The Spill-back and Spillover Effects of US Monetary Policy: Evidence on an International Cost Channel*

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This Version: September 2025

Abstract

We find that an unanticipated tightening of US monetary policy tends to raise US import prices using granular trade data. This empirical "spill-back" pattern differs from the predictions of typical open-economy macro models. We also document a new empirical "spillover" effect: import prices of other countries also rise following an unexpected US monetary tightening. To understand the mechanism, we examine Chinese exporters and identify a borrowing cost channel — their liquidity conditions generally deteriorate after a US monetary tightening. Indeed, the output price response is greater for those firms facing higher borrowing costs or tighter liquidity conditions.

Keywords: Monetary policy, Spillover, Export prices, Liquidity, Borrowing costs.

JEL Codes: E52, F14, F33, F42.

*We thank Yan Bai, Luigi Bocola, Bo Chen, Kaiji Chen, Maggie Chen, Giancarlo Corsetti, Carsten Eckel, Jingting Fan, Georgios Georgiadis, Robin Kaiji Gong, Yang Jiao, Wolfgang Keller, Ryan Kim, Megumi Kubota, Edwin L.-C. Lai, Eunhee Lee, Andrei Levchenko, Zheng Liu, Dmitry Mukhin, Xiaofei Pan, Ju Hyun Pyun, Veronica Rappoport, Kim Ruhl, Eric Swanson, Daniel Trefler, Tianyu Wang, Frederic Warzynski, Wenbin Wu, Jenny Xu, Stephen Yeaple, James Yetman, Miaojie Yu, Zhihong Yu and seminar and conference participants at Columbia University, Monash University, University of Glasgow, Peking University, HKUST, China Agricultural University, ES World Congress 2025, HKIMR-ECB-BOFIT Joint Conference, NBER Chinese Economy 2024, CICF 2024, 14th CTRG, NBER East Asian 2024, 1st HKUST-Fudan-SMU Conference on International Economics, 6th Melbourne Annual Macro Policy Meeting, Asia Meeting of ES in Vietnam 2024, IAAE Annual Conference 2024, AsianFA Annual Conference 2024, Asia Meeting of ES in China 2024, 1st IEJC in Shenzhen, 15th EITI Conference, and 17th Australasian Trade Workshop for helpful comments. All errors are our own.

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1 Introduction

A tightening of the monetary policy generally puts downward pressure on domestic prices. In a standard open-economy macro model, such policy tightening will lead to a reduction in the demand for imported goods. Alternatively, if the country is on a flexible exchange rate regime, such policy tightening leads to an appreciation of the domestic currency. In both cases, one expects a drop in the home-currency prices of the imports. In models that invoke a "pricing-to-market" assumption, the home-currency price of the imports may not change in the short run. In this paper, we will show that, in the data, an unanticipated tightening of US monetary policy generally leads to a rise rather than a fall in the dollar price of imports, which is different from the predictions of common open-economy macro models.

We also document a new empirical international spillover effect of US monetary policy: an unanticipated tightening of US monetary policy also tends to raise the import prices for other countries. This spillover impact could also be one factor that induces other countries to tighten their monetary policy following a contractionary US monetary policy. While people have observed a positive correlation in interest rates between the US and other countries (e.g. Dedola, Rivolta and Stracca 2017, Miranda-Agrippino and Rey 2020), this particular channel, to our knowledge, has not been documented in the existing literature.

The use of import price indices does not allow us to separate a true increase in the import price of a given product from a possible change in the composition of imports that puts a greater weight on either a higher-priced producer or a higher-valued variety of a given producer. To better understand the empirical "spill-back" and "spillover" effects described above, we study the pricing behavior of Chinese exporters - among the largest set of firms that engage in international trade.[1] With disaggregated data at the level of firm, product, destination market, and month, we confirm that the US dollar prices of their exports to the US generally rise in response to an unanticipated tightening of the US monetary policy. A one-unit unexpected contraction of the US monetary policy (equivalent to a 100 basis-point unexpected increase in the 2-year US treasury yield) raises the annual Chinese export prices to the US in US dollar terms by 21.8%. Moreover, the export prices to other countries also increase by a nearly equivalent magnitude, which constitutes a spillover effect.

