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Dynamic and spillover effects of USA import refusals on China's agricultural trade: Evidence from monthly data

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ELECTRONIC SUPPLEMENTARY MATERIALS (ESM):

- Supplementary materials (S1–2)
- Supplementary tables (Tables S1–2)
- Supplementary figures (Figures S1–5)

Supplementary materials

Supplementary material S1. Data Description

Between 2002 and 2016, the FDA import refusals on agricultural exports from China exhibited an overall upward trend, making China one of the top target countries of FDA in 2016. As Figure S2 reveals, the rise of refusal charges received by Chinese agricultural exporters was accompanied by substantial growth of agricultural exports from China to the United States. It implies that export surges may exhibit positive feedback to the decision of import refusals. An empirical model that takes both variables as endogenous is thus necessary. While sectors with greater exports tend to receive more refusals in general, the sectoral distribution of refusal charges differs from that of export value according to Figure S3. It suggests that the relationship between refusals and exports may vary across sectors, and the consideration of inter-sectoral heterogeneities is needed.

As a preliminary look into the effect of refusal charges on agricultural exports, Figure S4 considers an “average good” for all agricultural products of China that have been at least refused once during the sample period and tracks the monthly export value of this good around the date that it first received an import refusal. Specifically, we normalise the date of receiving the first import refusal to zero for each product, and consider the window between one year prior to this date and five years afterwards. Along these normalized dates, we calculate the average number of refusal charges and export values in logs by month across products. The figure shows a substantial one-month decline in exports when the first import refusal took place, implying a possible instantaneous export-reducing effect of refusal charges. Exports then rapidly picked up, and the number of refusal charges decreased, suggesting that refusal charges mostly served as a one-time shock with short-run impacts.

Supplementary material S2. Details of Methods

To be specific, we first de-trend the series of refusal charges, export quantities and unit prices, seeing that these variables are trend-stationary according to panel unit root test results presented in Supplemen-

tary Table S2. That is, we are concerned with de-trended variables $z_{it} = y_{it} - \bar{y}_i$, where $\bar{y}_i$ is the inter-temporal mean of vector y_{it} . The reduced form of Equation (1) is then estimated respectively for the 12 sectors with z_{it} and the VAR system of sectoral means with $\bar{z}_t = (\overline{charge_{it}}, \overline{lnuv_{it}}, \overline{lnq_{it}})$. The lengths of lag effects p_i and $\bar{p}$ are determined in each estimation using a likelihood ratio test with four information criteria.¹ Using estimated residuals of the reduced-form model as well as coefficient matrices of its associated vector moving-average representation, we next back out structural shocks and coefficient matrices of the structural-form model based on identifying restrictions. The implied structural shocks under these restrictions enables us to compute the loading matrix of common components in each sector, i.e. Λ_i , by estimating an OLS regression of ε_{it} on $\bar{\varepsilon}$. With this loading matrix, we can then decompose impulse response functions of each sector to composite shocks into impulse responses to common and idiosyncratic components.

It shall be noted that the decomposition above would yield a covariate matrix of idiosyncratic shocks with non-unity diagonal elements, though the covariate matrix of common shocks can be typically normalised to identity. That is, the estimated impulse responses should reflect impacts of unity sized common shocks and non-unity sized idiosyncratic shocks, making the two types of responses not directly comparable. We thus follow Pedroni (2013) by renormalising the covariance matrix of idiosyncratic shocks to identity as well.

Supplementary tables

Table S1. Reputation effects estimation

Variable	$Charge_{HS4,t}$
$Charge_{HS4,t-1}$	0.186*** (0.000)
$Charge_{HS4,t-2}$	0.160*** (0.000)
$Charge_{HS4,t-3}$	0.098*** (0.000)
$Charge_{otherHS4,t-1}$	0.007** (0.034)
$Charge_{othersector,t-1}$	0.00005 (0.897)
$lnval_all_{HS4,t-1}$	0.068*** (0.001)
R^2	0.1051
Observations	14 382

***, ** and * denotes statistical significance at the 1, 5 and 10% level, respectively; $Charge_{HS4,t}$ – dependent variable, the number of import refusal charges affecting a particular HS four-digit product in period t ; $Charge_{otherHS4,t-1}$ – lagged number of refusal charges affecting other HS four-digit products in the same sector, which reflects the sector reputation of refusals; $Charge_{othersector,t-1}$ – lagged number of refusal charges affecting other sectors in our sample; $lnval_all_{HS4,t-1}$ – lagged Chinese export value to the United States

Source: authors' elaboration

¹The four information criteria include the final prediction error (FPE), Akaike's information criterion (AIC), Schwarz's Bayesian information criterion (SBIC), and the Hannan and Quinn information criterion (HQIC).

