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August 2026

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The Inflationary Effects of Global Supply Chain Shocks: Evidence from Swedish Microdata*

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Sveriges Riksbank Working Paper Series
No. 470
August 2026

Abstract

We compile a unique dataset linking micro price data underlying the official Swedish producer price index with administrative firm level data and provide new evidence on the inflationary effects of global supply chain shocks. For identification, we interact exogenous shocks to global supply chains, obtained through a VAR model, with firm-specific import shares. Shocks to global supply chains lead to a significant and persistent increase in producer prices with a peak response after two years. Importantly, average responses mask heterogeneous responses across firms. Relatively larger firms, firms with lower labor costs and a higher market share raise prices more strongly.

Keywords: Global supply chain shocks, producer prices, microdata, firm characteristics, price setting

JEL Classification: E31, F14, F61

*We thank the editor of this journal and two anonymous reviewers, participants at the 17. FIW-Research Conference International Economics, the European Commission's Annual Research Conference 2024, the ECB-IMF-IMFER Conference 2024 on "Global Challenges and Channels for Fiscal and Monetary Policy", the 2024 Ghent University Workshop on Empirical Macroeconomics, the 2024 Annual Meeting of the Verein für Socialpolitik, the 2024 Padova Macro Talks, the 5th LB-NBP-CEBRA Biennial Conference on "Macroeconomic adjustments after large global shocks", the 24th IWH-CIREQ-BOKERI Macroeconometric Workshop in Halle, and the research seminar at UW Milwaukee, Lukas Hack and our discussants, Boreum Kwak, Tim Schmidt-Eisenlohr, Jakub Mućk, and Basile Grassi, for helpful comments. The opinions expressed in this article are the sole responsibility of the authors and should not be interpreted as reflecting the views of Sveriges Riksbank, the Deutsche Bundesbank or the Eurosystem.

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1 Introduction

Over the past decades, the globalization of trade has been accompanied by the increasing integration of global supply chains. The drawback is that firms become susceptible to disruptions in the supply of intermediate goods. Recent disruptions such as natural disasters, the COVID-19 pandemic and geopolitical tensions have highlighted the vulnerability of supply chains and their critical role in economic activity. In particular, supply chain bottlenecks are considered a major driver of the surge in global inflation since 2021.

A recent literature estimates the macroeconomic consequences of supply chain disruptions (Carrière-Swallow et al., 2023; Burriel et al., 2023; Ascari et al., 2024; Laumer, 2023; Khalil and Weber, 2022; Finck and Tillmann, 2023; Liu and Nguyen, 2023; Elsayed et al., 2025; De Santis, 2024; Bai et al., 2024; Andriantomanga et al., 2024; Mućk and Postek, 2025). The results suggest that supply chain shocks lead to a deterioration of economic activity and an increase in prices.[1] However, this literature focuses on the consequences for aggregate price levels, not dis-aggregate prices.

In this paper, we aim to fill this gap by estimating the causal effect of global supply chain disruptions on the price setting of Swedish firms. We add to the small branch of the literature that traces the effects of supply chain disruptions on the firm-level. For that purpose, we compile a unique dataset linking micro price data underlying the official Swedish producer price index (PPI) with administrative firm level data. The dataset includes around 280,000 individual price observations from a bit less than 2,600 unique firms.

[1] This result is also supported by quantitative general equilibrium models (Di Giovanni et al., 2022; Alessandria et al., 2023).

This allows us to estimate the price response at the product level. The rich set of firm level information enables us to shed light on differences in the responses of producer prices to supply disruptions across the distribution of firms. Thus, we can take account of the cross-sectional heterogeneity in the pass-through of cost shocks to prices, which is emphasized in Amiti et al. (2019), Dedola et al. (2021) and others.

Most papers drawing on microdata to study the adjustment to supply disruptions use data on firm-level quantities only. Boehm et al. (2019) estimate firm-level production elasticities between foreign and domestic inputs. The authors use the cross-country spillovers of the 2011 Tōhoku earthquake in Japan for identification and find a high degree of complementarity: a drop in imports leads to an almost one-for-one fall in output of Japanese affiliates based in the US. Carvalho et al. (2021) also exploit the 2011 earthquake and show that natural disasters propagate backwards and forwards through global supply chains. Lafrogne-Joussier et al. (2023b) compare firm-level data from France during the COVID-19 pandemic with pre-pandemic data. They show that firms relying on intermediate goods from China cut their input more strongly during the pandemic than other firms. Di Giovanni et al. (2024) also use microdata on the value added of French firms. They find that foreign aggregate shocks, e.g. a recession abroad, hit relatively large firms more than relatively small firms.

Our paper contributes to the literature using microdata to study the effects of supply chain bottlenecks on price setting rather than production quantities. Auer et al. (2019) estimate the cross-border transmission of cost increases based on input-output linkages using sectoral data across countries.

Santacreu and LaBelle (2022) study the inflationary effect of supply chain bottlenecks during the COVID-19 pandemic. They regress the industry-specific changes in U.S. producer prices on measures of supply chain pressure interacting with the cross-industry variation in the dependence of intermediate goods across countries. Isaacson and Rubinton (2023) use data on the commodity-specific exposure of U.S. imports to international shipping prices. They find a small pass-through of changes in shipping costs to import price inflation. Meier and Pinto (2024) study the propagation of supply chain disruptions originating in China to sectoral U.S. data during the COVID-19 pandemic. For identification, they exploit pre-pandemic cross-sectional variation in the intensity of imported goods from China.

Lafrogne-Joussier et al. (2023a) use microdata underlying the French PPI over the period January 2018 to July 2022 as well as data on firms' exposure to imported goods and energy costs. They find that an increase in foreign costs by 10 percent causes output prices to increase by 0.74 percent. Acharya et al. (2025) exploit the variation in the perception of supply chain disruptions across European firms at the product-country level. The authors interact this perception with a dummy that reflects the COVID-19 pandemic.

After constructing the micro dataset, we proceed in three steps. First, we estimate a structural vector autoregression (VAR) for Sweden. The model contains data on aggregate macroeconomic variables, as well as on international container trade such as the number of containers processed in the most important North Sea ports, the price of shipping a container and the Global Supply Chain Pressure Index of Benigno et al. (2022). The model is identified using a combination of conventional and narrative sign restrictions drawn from recent episodes of supply chain disruptions, i.e. the Tōhoku earthquake (2011), the Suez Canal obstruction (2021) and the Shanghai backlog as a consequence of the zero-COVID policy of Chinese authorities (2022). As a result, we obtain a series of structural shocks. We refer to these shocks that distort the supply of international container trade as supply chain shocks.[2]

In a second step, we use the granular dataset on product-level producer prices underlying the Swedish producer prices index, combine it with administrative firm level data, and estimate the response of individual prices to the aggregate supply chain shock in a panel local projection (Jordà, 2005) framework. Identification is achieved by combining the exogenous, aggregate shock with firm-specific import shares. We use the share of imports in total sales as a proxy of a firm's exposure to global supply chains. The interaction with the firm-specific import share translates the aggregate shock into firm-specific supply chain disturbances.

We find that global supply chain shocks cause a significant and persistent increase in producer prices. Following a shock of one standard deviation, firms raise prices by about two percent. The price response peaks after two years. Hence, firms pass the shortage of intermediate goods and their price increases to customers in terms of higher prices. Importantly, the average price response masks a considerable degree of heterogeneity in the extent of price adjustment across firms. In a third step, we distinguish firms along the lines of key characteristics such as size, cost structure and market share, among others. We find that relatively larger firms, firms with low labor costs

[2]This labeling is in line with many papers on the macroeconomic consequences of supply chain disruptions, though the literature has not yet reached a consensus on the terminology. An alternative label would be shipping shock.

and firms with a higher product-specific market share increase prices more strongly. Hence, the heterogeneity across firms matters for the pass-through of supply disruptions to prices.

2 Microdata from Sweden

This study relies on three main data sources for Swedish firm-level data.

Producer prices The price data is obtained from the micro dataset underlying the official Swedish producer and import price index (*Prisindex i producent- och importled, PPI*). The PPI data includes monthly product-level prices for a representative sample of goods produced in the Swedish domestic market or produced abroad and imported to Sweden. Thus, the data contains domestic, export and import producer prices. The unit of observation is the price of a product-level transaction, specifically referred to as a "product offering", which is a unique combination of a product sold by a particular firm. Firms that are selected from the sampling procedure are asked to report the price of a representative product they sell within a narrowly defined product code, given by the 8-digit Combined Nomenclature (CN) classification, and are required to respond by law. Our original PPI data covers the period January 1992 to September 2022 and include roughly 1.75 million product-month observations. The PPI dataset is maintained by Statistics Sweden.

Import share We interact the aggregate global supply chain shocks with a firm-level measure of import intensity to obtain different treatment intensities across firms. Data on firm-level import shares is constructed by using import data on foreign trade of goods (*Utrikeshandel med varor, UHV*) and data on total sales from the credit bureau *Upplysningscentralen (UC)*. The UHV data contains exports and imports of goods, both within and outside of the European Union (EU), and is maintained by *Statistics Sweden*.

UHV is a combination of the data collected on within-EU trade (*Intrastat*) and trade with countries outside of EU (*Extrastat*). Within-EU trade is collected using a survey that encompasses all trade above a specific threshold (in 2022, the threshold for export was 4.5 mil SEK and the threshold for import was 9 mil SEK), whereas trade outside of EU includes all trade and is collected by the customs agency. The UHV data includes all countries and products by CN code which a particular firm imports or exports in a given month. Our original UHV data on imports includes around 70 million firm-product-month-country observations spanning between 1997 and 2023.

