^*(j)\left(\frac{\mathcal{P}_{ch,row,t}}{\mathcal{P}_{row,t}}\right)^{-\gamma}\right]^{\frac{\gamma-1}{\gamma}}\right\}^{\frac{\gamma}{\gamma-1}} dj$$

which collapses to

$$\mathcal{P}_{row,t} = \left[(1-\omega)\,\mathcal{P}_{row,row,t}^{1-\gamma} + \omega\mathcal{P}_{ch,row,t}^{1-\gamma}\right]^{\frac{1}{1-\gamma}}$$

Note that when $\delta \to 0$ the model collapses to the standard Armington model with the usual ideal price index.

C.3 Retailers

An optimizing retailer chooses its price to maximize real profits, deflated by the final consumption price P_t^c:

$$\max_{P_t(i)} \mathbb{E}_t \sum_{s=0}^{\infty} \phi_p^s \Lambda_{t,t+s}\left[\frac{P_t(i)Y_{t+s}(i)}{P_{t+s}^c} - \frac{P_{ch,t+s}D_{t+s}(i)}{P_{t+s}^c}\right],$$

subject to the demand function above.

The solution is characterized by an optimal reset price, $P_t^{\#}$, given by

$$P_t^{\#} = \frac{\varepsilon_p - 1}{\varepsilon_p} \frac{\mathcal{J}_{1,t}}{\mathcal{J}_{2,t}}$$

where,

$$\mathcal{J}_{1,t} = P_{ch,t} Y_t + \phi_p \mathbb{E}_t \Lambda_{t,t+1} \mathcal{J}_{1,t+1}$$

$$\mathcal{J}_{2,t} = Y_t + \phi_p \mathbb{E}_t \Lambda_{t,t+1} \Pi_t^{\varepsilon_p - 1} \mathcal{J}_{2,t+1}$$

and the evolution of aggregate inflation dynamics, price dispersion and output are

$$1 = (1 - \phi_p) \left(P_t^{\#}\right)^{1-\varepsilon_p} + \phi_p \Pi_t^{\varepsilon_p - 1}$$

$$\varrho_t^p = (1 - \phi_p) \left(P_t^{\#}\right)^{-\varepsilon_p} + \phi_p \Pi_t^{\varepsilon_p} \varrho_{t-1}^p$$

$$D_t = Y_t \varrho_t^p$$

C.4 Proof of Proposition 1

Proof. In partial equilibrium aggregate prices and quantities are held constant at their steady state level. Then,

1. Log-linearize the aggregate demand for Chinese exports and its price,

$$\tilde{x}_{ch,row,t} = (1 - \delta) \left[-\gamma \left(\tilde{\mathcal{P}}_{ch,row,t} - \tilde{\mathcal{P}}_{row,t} \right) + \tilde{D}_t^* \right] + \delta \tilde{x}_{ch,row,t-1}$$

$$\tilde{p}_{ch,row,t} = \tilde{MC}_{ch,t} - \tilde{\zeta}_t$$

where

$$\tilde{\mathcal{P}}_{ch,row,t} = (1 - \delta\beta) \tilde{p}_{ch,row,t} + \delta\beta \mathbb{E}_t \tilde{\mathcal{P}}_{ch,row,t+1}$$

$$\tilde{\mathcal{P}}_{row,t} = (1 - \delta\beta) \tilde{p}_{row,t} + \delta\beta \mathbb{E}_t \tilde{\mathcal{P}}_{row,t+1}$$

$$\tilde{D}_t^* = \tilde{C}_t^* + \tilde{\varrho}_t^{p*}$$

Note that we can express the recursion of prices as,

$$\tilde{\mathcal{P}}_{ch,row,t} = (1 - \delta\beta)\mathbb{E}_t \sum_{j=0}^{\infty} \tilde{p}_{ch,row,t+j} \, (\delta\beta)^j$$

$$\tilde{\mathcal{P}}_{row,t} = (1 - \delta\beta)\mathbb{E}_t \sum_{j=0}^{\infty} \tilde{p}_{row,t+j} \, (\delta\beta)^j$$

and that under partial equilibrium $\tilde{C}_t^* = 0$ and $\tilde{\varrho}_t^{p*} = 0$ since $\tilde{P}_t^{\#*} = 0$ and $\tilde{\pi}_t^* = 0$, that is aggregate consumption and aggregate price dispersion in the ROW are constant.