Our main measure of unanticipated monetary shocks is based on Bu, Rogers and Wu (2021), which is computed for both the conventional and unconventional monetary policy periods, unpredictable from past information, and is less affected by the central bank information effect.[2] Our results are robust to alternative measures of unexpected changes in US monetary

[1] According to the Smartscout Insight report in 2024, China accounts for 13 of the top 20 cities in terms of the number of Amazon sellers, and 80% of all Amazon sellers from these 20 cities. See `https://www.smartscout.com/reports/top-cities-for-amazon-sellers`.

[2] Regarding the discussion on the information effect of monetary policy, see Nakamura and Steinsson (2018),

policy, including the shocks of Gürkaynak, Sack and Swanson (2005), Nakamura and Steinsson (2018), Jarociński and Karadi (2020), and Bauer and Swanson (2023*b*). Moreover, we perform several additional robustness checks, including (1) different measures of price changes, (2) different sub-samples, and (3) different empirical specifications, with all the results consistent.

We propose an international borrowing cost channel as a plausible explanation for these patterns. Specifically, an unexpected US monetary tightening worsens liquidity conditions for Chinese exporters despite the country's capital controls. This compels the exporting firms to rely more on external financing (e.g., bank loans, which are usually more expensive than internal financing), thus driving up their unit cost, which in turn induces them to raise their export prices. This pattern might be notably prominent for exporters as exporting are usually more demanding than domestic business and relies more on external capital (see Manova 2013). We provide empirical evidence on each part of this logic chain. Moreover, consistent with the cost channel, we show that the impact of the US monetary policy is more prominent on those Chinese exporters with either initially higher borrowing costs or initially tighter liquidity. A simple model with financial frictions can rationalize this mechanism: a US monetary tightening shock prompts firms to borrow more outside funds due to reduced liquidity. The resulting higher borrowing costs induce the firms to raise export prices.[3]

The proposed mechanism is also supported by additional checks. First, we provide suggestive evidence by decomposing the firm-level export price into markup and marginal cost, following the method proposed by De Loecker and Warzynski (2012) and Brooks, Kaboski and Li (2021). We find that the marginal cost responds significantly to the US monetary policy shock, while the reaction of the markup is weak on average. This suggests that a US monetary policy shock is a cost-push shock. Furthermore, the marginal cost shift is driven by financial costs rather than the costs of materials, wages, or imported inputs. As a side note, we examine the possibility that the financing cost effect comes through an increase in the domestic interest rate but find no significant evidence for this.

There are several complementary cross-sectional findings. First, foreign-invested firms are less affected by a tightening US monetary shock, presumably because they are less liquidity constrained. Second, Chinese firms exporting to financially underdeveloped countries exhibit a bigger increase in their export prices. Third, we find that the effect is stronger for the firms engaging in ordinary trade than those in processing trade, probably because ordinary traders are more reliant on external financing. We also investigate three alternative explanations, namely the global demand shift, international competition, and exchange rate pass-through, and find little support for them.

These findings have important implications for understanding the effectiveness of monetary

Jarociński and Karadi (2020), Bauer and Swanson (2023*a*), etc.

[3]Such a conclusion generalizes to models that consider price stickiness and different currency invoicing.

policy. First, the spill-back pattern points to an under-appreciated complication of an open economy for the effectiveness of monetary policy. In particular, a rise in import prices following an unanticipated tightening of US monetary policy makes the job of the US central bank a bit harder in an open economy than in a closed economy. Indeed, it is likely that the more open the US economy becomes, the stronger the spill-back effect will be since the imports will account for a bigger share of the domestic consumption basket. [4]

Second, the new spillover effect shows that an unanticipated tightening of the US monetary policy tends to raise the import prices into other countries, causing a reduction of the real income in these countries and making it harder for them to manage their inflation problem. This provides a (new) rationale for why other countries often synchronize their monetary policies with the United States even when they have a flexible exchange rate regime. The importance of this spill-over effect rises with the level of openness of the foreign economies in either consumption or production.[5]

A crucial feature of our empirical analysis is the use of detailed granular data at the level of exporting firm-product-destination-month/year. This allows us to investigate competing determinants of export prices and helps to better identify the transmission of US monetary policy beyond what one can learn from aggregate import prices alone.