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Table S2. Panel stationarity test results

Variable	Im-Pesaran-Shin test (IPS)			ADF-Fisher	PP-Fisher
	AIC	BIC	HQIC		
$lnval_{USA}$	–20.22 (0.0000)	–23.66 (0.0000)	–21.44 (0.0000)	29.24 (0.0000)	72.25 (0.0000)
lnq_{USA}	–17.80 (0.0000)	–24.56 (0.0000)	–23.65 (0.0000)	26.97 (0.0000)	68.61 (0.0000)
$lnval_{JPN}$	–22.09 (0.0000)	–24.81 (0.0000)	–22.30 (0.0000)	43.87 (0.0000)	71.10 (0.0000)
$lnval_{KOR}$	–14.81 (0.0000)	–19.93 (0.0000)	–17.01 (0.0000)	34.91 (0.0000)	60.35 (0.0000)
$lnval_{EU}$	–16.99 (0.0000)	–21.59 (0.0000)	–19.90 (0.0000)	28.47 (0.0000)	61.43 (0.0000)
$lnval_{ASEAN}$	–14.76 (0.0000)	–21.38 (0.0000)	–15.60 (0.0000)	28.66 (0.0000)	69.84 (0.0000)
$lnval_{HKG}$	–15.16 (0.0000)	–19.80 (0.0000)	–18.48 (0.0000)	36.25 (0.0000)	62.00 (0.0000)
$Price_{USA}$	–9.42 (0.0000)	–14.60 (0.0000)	–12.78 (0.0000)	19.54 (0.0000)	60.23 (0.0000)
$Charge_{USA}$	–24.26 (0.0000)	–30.98 (0.0000)	–24.83 (0.0000)	62.36 (0.0000)	106.99 (0.0000)
$Adulteration_{USA}$	–21.54 (0.0000)	–29.85 (0.0000)	–28.57 (0.0000)	53.98 (0.0000)	104.42 (0.0000)
$Non-Adulteration_{USA}$	–29.30 (0.0000)	–38.08 (0.0000)	–33.36 (0.0000)	80.99 (0.0000)	113.30 (0.0000)

$lnval_s$ – China's export value to each region; lnq_s – China's export quantity to each region; probabilities for IPS tests are computed using W-t-bar statistics; AIC – Akaike information criterion; BIC – Bayesian information criterion; HQIC – Hannan-Quinn information criterion; probabilities for Fisher tests are computed using a modified inverse chi-square distribution; ADF-Fisher – augmented Dickey-Fuller Fisher test; PP-Fisher – Phillips-Perron Fisher test