2. Differentiate the time-zero exports to a *permanent* increase in the VATRX ($\tilde{\zeta}$)

$$\frac{\partial \tilde{x}_{ch,row,0}}{\partial \tilde{\zeta}} = -(1-\delta)\gamma \quad \times \quad \frac{\partial \tilde{\mathcal{P}}_{ch,row,0}}{\partial \tilde{\zeta}} \quad + \quad \delta \quad \times \quad \frac{\partial \tilde{x}_{ch,row,-1}}{\partial \tilde{\zeta}}$$

$$\frac{\partial \tilde{x}_{ch,row,0}}{\partial \tilde{\zeta}} = -(1-\delta)\gamma \quad \times \quad \left[(1-\delta\beta) \sum_{j=0}^{\infty} \frac{\partial \tilde{p}_{ch,row,j}}{\partial \tilde{\zeta}} \quad (\delta\beta)^j \right] \quad + \quad \delta \times \frac{\partial \tilde{x}_{ch,row,-1}}{\partial \tilde{\zeta}}$$

Since $\frac{\partial \tilde{p}_{ch,row,h}}{\partial \tilde{\zeta}} = -1 \; \forall h$ (i.e permanent change in rebate translates into permanent change in export price) and $\frac{\tilde{x}_{ch,row,-1}}{\partial \tilde{\zeta}} = 0$ then,

$$\frac{\partial \tilde{x}_{ch,row,0}}{\partial \tilde{\zeta}} = -(1-\delta)\gamma$$