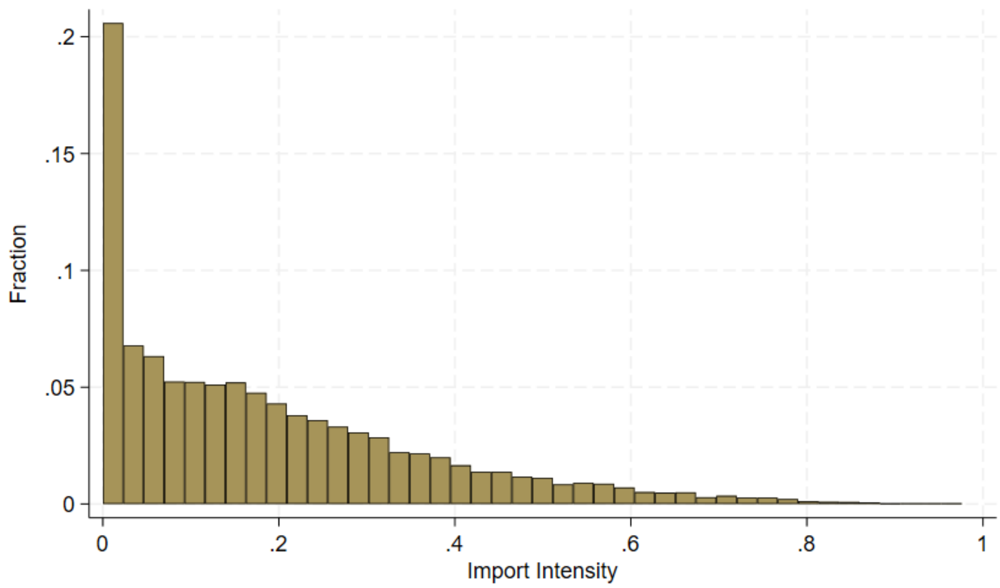
Figure (1) depicts the cross-sectional distribution of the import share. Around 20% of the firms in our sample have an import share close to zero. The distribution of the import intensity of the remaining firms is very wide. As in Lafrogne-Joussier et al. (2023a), firms differ substantially in their exposure to supply shocks. The import share exhibits sufficient variation to be used as a treatment variable that measures the firm-specific exposure to global supply chains.[3]

As an alternative exposure measure one could normalize imports by variable costs as done by Lafrogne-Joussier et al. (2023a). Such a measure might better capture marginal costs which matters in most models for firms' pricing

[3]See Antràs (2020) for a survey on the measurement of supply chain integration.

decisions. However, information on total variable costs, which we measure by the sum of the firm's wage bill and intermediate consumption, has lower coverage in our sample than total sales. Therefore, we use import shares as our baseline exposure measure but show below that the main results are robust to the alternative cost measure.

Figure 1: Distribution of import shares



Notes: The figure shows the distribution of the import share across Swedish firms. The data comes from *Statistics Sweden*'s collection of data on foreign trade of goods (UHV) and *Upplysningscentralen* (UC).

Balance sheet data We utilize additional annual balance sheet and income statement data, obtained from the credit bureau *Upplysningscentralen*, covering the entire population of Swedish corporations (*aktiebolag)*. In particular, we use data on total sales of firms in the computation of import shares. Additionally, we draw from this data source information on firms' total assets, inventories, and cost structure. The original UC data is available between 1989 and 2020.

To match the monthly frequency of the import data with the annual frequency of total sales from UC, we take the sum of the value of all imports made by a firm in a given year. The firm-specific import share is computed as the value of all imports within a year divided by the total sales for the same year.

Below, we draw on the following product and firm characteristics, respectively: firm size measured as (log) sales, labor costs as the ratio of the nominal wage bill and real sales[4], the ratio of inventories to sales, the product-specific market share, the number of sold products and the number of sold 4-digit CN product groups.[5] The online appendix reports summary statistics on the firm characteristics. One should keep in mind that due to the several rounds of data cleaning and matching across datasets, the final matched PPI dataset is not exhaustive.

A unique firm identifier allows us to merge information from all three datasets. Thus, we are able to link product-level prices from the PPI data to the firm-level for which we know import shares from the UHV data and other real and balance sheet information from the UC data. This micro dataset enables us to investigate the inflationary effects of global supply chain shocks at the very granular level and makes it possible to study differences in pricing

[4]In a large class of macroeconomic models, unit labor cost is a measure of marginal cost (Carlsson and Skans, 2012).

[5]Two products belong to the same product group if they have the same first 4 digits in their 8-digit CN codes.

decisions across firms.

We perform several rounds of cleaning the raw data. With regards to the PPI data, we drop a small number of missing, erroneous (negative) or duplicated price observations. Furthermore, we restrict the sample to products belonging to product groups B and C as defined by the Swedish Standard for Product Classification by Industry (SPIN), i.e. products sold within the industrial sector. In our analysis, we only include price series pertaining to output prices for sales in the domestic economy

We assemble the final dataset by merging the monthly product-level PPI price and yearly firm-level UHV import series, while at the same time merging additional variables constructed from the UC data. Our final dataset includes around 280,000 individual price observations from a bit less than 2,600 unique firms.

3 Deriving the supply chain shock: the VAR model

We extract the fundamental supply chain shock from a VAR estimated at the macro level and will later feed it into the micro-level model. Specifically, we are interested in the structural VAR of the form

$$\mathbf{y}_t'\mathbf{A}_0 = \mathbf{c} + \mathbf{y}_{t-1}'\mathbf{A}_1 + ... + \mathbf{y}_{t-p}'\mathbf{A}_p + \boldsymbol{\varepsilon}_t', \tag{1}$$

where $\mathbf{y}_t$ is an $n \times 1$ vector which contains the endogenous variables, $\mathbf{A}_1, ..., \mathbf{A}_p$ are $n \times n$ matrices of parameters and $\mathbf{c}$ is a $1 \times n$ vector of parameters. $\boldsymbol{\varepsilon}_t$ is an $n \times 1$ vector of structural shocks and $\mathbf{A}_0$ is an invertible $n \times n$ matrix which contains the contemporaneous relationships among the endogenous variables. The model can be rewritten in a compact form as

$$\mathbf{y}_t'\mathbf{A}_0 = \mathbf{x}_t'\mathbf{A}_+ + \boldsymbol{\varepsilon}_t', \tag{2}$$

where $\mathbf{x}_t$ is a $(np+1) \times 1$ vector given as $\mathbf{x}_t' = \left[1, \mathbf{y}_{t-1}', ..., \mathbf{y}_{t-p}'\right]$ and $\mathbf{A}_+' = \left[\mathbf{c}' \ \mathbf{A}_1'...\mathbf{A}_p'\right]$ is of the dimension $(np+1) \times n$.

We adopt the model specification as in Finck and Tillmann (2023), but estimate the model for the case of Sweden. The vector of endogenous variables comprises six variables, which are divided into two blocks. The first block includes three variables that reflect the business cycle in Sweden, i.e. industrial production, the index of consumer prices and the import price index. These series are included in natural logs (times 100) and are taken from Statistics Sweden. The estimation sample covers July 2007 to January 2023.

The second block consists of the remaining three variables reflecting international container shipping and global supply chains, respectively. The first is the RWI/ISL container throughput index provided by the Leibniz-Institut für Wirtschaftsforschung (RWI) in Essen, Germany, and the Institute of Shipping Economics and Logistics (ISL) in Bremen, Germany.[6] In the absence of data on container throughput at Swedish ports, this series should be a good approximation of container trade in Sweden. The appendix provides a

[6] The index reports the (seasonally adjusted) number of processed containers in the North Range, i.e. the ports of Le Havre, Zeebrugge, Antwerp, Rotterdam, Bremen/Bremerhaven and Hamburg. These are the most important ports for container trade in the euro area. See Döhrn and Maatsch (2012) and Döhrn (2019) for more information on the RWI/ISL index.

comparison between the monthly throughput index and the annual throughput of containers in Swedish ports. Since 2010, when the annual data starts, both series behave similarly.

The second shipping variable is the HARPEX PETERSEN Charter Rates Index which reflects the worldwide price development on the charter market for container ships. The HARPEX tracks prices for container shipment of semi-finished or finished products.

The last variable we include in our VAR is the Global Supply Chain Pressure index (GSCPI) provided by the Federal Reserve Bank of New York (Benigno et al., 2022). These authors construct a summary indicator of global supply chain pressure based on Purchasing Managers' Index (PMI) surveys for manufacturing firms in different countries, measure of transportation cost such as the HARPEX and the Baltic Dry Index and indicator of airfreight costs.

In order to identify global supply chain shocks, we follow Antolín-Díaz and Rubio-Ramírez (2018) and combine conventional sign restrictions and narrative restrictions. Narrative restrictions are meant to constrain the admissible set of structural parameters by ensuring that around selected historical events the structural shocks, historical decompositions (or both) align with the established narrative.

We restrict three of our six endogenous variables, while the impulse responses of the remaining three variable are left unrestricted. Importantly, we do not impose any price response following the supply chain shock and thus we are agnostic about the inflationary effects of the identified shock.[7]

[7] In contrast, other papers identify a supply chain shock by directly imposing a negative

We assume that a supply chain shock leads to an increase in the HARPEX, i.e. to an increase in charter rates for container vessels. This increase is accompanied by a reduction in the number of containers being processed, i.e. a reduction in container throughput. This negative co-movement therefore identifies a supply shock. Finally, in order to distinguish our supply chain shock from any other supply shock, such as technology or oil price shocks, we assume that the shock also increases supply chain pressure. We therefore additionally restrict the response of the global supply chain pressure index and assume that this index rises after a supply chain shock.[8] All restrictions hold on impact and for the two subsequent months.