Source: authors' elaboration

Supplementary figures

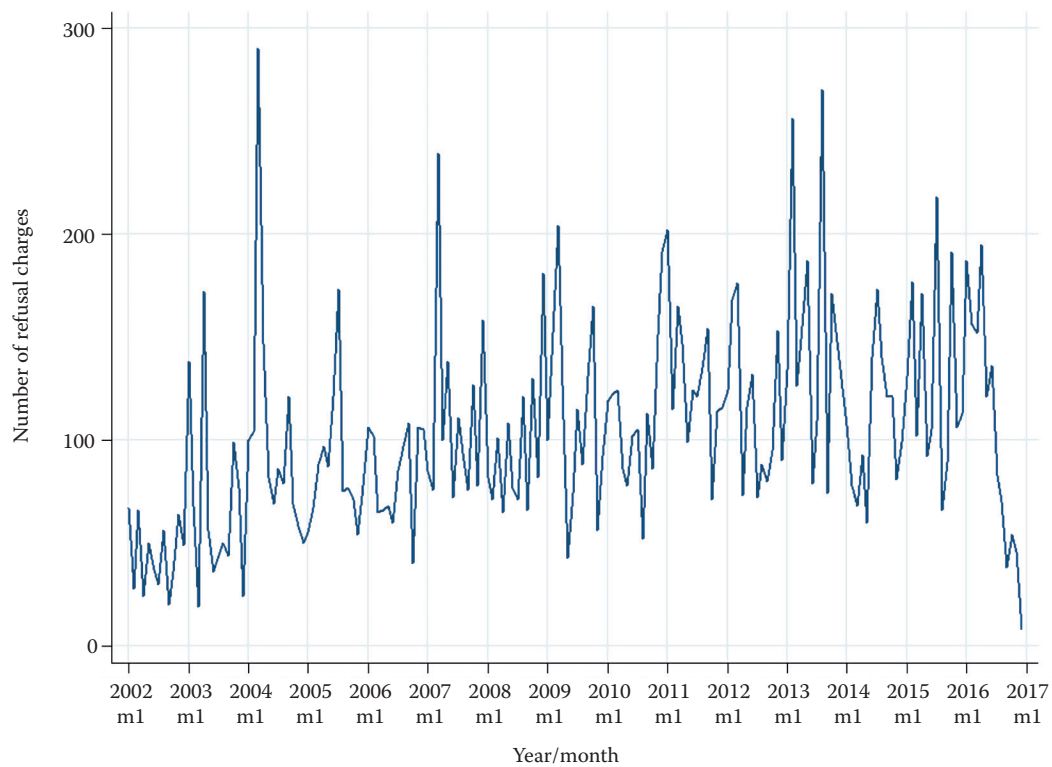


Figure S1. Monthly United States Food and Drug Administration (FDA) refusal charges on China's agricultural exports
Source: authors' elaboration

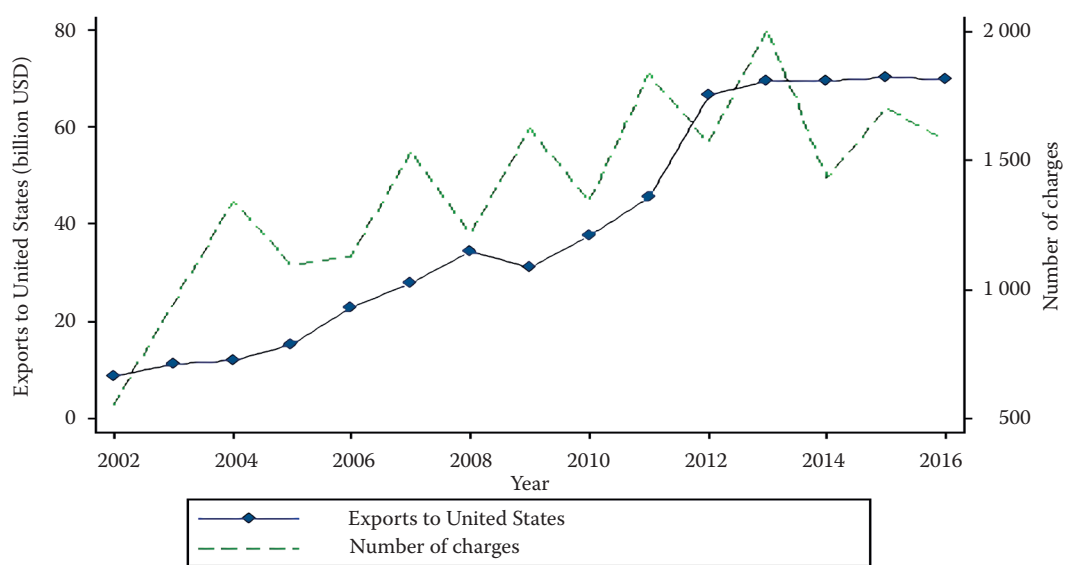


Figure S2. China's agricultural exports to the United States and the number of refusal charges
Source: authors' elaboration