□

D Estimation Appendix

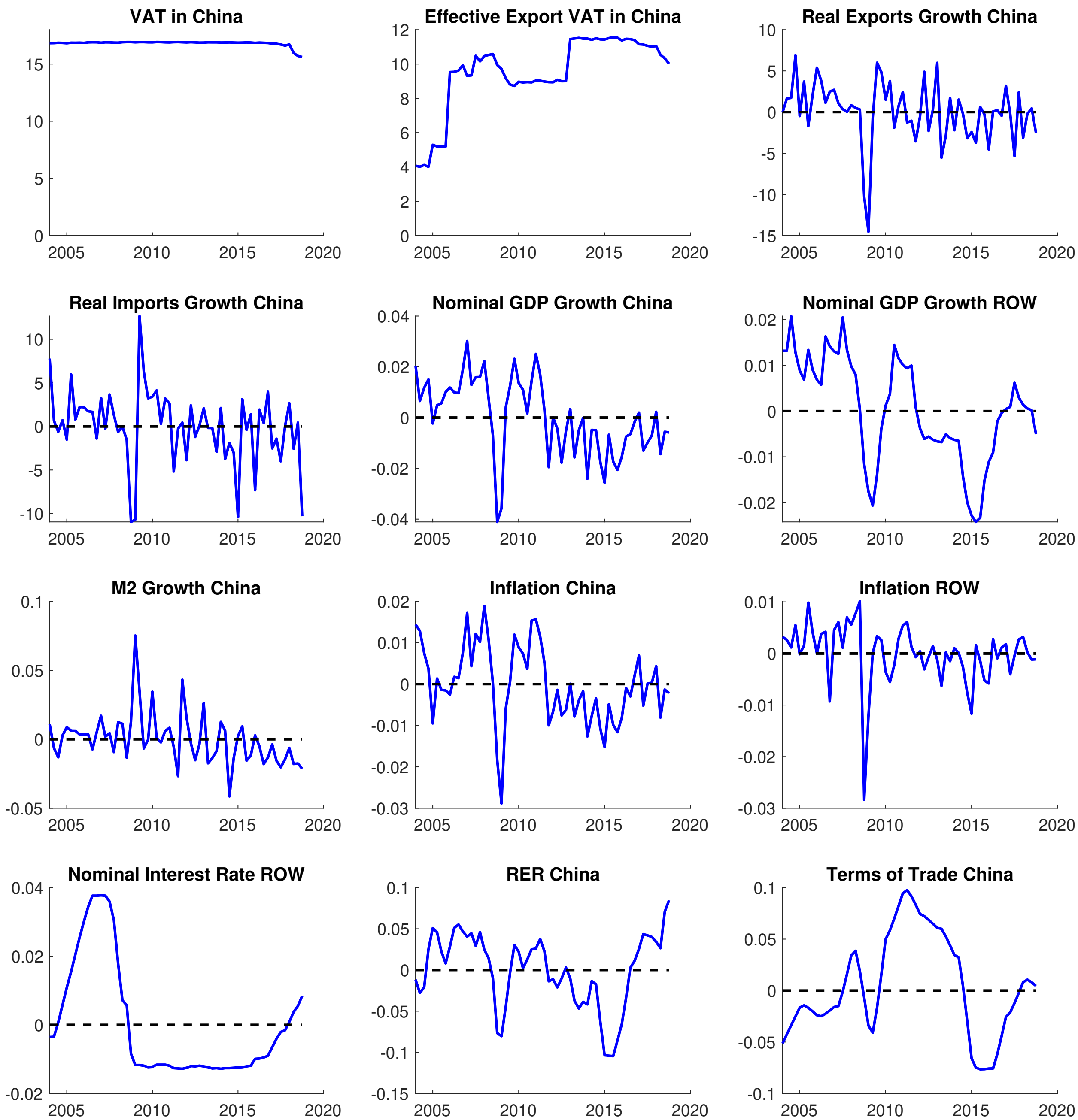


Figure D1: Data for Bayesian Estimation

D.1 Convergence and Identification Appendix

We assess convergence and identification to ensure the validity of the Bayesian estimation. The diagnostics point to two main conclusions: (i) the MCMC sampler converges satisfactorily, and (ii) the observables used in the estimation provide meaningful information for identifying the estimated parameters.

To evaluate convergence, we apply the Geweke (1992) diagnostic, which compares the mean of an early segment of the Markov chain to that of a later segment. Under convergence to the stationary posterior distribution, these two means should be statistically indistinguishable, yielding a test statistic close to zero. As shown in Table D1, the parameters exhibit Geweke p-values above conventional significance thresholds, implying that the null of equality of means cannot be rejected. This provides strong evidence that the chains have reached their stationary distribution and that the posterior moments reported in the paper are based on a well-behaved and well-mixed MCMC sampler.

We also verify that the model's observables deliver sufficient statistical information for identifying the structural parameters. Using the asymptotic information matrix we assess how sensitive the likelihood is to small perturbations in each parameter. All parameters display sizable curvature at the estimated values, indicating strong local identification, shown in Figure D2. We also applied complementary identification diagnostics to ensure robustness. Following Komunjer and Ng (2011), we checked rank conditions of the spectral density; Qu and Tkachenko (2012) used frequency-domain methods to trace non-identification curves; and Iskrev (2010) examined local sensitivity of autocovariances to parameters. All tests confirm that all parameters are well identified by our observables.

Together, these results confirm that the estimated parameters are both well-identified and supported by the data.

Table D1: Geweke (1992) Convergence Diagnostics

Parameter	Post. Mean	Post. Std	p-val 4% Taper	p-val 8% Taper	p-val 15% Taper
ρ_z	0.926	0.018	0.231	0.263	0.269
ρ_{z^*}	0.865	0.031	0.685	0.693	0.667
σ_z	0.094	0.010	0.823	0.820	0.818
σ_{z^*}	0.097	0.014	0.971	0.970	0.967
$\rho_{\xi_{ch,row}}$	0.966	0.022	0.793	0.810	0.824
$\sigma_{\xi_{ch,row}}$	0.088	0.010	0.629	0.620	0.616
$\rho_{\xi_{row,ch}}$	0.937	0.023	0.352	0.351	0.371
$\sigma_{\xi_{row,ch}}$	0.081	0.009	0.053	0.036	0.011
ρ_ψ	0.562	0.041	0.855	0.856	0.857
σ_ψ	0.046	0.005	0.288	0.265	0.242
ρ_τ	0.113	0.052	0.812	0.781	0.744
σ_τ	0.002	0.000	0.282	0.292	0.235
ρ_ζ	0.976	0.016	0.710	0.729	0.743
σ_ζ	0.007	0.001	0.793	0.799	0.771
ρ_{i^*}	0.819	0.017	0.814	0.806	0.790
σ_{i^*}	0.004	0.000	0.539	0.563	0.579
ρ_m	0.433	0.061	0.475	0.452	0.402
$\phi_{m,\pi}$	0.568	0.069	0.876	0.884	0.892
$\phi_{m,e}$	0.482	0.058	0.713	0.680	0.656
$\phi_{m,gdp}$	0.060	0.043	0.654	0.659	0.660
σ_m	0.020	0.002	0.572	0.573	0.516
$\rho_{\mathcal{M}}$	0.986	0.002	0.038	0.038	0.054
$\sigma_{\mathcal{M}}$	0.469	0.050	0.046	0.060	0.076
ρ_β	0.797	0.033	0.989	0.988	0.987
σ_β	0.033	0.004	0.275	0.312	0.304
ρ_{β^*}	0.671	0.030	0.499	0.472	0.439
σ_{β^*}	0.028	0.003	0.129	0.099	0.077