In addition to the conventional sign restrictions, we also use narrative sign restrictions. We exploit the same episodes of supply chain disruptions as in Finck and Tillmann (2023), i.e. the Tōhoku earthquake (2011), the Suez Canal obstruction (2021) and the Shanghai backlog as a consequence of the zero-COVID policy of Chinese authorities (2022). We impose the restriction that the supply chain shock in all three periods must be positive, i.e. restrictive. Moreover, we assume that the shock in March 2011 (Tōhoku earthquake) is the most important driver of the GSCPI. The online appendix contains the aggregate impulse response functions. In line with other studies (Ascari et al., 2024), we find that an exogenous global supply chain disruption leads to a fall in industrial production coupled with a significant rise in import and consumer prices.

We discuss the main results in the online appendix. For the remainder

co-movement between prices and output (Ascari et al., 2024).

[8] Importantly, an increase in the GSCPI is supply-side driven by design, as the GSCPI is purged of demand-driven factors.

of the paper, we need the series of identified supply chain shocks only. The appendix shows the impulse responses of the endogenous variables to a supply chain shock and a figure showing the supply chain shock itself, which we use later for the local projections. Importantly, our supply chain shock is robust with respect to changing the narrative restrictions. The structural shock series we obtain if we include the narrative restrictions one-by-one generates very similar product-level price responses as documented in the appendix.

4 The firm-responses to supply chain shocks

In this section, we investigate the average response of firms to supply chain shock. The next section sheds light on the role of firm heterogeneity.

4.1 Model

To evaluate the effects of global supply chain shocks on Swedish producer prices, we use panel local projections (Jordà, 2005) at the individual product level and estimate the following for each horizon $h = 0, ..., 36$ equation

$$\begin{aligned} \log\left(p_{i,j,f,t+h}\right) - \log\left(p_{i,j,f,t-1}\right) &= \alpha_{j,h} + \alpha_{m,h} + \beta_h \left(share_{f,t} \times shock_t\right) \\ &\quad + \gamma_h X_{t-1} + u_{i,j,f,t+h}, \end{aligned} \tag{3}$$

where $p_{i,j,f,t}$ is the price of product i, belonging to product group j and produced by firm f in month t. $\alpha_{j,h}$ are product group fixed effects to filter out any unobserved heterogeneity across product groups. The $\alpha_{m,h}$ coefficients are monthly fixed effects to control for seasonal price movements.

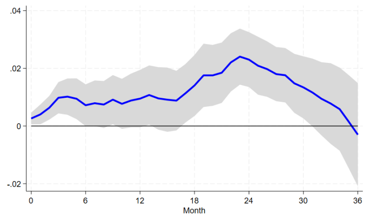
A vector X_t collects additional aggregate control variables, including 12 lags of the unemployment rate and of log industrial production. The projection residual is $u_{i,j,f,t+h}$. $shock_t$ is the aggregate global supply shock that we obtain from the VAR model as described in Section 3. The variable $share_{f,t}$ is the firm-specific import share that we introduced before. By interacting the aggregate shock measure with the firm-specific import share, we allow for different shock treatments across firms. In particular, we assume that a firm with a higher (lower) import share is more (less) affected by disruptions to global supply chains. The import share can be seen as a proxy for the integration into supply chains and thus the more (less) integrated a firm is, the stronger (weaker) the impact of global supply chain shocks should be. We are mainly interested in the coefficient β_h, which measures how prices change over the h periods when a firm-specific shock to global supply chains happens in t. Throughout, we use Driscoll and Kraay (1998) standard errors, which take into account the potential residual correlation across firms, as well as serial correlation and heteroskedasticity among the residuals over time.

4.2 Baseline results

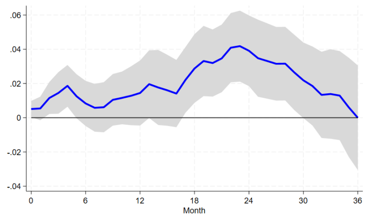
We now study the response of producer prices to the global supply chain shock. The size of the shock is one standard deviation. All figures show the estimated β_h coefficients together with a 90% confidence band constructed from Driscoll-Kraay (Driscoll and Kraay, 1998) standard errors.

Panel (a) of Figure (2) shows the response in our baseline model. A supply chain disruption causes a significant increase in domestic producer prices. The price increase becomes significant one month after the shock and reaches a maximum of two percent after about two years. Thus, producers raise goods prices and pass a part of the higher costs of intermediate goods to customers. As the model is linear, this also implies that a relaxation of supply chain stress contributes to a drop in producer prices.

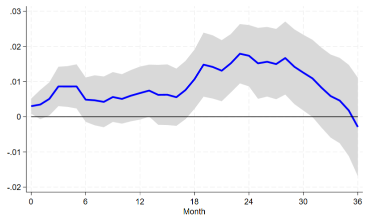
Figure 2: Response of product-level producer prices



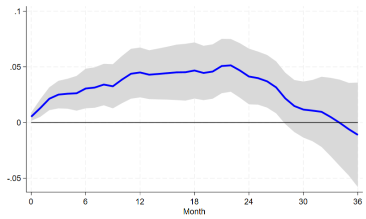
(a) baseline model



(b) weighted price observations model



(c) including time fixed effects



(d) reduced form shock

Notes: The figure shows the response of firm-level producer prices to a global supply chain shock. The 90% confidence band is based on Driscoll-Kraay standard errors.

Panel (b) of Figure (2) reports the estimated impulse response function from an alternative specification, in which we weight each price by its weight in the aggregate Swedish producer price index. The shape of the response remains unchanged: the global supply chain shock prompts firms to set a significantly higher domestic price. Interestingly, the size of the effect is much larger than in the baseline model with prices increasing by about four percent. If all firms respond similarly to the shock, the weighting of the responses would not change the overall response. The fact that the weighted responses generates an effect twice as large as the effect of the unweighted firm responses points to an important role of the cross-sectional distribution of price responses.

In order to control for aggregate determinants of price setting not captured by the control variables in the vector X_t, we estimate a model specification with time fixed effects. The results are shown in Panel (c) of Figure (2). Producers start to raise prices in the first months after the shock with the peak response occurring about two years after the shock. The magnitude and the shape of the impulse response is almost indistinguishable from the baseline response, which suggests that there is no systematic information left for the time fixed effects to pick up.

The previous results were based on the structural supply chain shock derived from the identified VAR model. To gauge the robustness of the estimated effects, we use an alternative shock series. As a reduced-form shock, we use the residual from a regression of the Global Supply Chain Pressure Index (Benigno et al., 2022) of three lags of itself as well as current realizations and three lags of world industrial production (in log differences), Global Economic Conditions and the real price of oil (in log differences). The variable Global Economic Conditions is taken from Baumeister et al. (2022). In addition, we include a dummy for the global financial crisis, the COVID-19 pandemic and the full-scale Russian invasion of Ukraine.

Panel (d) of Figure (2) shows response of producer prices to the reduced-form shock. Prices increase more strongly than in the the baseline model shown in Panel (a) with a peak response of about five percent reached after about two years. Furthermore, the response to the structural shock is estimated more precisely. Take the response at $h = 12$ as an example. The 90% confidence band is narrower if we use the reduced shock compared to the confidence band surrounding the structural shock. Despite the somewhat narrower confidence band, we proceed with the structural shock from our VAR model. We believe the additional structure imposed by the VAR model is essential for identifying the shock.

The baseline results are also robust to using an alternative exposure measure. When measuring the exposure to supply shocks by the ratio of imports over total variable cost (instead of imports to total sales as in the baseline), we find a very similar price response as shown in the appendix.[9]

In addition, in the appendix, we show that our baseline results are robust to including additional disaggregated fixed effects and controls. In particular, the price response is very similar when, instead of of product group and month fixed effects, we include i) product and month fixed effects, ii) product group, firm and month fixed effects, or iii) product-group times month fixed effects. Moreover, we obtain a similar price response when additionally including lags of the firm-specific import share and of the product-specific month-to-month inflation rate.

Finally, the appendix presents the results of an interacted model where we include the aggregate shock, the firm-level import share and the interaction of both as separate regressors. The estimated coefficients on the interaction

[9] We measure total variable costs as the sum of wage cost and intermediate consumption, which includes expenditure on raw materials, goods, and services.

term are very similar to the ones of our baseline model.

5 Heterogeneous firm-responses to supply chain shocks

The key contribution of this section is to investigate potential dependencies of the inflationary effects of global supply shocks across firms on firm characteristics.

5.1 Model

We extend the linear specification (3) and estimate a set of different interacted panel local projections of the following form

$$\begin{aligned}\log\left(p_{i,j,f,t+h}\right)-\log\left(p_{i,j,f,t-1}\right) \;=\;& \alpha_{j,h}+\alpha_{m,h} \\ &+I_{f,t-1}\left[\alpha_h^{High}+\beta_h^{High}\left(share_{f,t}\times shock_t\right)\right] \\ &+\left(1-I_{f,t-1}\right)\left[\alpha_h^{Low}+\beta_h^{Low}\left(share_{f,t}\times shock_t\right)\right] \\ &+\gamma_h X_{t-1}+u_{i,j,f,t+h},\end{aligned} \tag{4}$$

where $I_{f,t-1}$ is a dummy variable that captures the specific firm property we are looking at.

We are interested in studying particular firm characteristics that could significantly influence the inflationary effects of supply chain shocks such as size or market share. The dummy variable $I_{f,t-1}$ indicates whether a firm f belongs to the group of firms for which this specific characteristic lies above the median across firms at a given point in time. We will refer to this firm characteristic as the state variable. We include a one-period lag of $I_{f,t-1}$ in the regressions in order to minimize contemporaneous correlations between the supply shock and the realization of the state variable. Given our specification, β_h^{High} provides an estimate of the price response in state *High*, i.e. a firm's realization of the state variable is above the cross-section median, whereas β_h^{Low} provides the estimate in state *Low*. We need the state variable only to distinguish between between firms above or below the median with respect to a specific property. We do not look at variation of the state variable as such. If the state variable fluctuates over time but the firm does not switch between the *High* and the *Low* state, these fluctuations do not affect our estimates. We use the maximum length of the price data (until September 2022). As we make a projection over three years, the last shock we can use for the estimation of the local projections is from 2019. Since the state variable is lagged relative to the moment of the shock, the last realization of the state variable is taken from the year 2018. Hence, the missing balance sheet data after 2020 is not a binding constraint.