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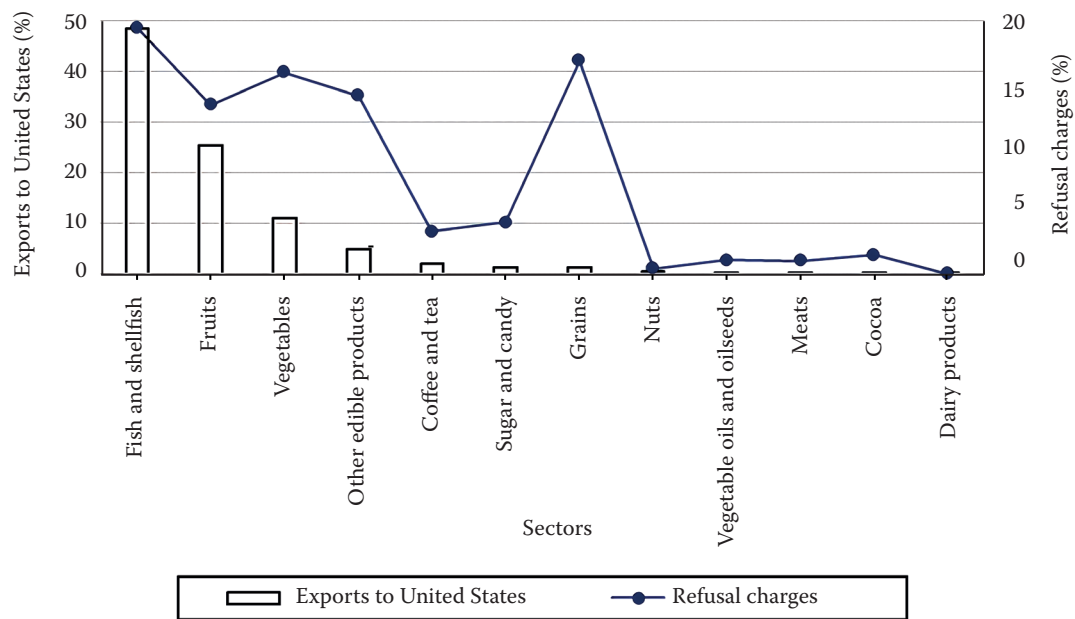


Figure S3. Sectoral distribution of China's agricultural exports and refusal charges on them