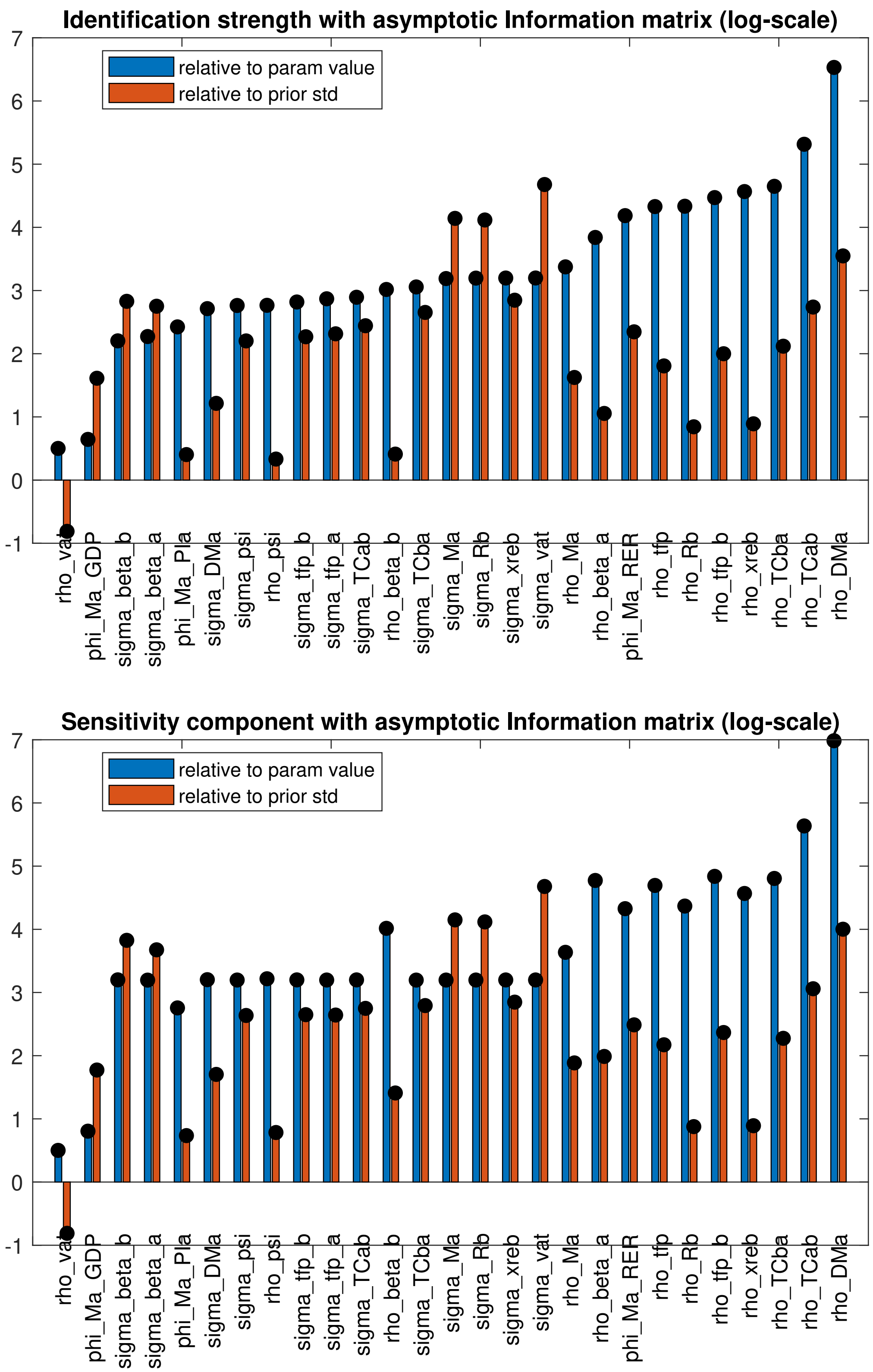


Figure D2: Identification Strength

E Quantitative Results Appendix

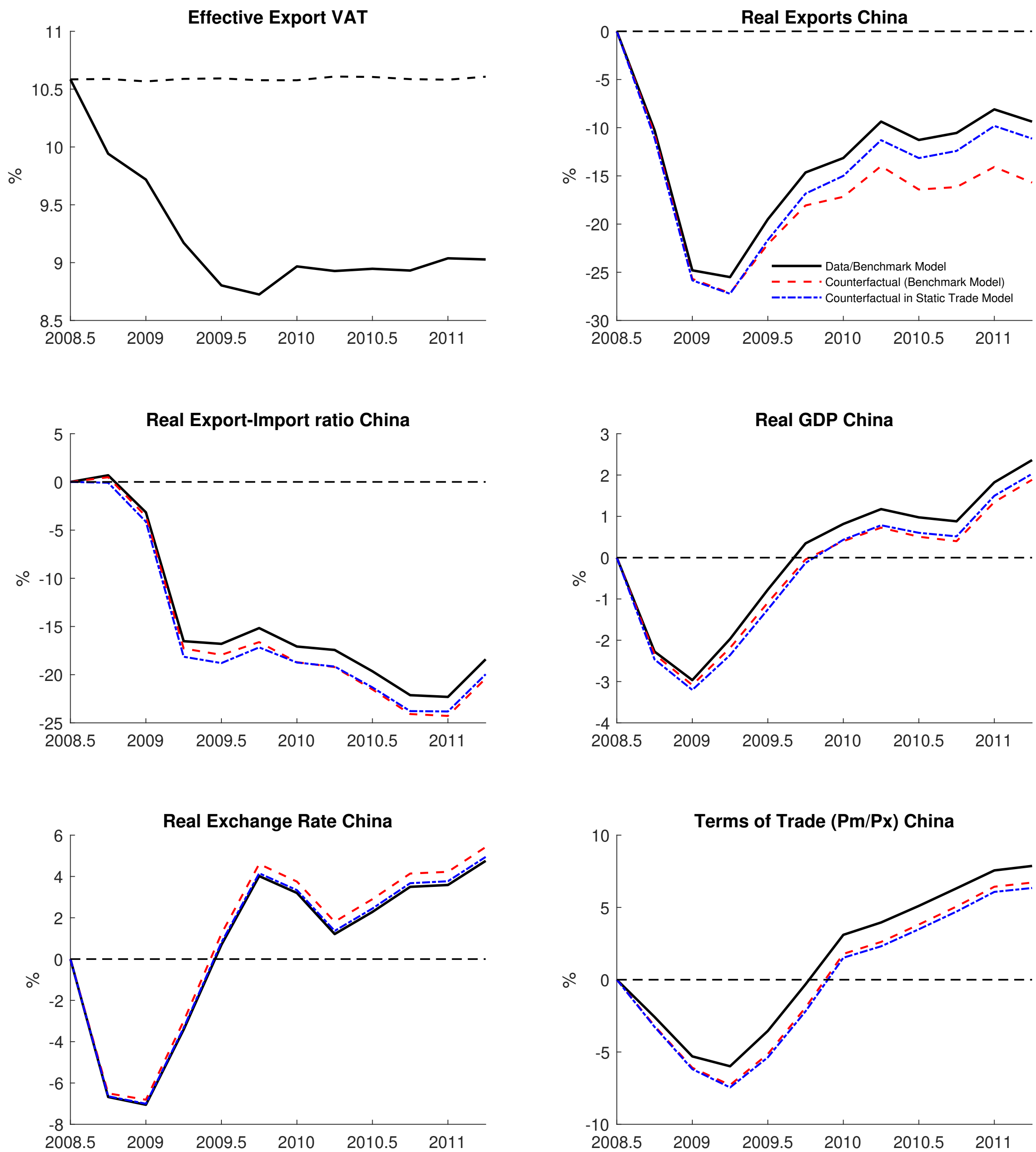


Figure E1: Global Financial Crisis: Counterfactual with Static Short-Run Trade Elasticity

Table E1: Semi-Elasticities of Key Variables under Benchmark vs. Static Trade Elasticity Scenarios

Variable	Short Run		Medium Run	
	Benchmark	Static Trade	Benchmark	Static Trade
Export	1.04	1.22	4.00	1.11
Export-Import Ratio	0.48	1.15	1.25	0.98
GDP	0.14	0.28	0.30	0.21
Terms of Trade	0.92	1.04	0.97	0.92
Real Exchange Rate	0.28	0.06	0.42	0.12