When testing for state-dependencies across time rather than across firms, e.g. as a result of fluctuations of aggregate pressure on supply chains, $I_{f,t-1} = I_{t-1}$. In this case, the state is the same across all firms but varies across the time dimension, months in our data. In general, we condition the effect of supply chain disruptions on the firm characteristics, each transformed into a binary state variable $I_{f,t-1}$.[10]

[10] As an alternative, one could fix the firm characteristics across time to minimize the noise resulting firm operating close to the binary threshold and thus repeatedly switch categories. As shown below, our main qualitative findings are robust to using such a

5.2 Firm characteristics

Before we show the results, we briefly survey the recent literature on the role of selected firm properties for price setting.

Size and market share A natural starting point is to compare the response of small and large firms. Hence, the first characteristic is size. Firms are considered small (large) if their (log) sales are below (above) the median of the distribution of firms at time t. It is important to stress that in all our estimated state-dependent models, the distinction across firms is dynamic. Thus, the status of firms can change if their sales increase or decrease relative to their peers over the sample period.

We also study the role of market shares. This variable is constructed as follows. First, we take the value of production for each product and firm, and divide it by the total value of production by any firm for products who have the same first two digits in their CN code. This is done on a year-by-year basis. Second, by firm and year, we take the average of the product specific market shares (computed in step one). The average is weighted by the value of production for that product by that firm. This way, even if a firm has a large market share for one product, if they produce very little of this product, then it will have a lower weight in the average.[11]

Theoretically, the impact of specific firm characteristics on the pass-through of supply chain shocks is ambiguous. Let us start with firm size

time-invariant threshold.

[11]As an alternative, we construct market shares at the product-firm-year level and thus classify products according to their market shares. We show below that results are qualitatively similar when using this alternative measure.

and market share. Textbook logic suggests that larger firms have more market share and, hence, are able to pass on a larger share of the increase in the price of intermediate goods imported from upstream suppliers to producer prices. However, recent research sheds light on two factors which could turn this effect around. First, Amiti et al. (2019) focus on the role of strategic complementarities in price setting. If small firm exhibit less strategic complementarity in pricing with respect to their competitors than large firms, they will more strongly pass through cost increases than large firms. Second, Guilloux-Nefussi (2020) uses a New-Keynesian model with strategic interaction among firms. She argues that globalization causes a shift in resources to large firms, which also have high market power. To the extent the fall in trade costs due to globalization expands the production of large firms, these firms exhibit a smaller pass-through of cost shocks to prices. The reason is that firms strategically allow their desired markup to fall in order to absorb the cost pressure.

The empirical findings of the literature on the role of size and market power for the price response to supply shocks are indeed mixed. Dedola et al. (2021) find that smaller firms respond less to a shock to the costs of imported goods than larger firms. This finding suggests that large firms with pricing power try to gain market share as stressed in the literature on strategic as in Amiti et al. (2019). In contrast, Parlapiano (2025) uses data for Italian firms and shows that large firms respond more strongly to a shock to intermediate goods prices. Auer and Schoenle (2016) examine how market shares and price complementarities influence the pass-through decisions of firms. They find that the response of import prices to changes in the exchange rate is U-shaped in the market share. Their theoretical model shows that large firms can pass-through a large share of the cost increase due to its market power. A small firm behaves similarly, not because of market power, but because it has very little market share to lose.

Franzoni et al. (2024) use firm-level data from a cross-country data set and find that when shortages occur, larger firms gain higher market share, profitability, markups and stock returns, while benefiting from smaller cost increases. Overall, they find that shortages are associated with larger price adjustments, especially in ex-ante concentrated markets. Bräuning et al. (2023) estimate that margins decrease less after an adverse cost shock in more concentrated industries. In contrast, Kharroubi et al. (2023) study the pass-through of oil price shocks to producer prices in U.S. sectors. As a key result, they show that sectors with a higher markup exhibit a weaker pass-through. Amiti et al. (2014) argue that import share and market shares are key determinants of the pass-through of cost shocks. They find that large firms adjust prices less strongly than small firms after exchange rate fluctuations. A small exporting firm, which at the same time does not rely on imported goods, exhibits a complete pass-through. In contrast, a firm that is large in terms of market share and in terms of import share has a much lower pass-through.

Cost structure We shed light on the cost structure of firms as a determinant of the size of the price adjustment. We define labor costs as the ratio of the nominal wage bill and real sales. Suppose the production is labor intensive, that is, the labor share is relatively high. This implies that the firm spends relatively less on intermediate inputs and the shock to supply chains has a small impact on total costs. In the absence of strong complementarities of factor inputs, the firm should raise prices modestly if exposed to supply disruptions. Another firm with relatively low labor costs should raise prices more aggressively. In the presence of complementarities in inputs, labor costs alone are not particularly informative about the pressure to raise product prices.

Inventory holdings To the extent the global supply shock disrupts the sourcing of raw materials, we expect that a low stock of inventories prompts firms to raise prices more strongly. Hence, we condition the impact of the shock on the stock of inventories. We scale the stock of inventories by the level of sales for each firm and again split the sample at the median of the cross-sectional distribution.

Intuitively, firms with a higher stock of inventories should raise prices less than firms with lower inventory holdings. From a general equilibrium model with heterogeneous firms, Alessandria et al. (2023) conclude that an increase in shipping times of intermediate goods causes a stronger increase in prices for lower levels of inventories.[12] Mehrotra et al. (2025) utilize industry-level data for the U.S. economy. They find that the elasticity of retail prices with respect to inventories is negative. A shock that hits an industry with a lower stock of inventories causes a stronger price response. Alessandria et al. (2010) show that supply chain disruptions have an effect on firms'

[12] Barrot and Sauvagnat (2016) show that relatively large stocks of inventories cushion the contractionary effects of natural disasters interrupting supply on firms' production levels. However, they cannot show the effect on producer prices.

inventory levels, which in turn influence pricing. They build a model with inventory frictions, i.e. time-lags between the order and the delivery of goods as well as fixed transaction costs, which ultimately translate into trade costs. After an unexpected exchange rate devaluation, inventories are sub-optimally high, which in turn suggests that importing firms should not increase prices strongly.

Product portfolio Another determinant of the price response is the firm's product portfolio. The theoretical effect is ambiguous. A firm selling multiple products relies on a broader set of inputs and is more exposed to supply shortages than a single-product firm. Multi-product firms may benefit from diversification within their product portfolio, potentially leading them to react less to disruptions affecting some inputs. At the same time, the theoretical prediction crucially depends on assumptions regarding the separability across products.

Chatterjee et al. (2013) study the exchange rate pass through of multi-product firms. Their key finding is an increase in the within-firm heterogeneity of prices after a devaluation. They also let the impact of the real exchange rate on producer prices interact with the number of products the firms sells. The estimated coefficient on the interaction term is positive, thus implying that the pass-through intensifies in the number of firms. In contrast, Friberg and Rohman (2018) come to the opposite conclusion analyzing the pass-through of cost shocks for Swedish beer producers. Firms with a larger product portfolio exhibit a lower pass-through to prices.[13]

[13]See also Dedola et al. (2021) and Duprez and Magerman (2021) for models of the complex price setting in multi-product firms as a response to exogenous shocks.

Table 1: Firm Characteristics Overlap

	total sales	labor costs	market share	inventory ratio	no of products	no. of product groups
total sales	100	34	70	46	60	51
labor share	33	100	35	54	52	43
market share	68	35	100	50	59	50
inventory ratio	46	55	51	100	58	47
no. of products	54	48	54	52	100	83
no. of product groups	55	48	55	51	100	100

Notes: The table displays the percentage overlap between firm indicators. Rows denote the indicator, and column values refer to the percentage overlap. For example, 46% of all month-firm observations for which a firm is classified "large" (as measured by total sales) overlaps with observations in which the same firm also had a high (i.e., above average) inventory ratio.

Given the fact that the theoretical literature offer several mechanism through which firm characteristics interact with shocks to global supply, we remain agnostic about both the precise mechanism and the sign of the interaction. Rather, we use our unique data set together with the identified shock and let the data speak.

Importantly, these properties of firms are not orthogonal. To what extent do the classifications of firms into two types (i.e. high and low) overlap across indicators? For example, are firms classified as large also classified as having a high inventory ratio? We calculate the percentage overlap between selected firm indicators and document it in Table (1). For each indicator, the row provides the percentage overlap with other indicators.[14] Consider again firm size: 46% of all month-firm observations for which a firm is classified large based on total sales overlap with observations in which the same firm is

[14]The appendix shows the unconditional correlations of selected firm characteristics.

classified as having a high inventory ratio. Likewise, large firms tend to be classified as multiple product firms and firms selling multiple product groups. This is consistent with the findings of Bhattarai and Schoenle (2014) on multi-product firms. Large firms are also more frequently classified as firms with a high market share. Overall, the extent of overlap is not too large such that our analysis of the response to shocks conditional on different firm characteristics is not obfuscated by coinciding classifications.