The paper makes three contributions to the literature. First, it provides novel insight into the literature on the transmission of monetary policy, illustrating new evidence on spill-back and spillover effects through import and export prices.[6] Several papers suggest a cost channel - that the marginal cost of (domestic) production could rise with the interest rate - may produce a "perverse" result that a tightening of the monetary policy could lead to an increase in inflation (e.g., Barth III and Ramey 2001, Ravenna and Walsh 2006, Gaiotti and Secchi 2006, Boehl, Goy and Strobel 2022, and Beaudry, Hou and Portier 2024).[7] The "unconventional" sign on domestic CPI inflation only works when the so-called Patman condition is satisfied (Beaudry, Hou and Portier 2024). Our paper reveals an international dimension of the cost channel.[8] In particular, the marginal cost of foreign firms that export to the US may rise in response to an unanticipated US monetary tightening. It is also interesting (and novel) to note that the international cost channel can work even in those foreign countries with a restrictive capital control regime (such as China) through liquidity reduction. Unlike most of

[4]The pass-through of import prices to the U.S. price index has been estimated by Amiti et al. (2022) to have risen in recent years, which would have compounded the problem.

[5]Goldberg and Campa (2010) find that, in a sample of 21 industrialized economies, an important channel for domestic CPI to be affected by import prices is the use of imported inputs in domestic production. This is often more important than the direct effect of imported goods on the domestic consumption basket.

[6]See Bhattarai and Neely (2022) for a survey of the existing literature.

[7]In some VAR studies, monetary easing is followed by falling prices rather than rising ones—a finding referred to as the "price puzzle", e.g., Sims (1992), Bernanke and Mihov (1998), and Wu and Xia (2016).

[8]This is in addition to the dollar exchange rate channel, such as Bruno and Shin (2023).

the existing literature, which focuses on the aggregate macro price index, this paper provides a new perspective on this question using firm/product-level information to differentiate different determinants of price movements.

Second, this paper sheds new light on the determinants of export prices. Complementing the literature that has explored the roles of firm or country characteristics or the effects of trade liberalization[9], we examine the role of US monetary policy in affecting international export prices via an international cost channel. Indeed, our paper appears to be the first that uses granular firm-product-destination data to examine how export prices respond to external monetary policy shocks. A subset of this literature examines the "pricing to market" behavior in which export firms adjust their markups to absorb demand/supply shocks. In comparison, we find a relatively limited role of a markup adjustment and correspondingly a bigger role of a direct effect of a cost shock on export prices.

Third, we offer a new lens on how financial frictions affect international trade prices. While the literature has already shown that credit constraints are important[10], we show that the financial conditions of foreign exporting firms are intimately tied to US monetary policy. Importantly, we show that this linkage is present even in foreign countries with restrictive capital controls.

The remainder of this paper is organized as follows. Section 2 shows some motivating facts. Section 3 describes the data and measurements. Section 4 presents our main empirical results. Section 5 demonstrates the mechanism. Section 6 provides more discussion. Finally, Section 7 concludes.