Source: authors' elaboration

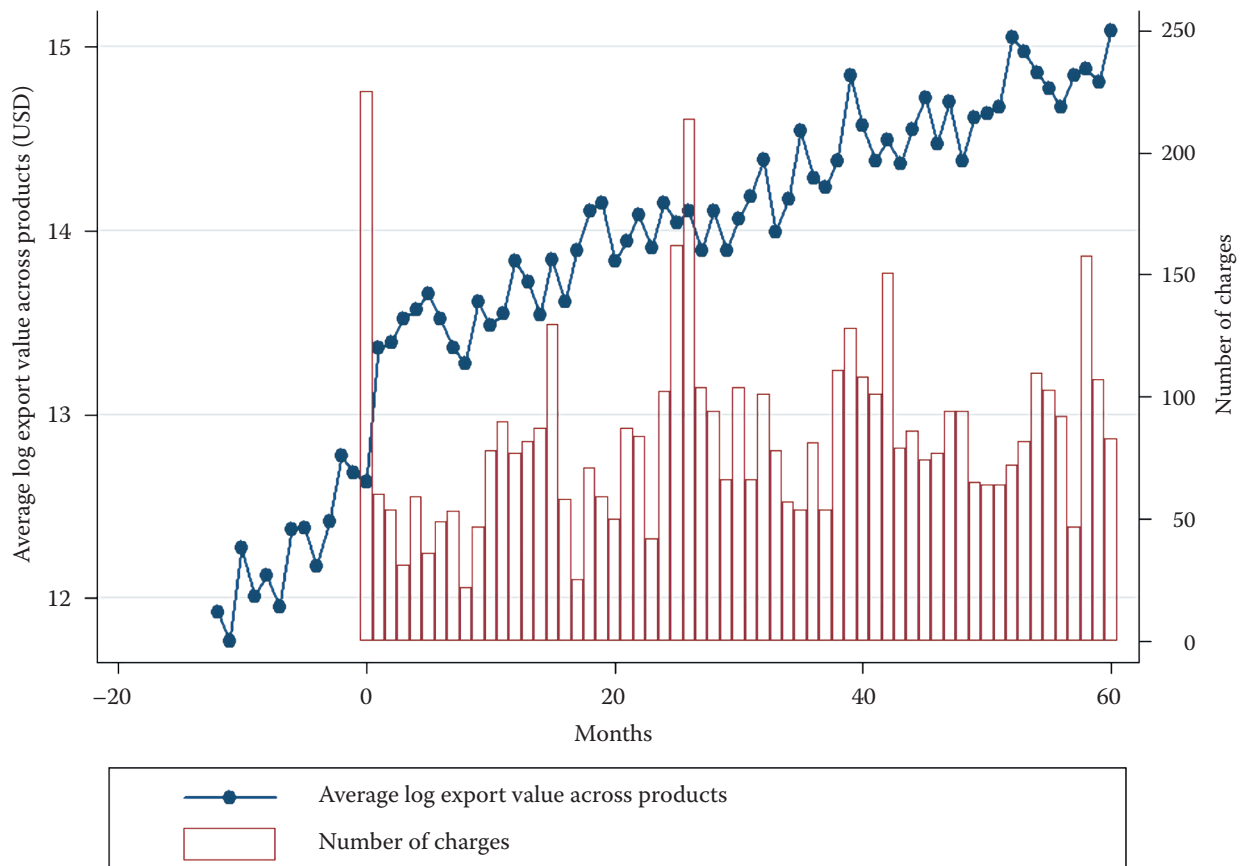


Figure S4. Evolution of China's agricultural exports and FDA refusal charges along normalized dates