5.3 Results

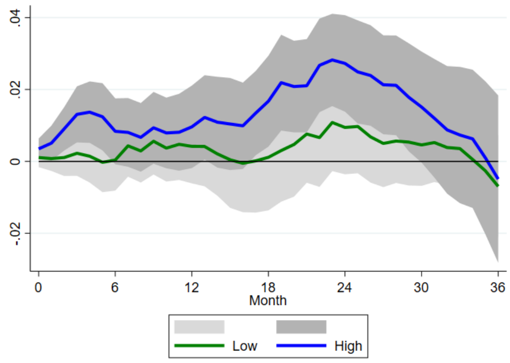
We now turn to the results from the empirical model in which the effect of global supply chain shocks depend on firm properties. This allows us to shed light on the heterogeneity of price responses across firms.[15]

Panel (a) of Figure (3) reveals differences between small and large firms in the pass-through of supply chain shocks. We find that large firms raise prices significantly, while small firms do not. Put differently, large firms pass-on higher costs to their customers through higher producer prices. The price response of large firms peaks after two years at a level of almost three percent. Smaller firms, in contrast, do not charge significantly higher prices. Their response remains indistinguishable from zero through the projection horizon. As discussed before, the connection between firm size and price pass-through in the empirical literature is less clear-cut.

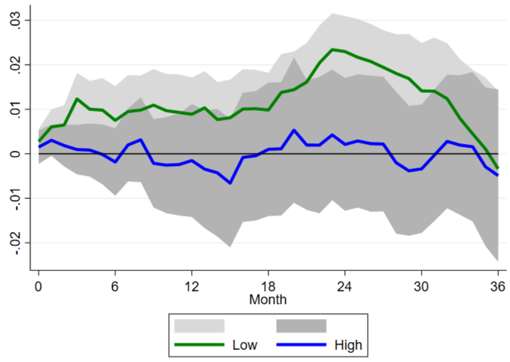
We split the observations between firms with high and low labor costs,

[15] The magnitude of the impulse responses in the state-dependent model is not directly comparable to the responses in the linear baseline model. Since the intercept term of the regression model interacts with firm characteristics, we control for an additional average effect across firms in the state-dependent model, which affects the size of the coefficients.

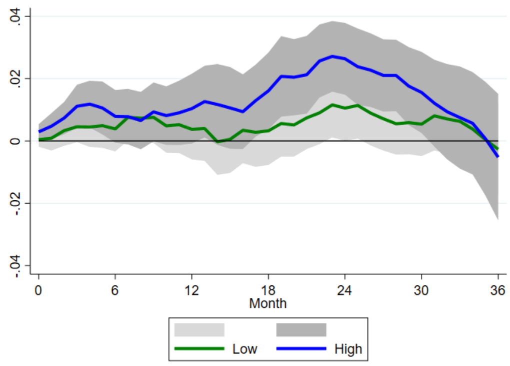
Figure 3: State-dependent response of product-level producer prices



(a) firm size



(b) labor costs



(c) market share

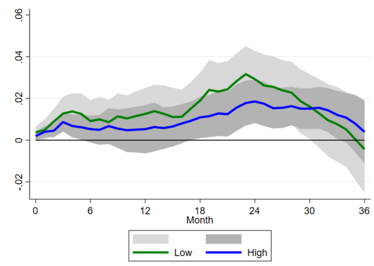
Notes: The figure shows the response of firm-level producer prices to a global supply chain shock. The 90% confidence band is based on Driscoll-Kraay standard errors.

defined as the ratio of the nominal wage bill and real sales. At each point in time, half of the firms belong to either the high or low labor cost group. Panel (b) of Figure (3) supports our notion: after a supply chain shock, firms with relatively low labor costs raise prices. In contrast, firms with relatively high labor costs do not raise prices significantly. This is consistent with the presumption that firms with high labor costs labor do not use price increases as a margin of adjustment in order to react to supply chain disruptions.

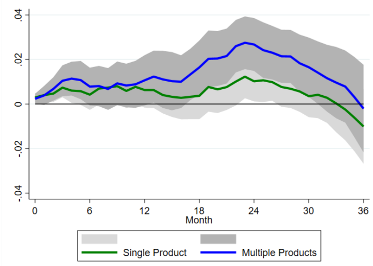
Lastly, we assess the market share of firms as a determinant of shock pass-through. In Panel (c) of Figure (3), it is evident that firms with a higher market share respond more strongly compared to low-market share firms. The price response of the latter set of firms even lacks statistical significance. Put differently, firms with high market share pass-through the supply disruptions to prices, while firms with less market share do not. We find qualitatively similar results when measuring market share at the product-firm-year level as shown in the appendix. Moreover, the main qualitative findings presented in Figure (3) are robust when, instead of allowing for a time-variant state variable, fixing the state across time (see the appendix).

Panel (a) of Figure (4) reveals that firms with a low stock of inventories increase prices more strongly than firms with a high stock of inventories. Multi-product firms increases prices by more than two percent within two years, while single-product firm raise their prices by one percent only, see Panel (b). Panel (c) of the figure shows the impulse responses when we distinguish between the number of different product groups. The empirical finding remains unchanged: firms with a more diversified set of products, here approximated by multiple product groups, respond to the supply chain

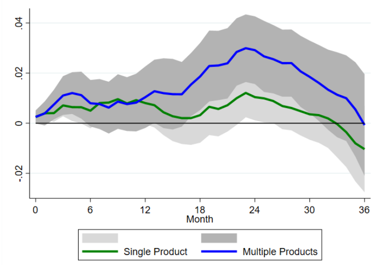
Figure 4: State-dependent response of product-level producer prices



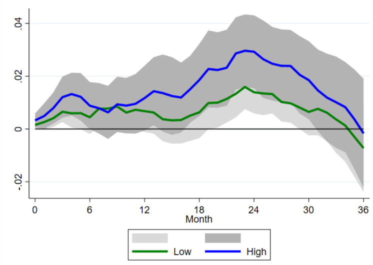
(a) inventories



(b) single- vs multi-product firms



(c) single- vs multi-product firms (product groups)



(d) number of products over number of product groups

Notes: The figure shows the response of firm-level producer prices to a global supply chain shock. The 90% confidence band is based on Driscoll-Kraay standard errors.

shock by setting a higher price than firms producing a single product group.

An alternative way to study the response of firms with a diversified set of products to supply bottlenecks is to look at the ratio of the number of products to the number of product groups. This ratio is a measure of the degree of specialization of a firm's set of products. If a firm produces five products from just one product group, this firm is considered very specialized. If the firm produces five products in two product groups, the firm is less specialized and the ratio is smaller. The more specialized the firm is, the more likely the firm is to suffer from shortages of intermediate products and, hence, the higher will be the pressure to raise prices. Panel (d) of Figure (4) shows the responses when we split the firms along the median of the ratio of the number of products and the number of product groups. The results support our intuition: more specialized firms, i.e. firms in the upper half of the distribution of the ratio of products to product groups, set a higher price than less specialized firms. Thus, our results remain qualitatively unchanged if we condition the effect of the supply chain shock on the number of different products, the number of product groups or the ratio of the two characteristics.

6 Conclusions

Disruptions of global supply chains are seen as a major driver of the surge in inflation after the COVID-19 pandemic. In this paper, we studied the quantitative impact of supply chain shocks on producer prices in Sweden. We assemble a unique data set linking microdata underlying the official Swedish producer price index with administrative firm level data. Importantly, we interact the supply chain shock with firm-specific import shares as a measure of value chain participation.

An adverse supply chain bottleneck causes a significant and relatively persistent increase in producer prices. Our microdata allows us to uncover a significant degree of heterogeneity in the responses across firms. Firms which are relatively large, have relatively low labor costs and a higher market share exhibit a stronger price response than other firms. Thus, these factors amplify the pass-through of supply chain disruptions to producer prices.

Our findings help understand the source of the surge in inflation after 2021 and its fall in 2023, when pressure on global supply chains eased. The enormous heterogeneity in the adjustment of prices across firms makes the design of appropriate stabilization policies in light of supply chain shocks, in particular monetary policy, challenging.

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Appendices

These online appendices are not part of the published paper.

Appendix A Summary statistics

Table (A.1) shows the summary statistics on the main firm characteristics in our data set. Each variable is the average of the available firm observations. The unconditional moments reported in the table reflect the distribution of the firm-specific averages across firms. It is particularly noteworthy that the inventory ratio is surprisingly low. On average, firms hold inventories worth only 2% of their total sales. Despite the low average level of inventories, there is substantial variation in the inventory ratio across firms.

Table A.1: Summary statistics

	mean	median	std. dev.	p25	p75
import share	0.16	0.12	0.17	0.02	0.26
total sales	16.86	16.76	1.27	16.05	17.58
labor costs	24.55	23.22	12.96	15.26	32.03
market share	0.03	0.00	0.09	0.00	0.01
inventory ratio	0.02	0.01	0.03	0.00	0.02
no. of products	3.58	2.00	5.74	1.00	3.60
no. of product groups	2.37	1.40	2.78	1.00	2.62

Notes: The table shows the summary statistics on the main firm characteristics in our data set. Each variable is the average of the available firm observations. The statistical moments reflect the distribution of the firm-specific averages across firms.

Appendix B Firm characteristics

In the main paper, we show the overlap of state variables. In order to complement this, Table (B.2) documents the unconditional correlation of selected firm properties. Large firms, i.e. firms with high log sales, tend to have a higher import share, though the correlation is relatively small ($\rho = 0.15$). Large firms also are more likely to be multi-product firms ($\rho = 0.36$) and sell multiple product groups ($\rho = 0.29$). Overall, the correlations are relatively low.