2 Motivating Facts on Import Prices

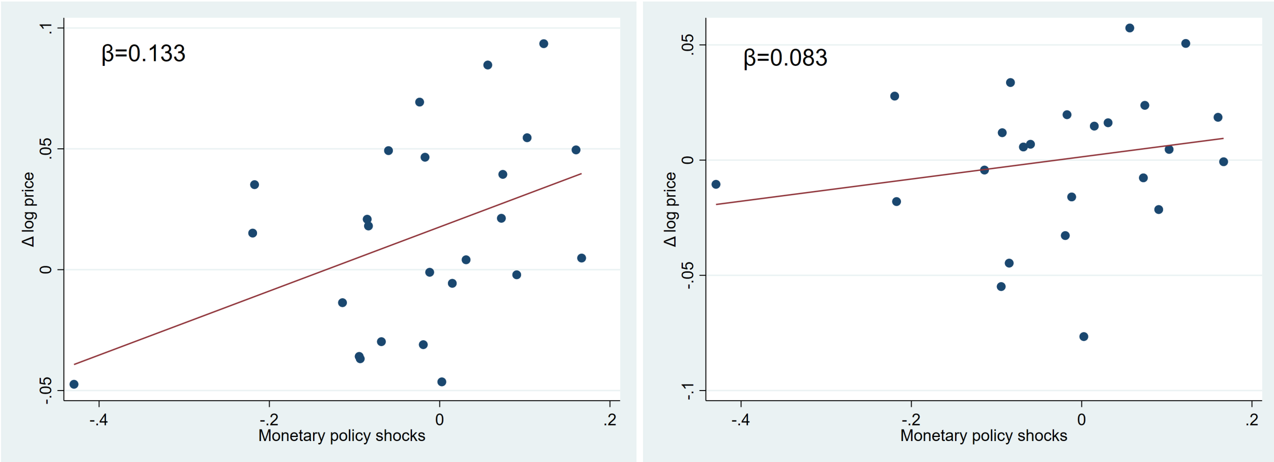
Using the import data from the US Census Bureau (Schott, 2008), we first document the changes in the average product-level annual US import prices in US dollars in response to one unit Fed monetary policy shock, which means an unexpected increase in the daily 2-year US treasury rate by 100 basis points (Bu, Rogers and Wu 2021). These shocks differ from the observed gross treasury yield change, which captures both the expected and unexpected variation. We calculate the US import price of each product as the weighted average unit value at the HS6 product category level from all sources.[11] The correlation of price change and monetary shock is visualized in the scatter plot Figure 1. It is seen that the average

[9]See Obstfeld and Rogoff (2000), Amiti, Itskhoki and Konings (2014), Li, Ma and Xu (2015), Manova and Zhang (2012), Fan, Lai and Li (2015), Harrigan, Ma and Shlychkov (2015), Fan, Li and Yeaple (2015).

[10]See Manova (2013) and Manova, Wei and Zhang (2015).

[11]For each product, the price is its total value over total quantity. Note that the product-level price index is a weighted average of the prices of similar products exported by different firms, which differs from the results of our later baseline regressions that use firm-level price indices.

product-level US import prices seem to rise when the US monetary policy tightens and decline when it loosens. This is inconsistent with the prediction of canonical macro models that the import price would reliably decrease.



(a) Unconditional correlation (without control) (b) Conditional correlation (with controls)

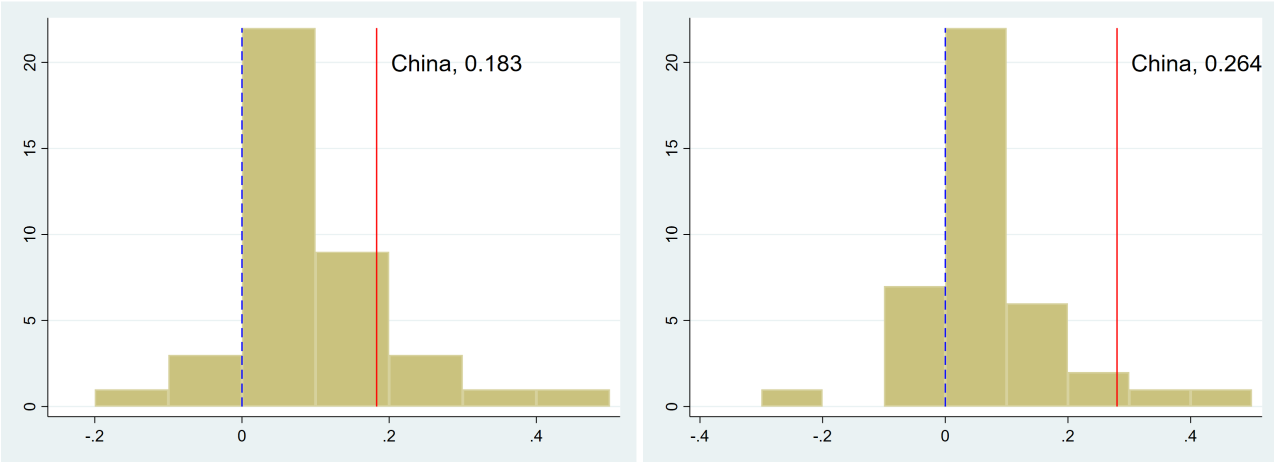
Figure 1: US Monetary Policy Shocks and Average US Import Price