Source: authors' elaboration

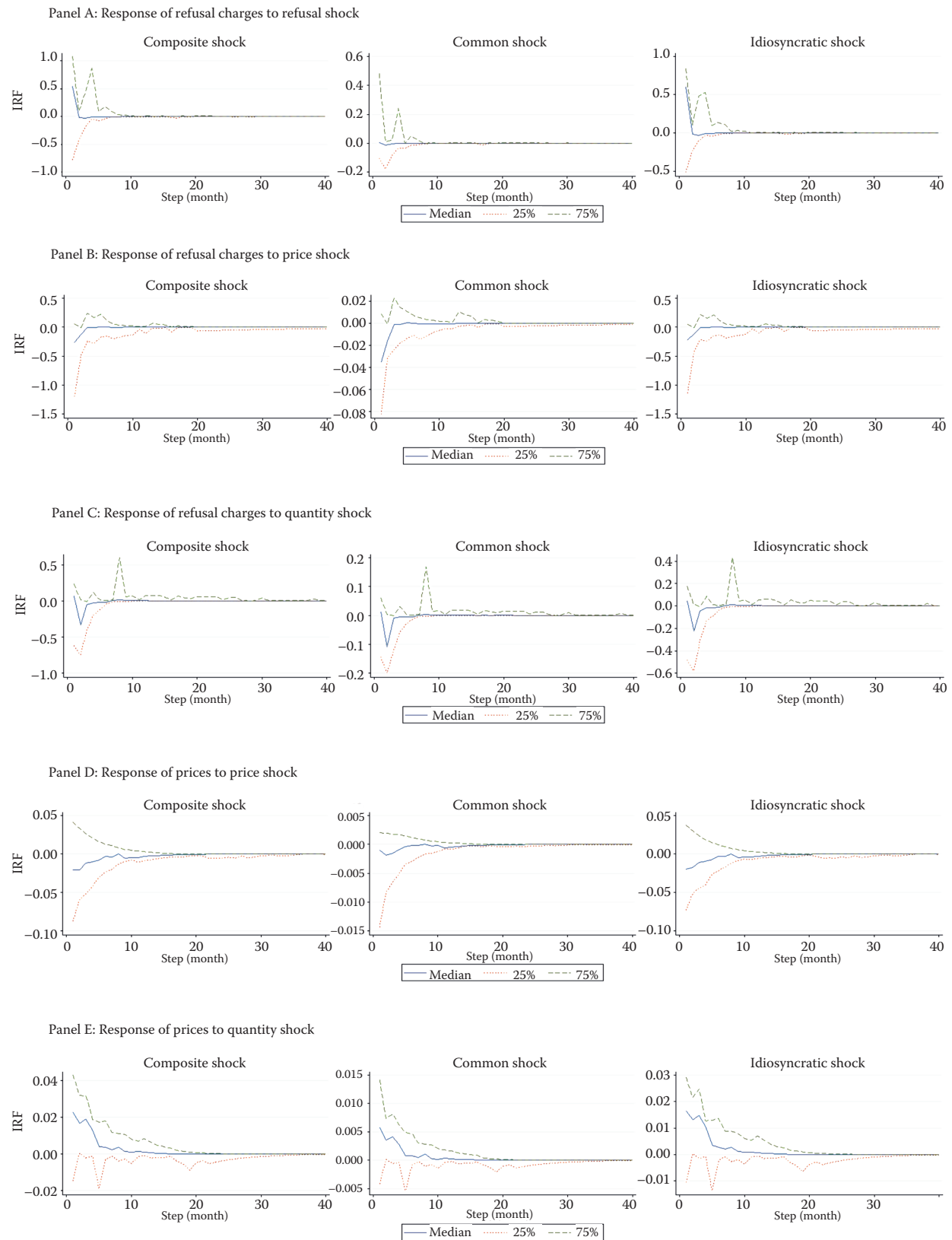


Figure S5. Other impulse response functions (IRFs) estimated from the benchmark model

Source: authors' elaboration

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Continuation Figure S5

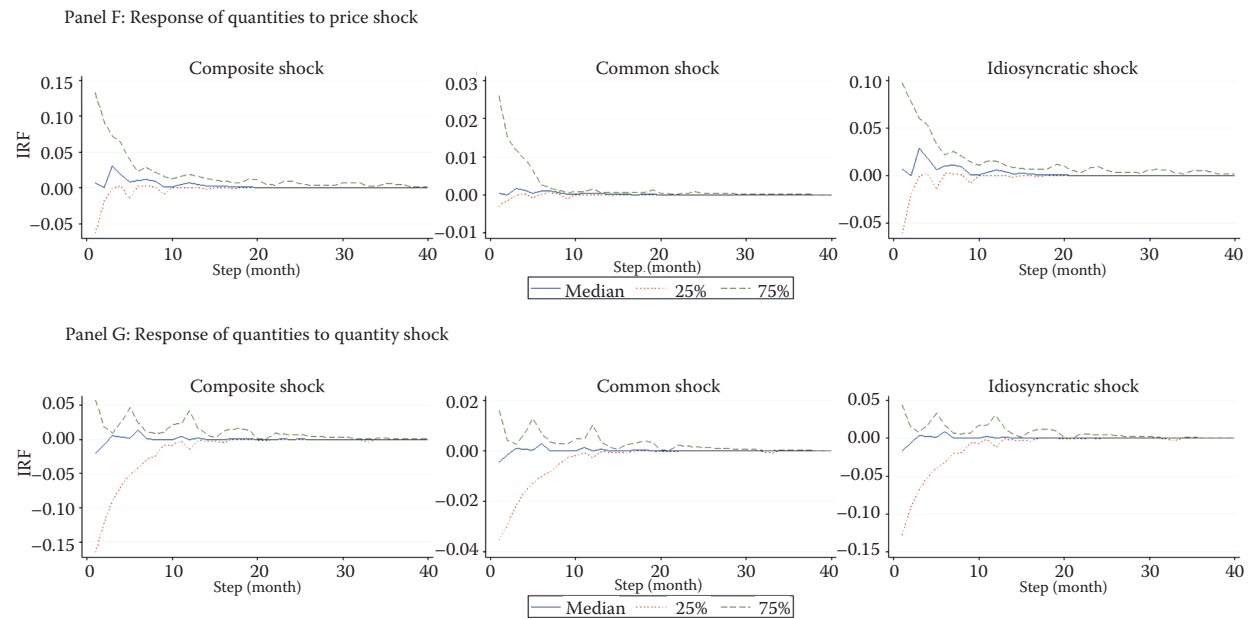


Figure S5. Other impulse response functions (IRFs) estimated from the benchmark model

Source: authors' elaboration