Table B.2: Correlation of firm characteristics

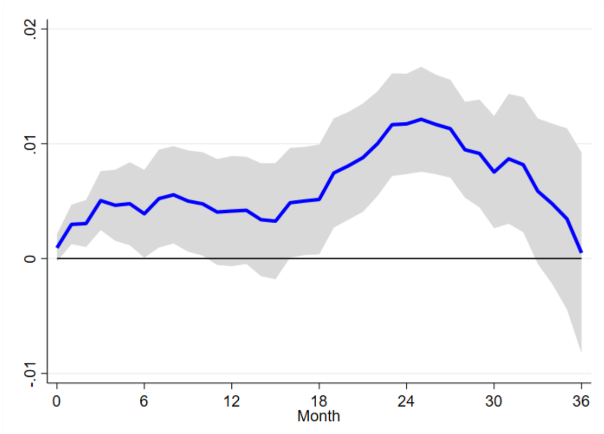
	import share	total sales	labor costs	market share	inventory ratio	no of products	no. of product groups
import share	1.000						
total sales	0.154	1.000					
labor costs	-0.283	-0.458	1.000				
market share	0.118	0.199	-0.096	1.000			
inventory ratio	-0.038	-0.103	0.100	-0.009	1.000		
no. of products	0.128	0.356	-0.169	0.246	-0.035	1.000	
no. of product groups	0.128	0.295	-0.132	0.229	-0.054	0.892	1.000

Notes: The table shows the unconditional correlation of firm characteristics.

Appendix C Alternative exposure measure

In the baseline, we use firms' import shares, the ratio of imports to total sales, as our exposure measure. As an alternative, following Lafrogne-Joussier et al. (2023a), we normalize imports by total variable costs measured as the sum of total wage costs and intermediate consumption which includes expenditure on raw materials, goods, and services. Figure (C.1) presents the results of

Figure C.1: Response of product-level producer prices (alternative exposure measure)



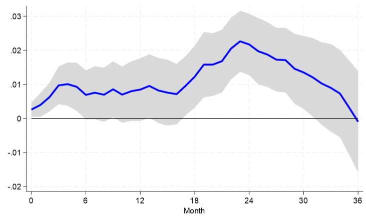
Notes: The figure shows the response of firm-level producer prices to a global supply chain shock. The 90% confidence band is based on Driscoll-Kraay standard errors.

the baseline estimation when using this alternative exposure measure. The results remain unchanged compared to the baseline findings.

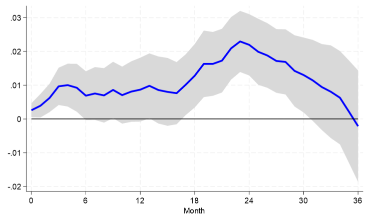
Appendix D Additional fixed effects and controls

Figure (D.2) presents the results of including additional fixed effects and controls in our baseline model. In particular, Panel (a) shows results when including product and month fixed effects, Panel (b) when including product group, firm and month fixed effects, and Panel (c) when including product-group times month fixed effects. Panel (d) presents results of the baseline model when additionally controlling for six lags of the firm-specific import share and six lags of the product-specific month-to-month inflation rate. The estimated responses of all the models deliver similar results compared to the baseline model shown in the main text.

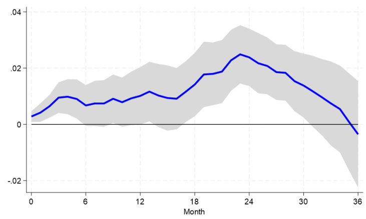
Figure D.2: Response of product-level producer prices (additional fixed effects and control variables)



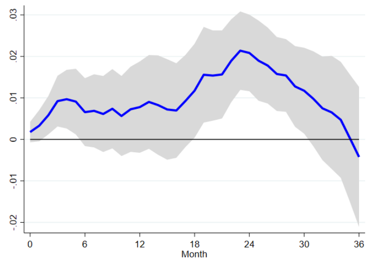
(a) Product and month fixed effects



(b) Product group, firm, and month fixed effects



(c) Product group times month fixed effects



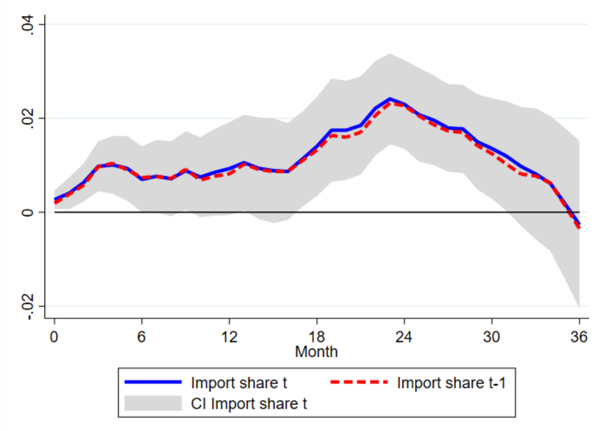
(d) Additional firm controls

Notes: The figure shows the response of firm-level producer prices to a global supply chain shock. The 90% confidence band is based on Driscoll-Kraay standard errors.

Appendix E Lagged import share

We condition the supply chain shock on the firm-specific import share. In our baseline model, we use the contemporaneous import share dated t. It could be argued that a supply chain shock affects the volume of imports differently than total sales. In this case, the import share would endogenously respond to the exogenous shock, which would also affect our estimate of β_h. One way

Figure E.3: Response of product-level producer prices (lagged import share)



Notes: The figure shows the response of firm-level producer prices to a global supply chain shock. The 90% confidence band is based on Driscoll-Kraay standard errors.

to rule out this endogenous adjustment is to use the lagged import share, i.e. $share_{f,t-1}$. Figure (E.3) reports the estimated impulse responses for this specification and compares it to our baseline results. We find that both impulse responses are virtually indistinguishable, which suggests that we can rule out a sizable adjustment of the import share.

Appendix F The interacted model

In our baseline specification, we regress product-specific prices on the interaction term between the supply chain shock and the firm-specific import share. The coefficient on this interaction term provides us with the effect of the supply chain shock on the prices set by firm f conditional on the supply chain integration of that specific firm. We now report the results from an augmented specification that does not only include the interaction term, but also includes the supply chain shock and the import share as separate explanatory variables. Our aim is to show that the main findings of the paper remain unchanged if we include these variables separately.

In Equation (F.1), $\beta_{1,h}$ reflects the unconditional effect of the supply chain shock, $\beta_{2,h}$ reflects the effect of the import share on price setting and $\beta_{3,h}$ is the coefficient on the interaction term that corresponds to the estimated β_h coefficient in the main part of the paper, i.e.

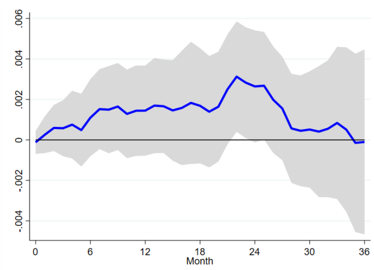
$$\begin{aligned} \log\left(p_{i,j,f,t+h}\right) - \log\left(p_{i,j,f,t-1}\right) &= \alpha_{j,h} + \alpha_{m,h} + \beta_{1,h}\left(shock_t\right) + \\ &\quad \beta_{2,h}\left(share_{f,t}\right) + \\ &\quad \beta_{3,h}\left(share_{f,t} \times shock_t\right) + \\ &\quad \gamma_h X_{t-1} + u_{i,j,f,t+h}. \end{aligned} \tag{F.1}$$

The remaining parts of the regression equation are left unchanged. Panel (a) in Figure (F.4) depicts the estimated coefficient on the supply chain shock as a function of the projection horizon h. The impact of the shock remains insignificantly different from zero throughout the projection horizon. It remains imperative to study the impact of supply bottlenecks across firms with different characteristics as the unconditional effect remains small.

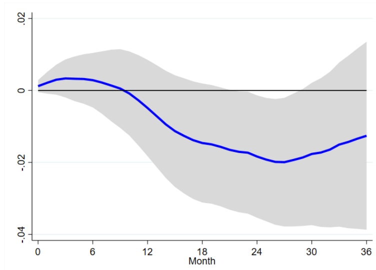
The effect of the import share on prices is shown in Panel (b) of Figure (F.4). An increase in the import share results in a slightly lower price after two years. The effect becomes significant after two years only. Importantly, we should not over-interpret this coefficient because there is little variation in the import share over time and, more importantly, fluctuations in the import share are to some extent anticipated. Since these fluctuations do not come as a surprise, the estimated coefficient does not allow for a causal interpretation.

Panel (c) of Figure (F.4) reports the key finding of this robustness check. The coefficient on the interaction term remains significantly positive if we include the shock and the import shares as separate variables. Compared to the baseline results reported in the main paper, the response is a bit smaller.

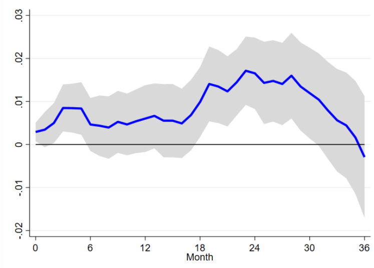
Figure F.4: State-dependent response of product-level producer prices



(a) coefficient on shock



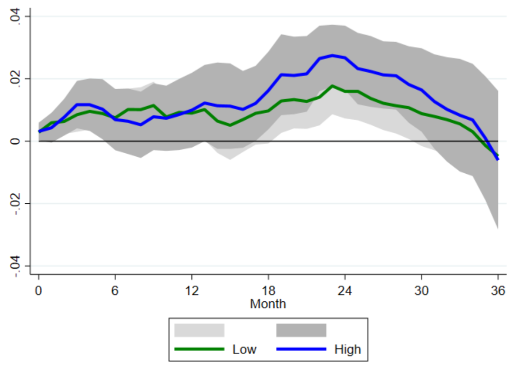
(b) coefficient on import share



(c) coefficient on interaction term

Notes: The figure shows the response of firm-level producer prices to a global supply chain shock. The 90% confidence band is based on Driscoll-Kraay standard errors.

Figure G.5: State-dependent response of product-level producer prices (market share at the product level)



Notes: The figure shows the response of firm-level producer prices to a global supply chain shock. The 90% confidence band is based on Driscoll-Kraay standard errors.

Appendix G Measuring market share at the product level

In the main text, we aggregate market shares at the firm level. As an alternative, we measure market shares at the product-firm-year level and thus classify products according to their market shares. Figure (G.5) presents the estimation results when using this alternative measure as a state variable in the local projections. The results remain unchanged.

Appendix H Fixing state variables across time

In the main text, we allow for a time-varying state variable in the local projections, $I_{i,t}$. As an alternative, one could fix the state variable across time to rule out any noise stemming from firms that frequently switch categories.

Figure (H.6) presents results when fixing the state variable. In particular, we first calculate firm-specific averages of the variables of interest (size, labor costs, market share) across the sample period and then create a dummy variable, which takes the value of one if the average variable of a firm is greater than the median of the firm level averages and zero otherwise. For these calculations, we restrict the sample to firms that are included in at least half of the sample. The results do not differ from our findings in the main part of the paper.

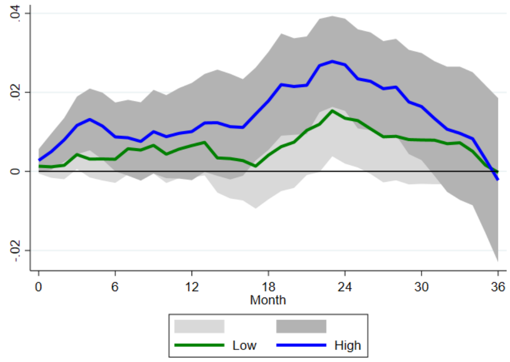
Appendix I Information on the VAR model

I.1 Aggregate effects of a global supply chain shock

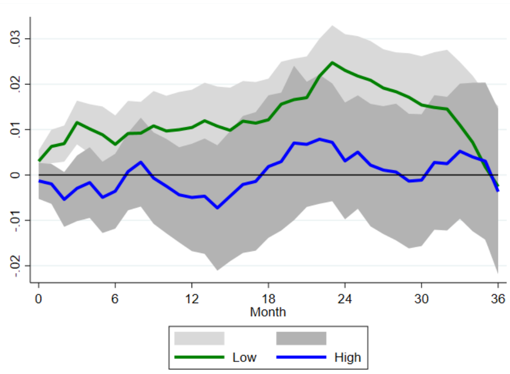
Figure (I.7) shows the aggregate effects of a global supply chain shock (one standard deviation) on the Swedish economy. The red-solid lines correspond to the medians across all models that satisfy the conventional sign restrictions, while the blue-solid lines correspond to the medians across all models that additionally satisfy our narrative restrictions. The dashed lines mark the 90 percent credible bands for the conventional restrictions, while the shaded areas correspond to the 90 percent credible bands that additionally satisfy the narrative restrictions.

On impact, container prices rise by about three percent and continue to rise until they reach a peak after one year, with prices rising by about nine percent. Hereafter, container prices fall below the mean, even though this effect is only borderline significant. We see that the number of containers

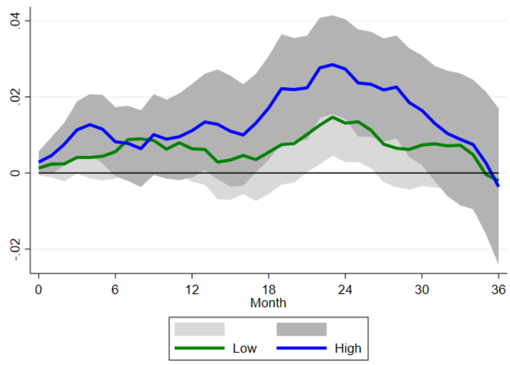
Figure H.6: State-dependent response of product-level producer prices (fixed state variable)



(a) firm size



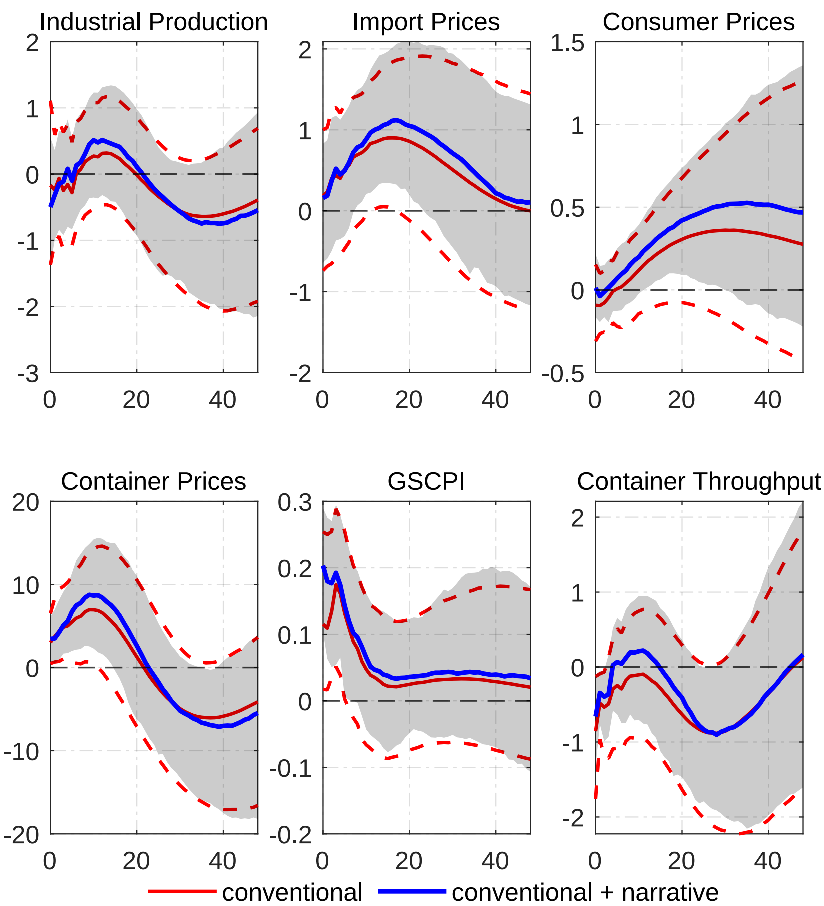
(b) labor costs



(c) market share

Notes: The figure shows the response of firm-level producer prices to a global supply chain shock. The 90% confidence band is based on Driscoll-Kraay standard errors.

Figure I.7: Aggregate effects of a supply chain shock



Notes: The red-solid line corresponds to the median of the draws that satisfy the conventional sign restrictions, while the blue-solid line corresponds to the median of the models that satisfy both the conventional sign restrictions and the narrative restrictions. The dashed lines mark the 90 percent credible bands for the conventional restrictions, while the shaded areas correspond to the 90 percent credible bands that additionally satisfy the narrative restrictions being processed falls by almost one percent on impact. This effect is significant only for the first three periods. Also, we see that the pressure on supply chains jumps on impact. Supply chain pressure decreases very gradually and only returns to its mean after one and a half years.

Importantly, the information from the three narrative restrictions matters: we see that the impact response of the GSCPI is greater in magnitude if we impose narrative restrictions on the model. Also, note that uncertainty around the impulse responses of container prices, the GSCPI and container throughput decreases noticeably once narrative restrictions are imposed.

Turning to the Swedish business cycle variables, a few things stand out. First, although the responses of these three variables remain unrestricted, all impulse responses point in the direction we would expect. This is true for both the draws that only satisfy the traditional sign restrictions and the draws that also satisfy the narrative restrictions. For instance, after a global supply chain shock, industrial production in Sweden falls by almost 0.8 percent after about two years, even though the effect is not significant on the 90% credible interval. The delayed response of industrial activity could be due to the fact that on average firms can still draw on stocks of intermediate products before supply bottlenecks become binding. Consumer prices and import prices, on the other hand, increase by about half a percent and one percent, respectively. Second, all these responses are insignificant when only the conventional sign restrictions are considered. If, on the other hand, the narrative restrictions are also taken into account, we see that the uncertainty of the responses decreases noticeably. This can be observed especially for the reactions of consumer prices and import prices. Hence, we now find that a supply chain shock triggers a significant adjustment of consumer prices and import prices.

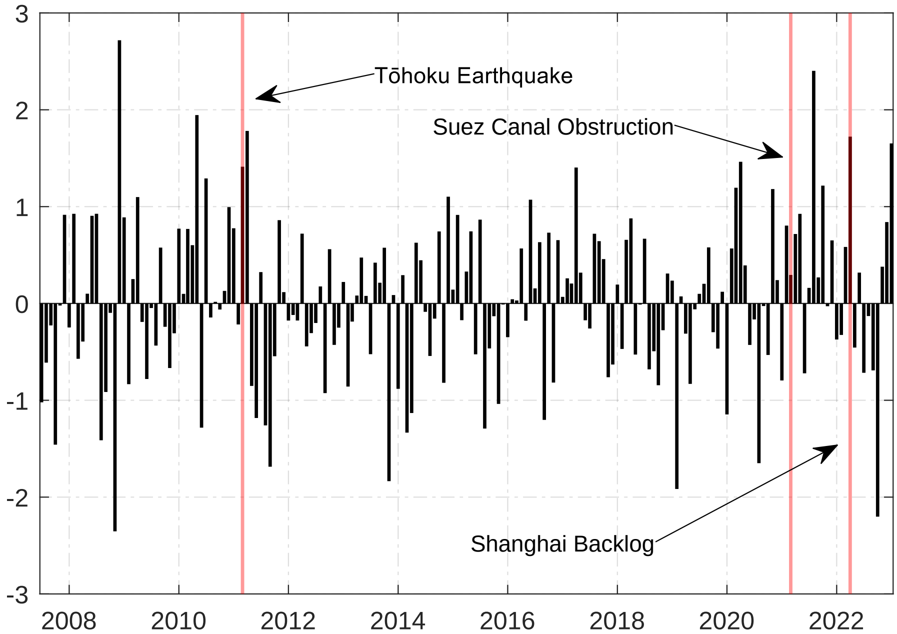
Overall, we find that a supply chain shock has qualitatively comparable effects on consumer prices and industrial production as a conventional supply shock identified by a rise in prices and a fall in production. In our case, however, the reactions of these variables are left unrestricted.[1]

The results are qualitatively comparable with those found in Finck and Tillmann (2023). In their paper, container prices rise by 5% instead of almost 10%, but the effects on consumer prices are only half-sized. The reactions of industrial production and import prices, on the other hand, are roughly the same. De Santis (2024) also uses an SVAR with sign restrictions and narrative restrictions to examine the effects of supply chain disruptions. Here, a shock of one standard deviation leads to an increase in consumer prices of around 0.25% after two years. In another paper, Laumer and Li (2024) use TVP-VARs to measure the effects of GSCPI shocks across different countries. For the euro area, they find that a shock of one standard deviation leads to an increase in inflation of around 0.25 percentage points after about one year.

Figure (I.8) shows the posterior median of the supply chain shock from the VAR model that we use for the estimation of local projections in the main part of the paper.

[1]The Swedish krona has been characterized by a strong depreciation against other currencies since 2014, especially against the US dollar and the euro. In two alternative specifications, we include the broad (narrow) effective exchange rate in our VAR model and re-estimate our model. The identification of the supply chain shock remains the same as in our baseline model, so we leave the exchange rate unrestricted. We find that the exchange rate tends to appreciate following the supply chain shock. Moreover, the impulse responses of the other variables remain almost unchanged, so that the role of the exchange rate via feedback effects is negligible.

Figure I.8: Estimated global supply chain shock over time



Notes: The figure shows the posterior median of the estimated global supply chain shock that satisfies both conventional and narrative sign restrictions. The arrows highlight the three established historical events mentioned in the main text.

I.2 Properties of the supply chain shock

We want to investigate whether our supply chain shock is orthogonal to other global supply-side shocks. For that purpose, we study the contemporaneous correlation with four series of oil supply shocks from the literature. Table (I.3) reports the correlations. The first shock series is taken from Baumeister and Hamilton (2019). From Känzig (2021), we take the series of oil news surprises. As for the Kilian and Murphy (2012) oil supply shock, we update the estimation sample until April 2020 and estimate the model with the same specification (same lag length, priors and identification strategy) as in the

Table I.3: Correlation of the global supply chain shock with other shocks from the literature

Source	ρ	p-Value	n	Sample
Baumeister and Hamilton (2019)	0.01	0.92	183	Jul 2007 - Sep 2022
Känzig (2021)	0.11	0.15	183	Jul 2007 - Sep 2022
Kilian and Murphy (2012)	-0.02	0.82	154	Jul 2007 - Apr 2020
Antolín-Díaz and Rubio-Ramírez (2018)	-0.00	0.98	154	Jul 2007 - Apr 2020

Notes: Correlation of the supply chain shock with selected shocks from the literature, where ρ is the Pearson correlation coefficient. The p-value corresponds to the test whether the correlation is equal to zero and n is the sample size.

original paper. That is, we use the month-on-month growth rate of global crude oil production, an index of real economic activity and the log of the real dollar oil price. We also estimate the Kilian and Murphy (2012) model for the updated sample using the same narrative sign restrictions as in Antolín-Díaz and Rubio-Ramírez (2018), except for the restrictions on the elasticity bounds. In each case, the correlation is very low and indistinguishable from zero.

I.3 Robustness of the VAR identification

The shock series we use in the main part of the paper is drawn from a VAR model identified with sign restrictions and narrative restrictions based on our knowledge of selected episodes. In this section, we address the robustness of our findings with respect to the identification of the shock series. Specifically, we estimate a range of VAR models, each with the full set of sign restrictions but with alternative narrative restrictions, and extract the structural shock series of each model. We then estimate the baseline local projections for each alternative shock series.

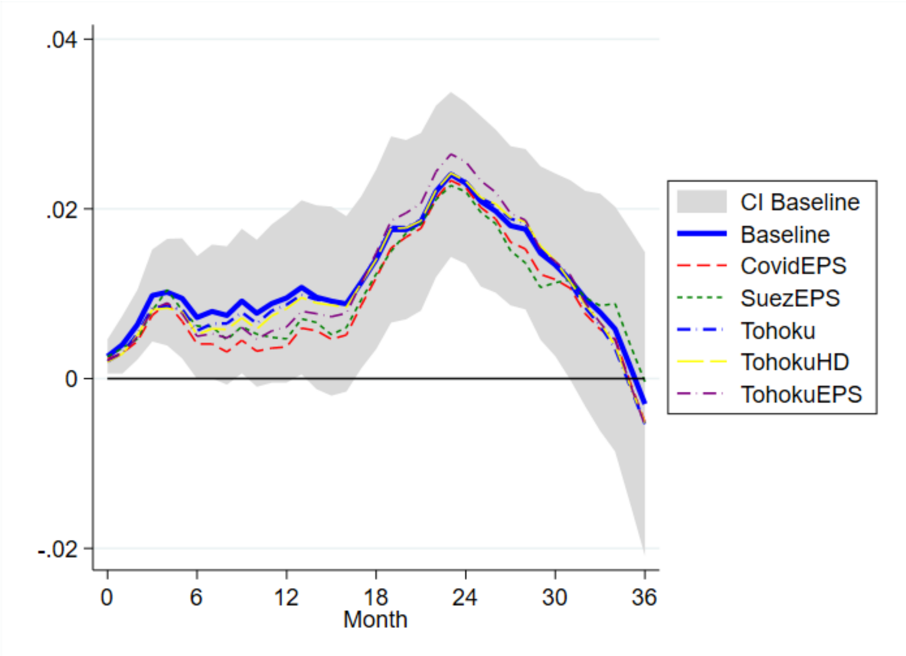
In each alternative model we maintain the restrictions on the signs of the responses to the structural shock. In the first alternative VAR model, we additionally impose only the narrative restriction that the supply chain shock is positive, i.e. restrictive, in March 2011 (TōhokuEPS). In the second model, we impose only the narrative restriction that the supply chain shock is the largest shock in the system in March 2011 (TōhokuHD). In the third model, we impose the previous two narrative restrictions jointly (Tōhoku). In the fourth model, we impose the narrative restriction of a positive shock in March 2021 (SuezEPS). Finally, in the last alternative model we impose that the shock is restrictive in April 2022 (CovidEPS).

Figure (I.9) shows the responses of product-level producer prices to each alternatively identified structural shock as well as the response to our baseline shock. We also include the confidence interval around the baseline estimates. All shock series generate very similar response patterns. In each case, prices increase and reach a maximum 24 months after the initial disruption. Hence, our results remain very robust with respect to the specific narrative identification used in the first stage of the analysis. We decide to include all four narrative restrictions in our baseline model in order to reduce the uncertainty around the responses in the aggregate model.

I.4 Container throughput

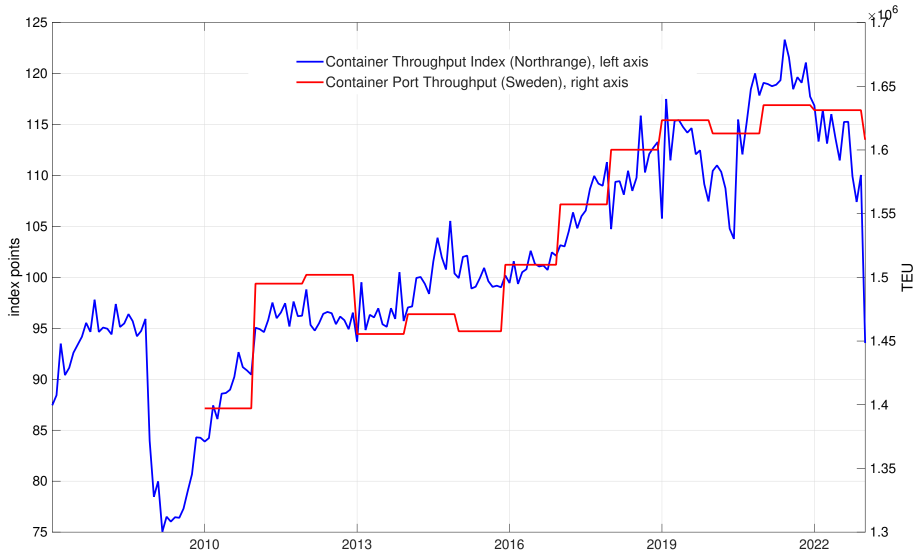
We derive the shock series from an estimated VAR model which includes data on international container trade. In the absence of monthly data on

Figure I.9: Response of product-level producer prices to alternative shocks



Notes: The figure shows the response of firm-level producer prices to a global supply chain shock. The 90% confidence band is based on Driscoll-Kraay standard errors.

Figure I.10: Comparing container throughput in the Northrange and in Sweden



Notes: The figure shows the Container Throughput Index for the Northrange (monthly) and the throughput of containers in Swedish ports (annual). The annual data is taken from UNCTAD and is measured in twenty-foot equivalent unit value (TEU).

Swedish container throughput, we estimate the model using the RWI/ISL container throughput index. This index measures container traffic in the most important North Sea ports. We want to compare this index to data from UNCTAD on container throughput in Swedish ports, which is available on an annual frequency since 2010. Figure (I.10) shows both series, which behave similarly over time.

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