Notes: The horizontal axis represents monetary policy shocks, where one unit of positive shock means an unexpected increase in the daily 2-year US treasury rate by 100 basis points (Bu, Rogers and Wu, 2021). Please refer to Section 3 for details about the shock measure. The vertical axis denotes the simple average of US annual import price changes across all the products. The price of each product is calculated as the unit value at the HS6 product level. Panel (a) shows the unconditional correlation between price change and monetary policy shocks, $\Delta \ln P_{ht}^{imp} = \alpha + \beta m_t + \varepsilon_{ht}$, where P_{ht}^{imp} is the import price for HS6 product h in year t. Panel (b) adds control terms for price changes in the previous year, US GDP growth, and dollar index changes, $\Delta \ln P_{ht}^{imp} = \alpha + \beta m_t + \gamma_1 \Delta \ln GDP_t^{US} + \gamma_2 \Delta \ln DXI_t + \gamma_3 \Delta \ln P_{ht-1}^{imp} + \varepsilon_{ht}$. The vertical axis is the residual of regressing the price change on specified control variables. The sample period is from 1995 to 2019.

We then decompose the responses of the US import prices from each major trading partner and plot the results in Figure 2. For representation, we focus on imports from the top 40 countries (excluding the United States) in terms of nominal GDP in 2006 (US dollar price), including 21 developed countries and 19 developing countries. These trading partners account for more than 95 percent of the total US import value.[12] We may note that import prices from China fall into this majority category, and China's price response coefficient is close to the median among all major trading partners of the US and is therefore representative.[13]

[12] In particular, they are Japan, Germany, China, UK, France, Italy, Canada, Spain, Brazil, Russia, South Korea, Mexico, India, Australia, Netherlands, Turkey, Switzerland, Sweden, Belgium, Indonesia, Saudi Arabia, Norway, Poland, Austria, South Africa, Denmark, Iran, Greece, Argentina, Ireland, Nigeria, United Arab Emirates, Thailand, Finland, Portugal, Venezuela, Malaysia, Pakistan, Colombia, Israel. These countries account for more than 95 percent of total trade with the United States. The combined GDP of these countries accounts for more than 90 percent of the rest of the world's total GDP (other than the United States).

[13] Apart from the import prices in the US, we also find a similar pattern for other countries' import prices



(a) Unconditional price responses of the US from 40 countries (without control)

(b) Conditional price responses of the US from 40 countries (with controls)

Figure 2: US import price responses from 40 major trading partners

Notes: The horizontal axis shows the regression coefficients of US import price (simple average of HS6 product level unit value) changes to US monetary policy shocks (Bu, Rogers and Wu, 2021), from each source among top 40 countries (excluding the United States) in terms of nominal GDP in 2006. The regression is similar to Figure 1. Panel (a) shows the distribution of unconditional US import price responses. Panel (b) adds control terms for price changes in the previous year and the GDP growth of the US. For display purposes, we place all observations less than -0.2 into the leftmost bin and all observations greater than 0.6 into the rightmost bin.

To obtain an intuitive understanding of the importance of the effect through import prices. We can do a back-of-the-envelope calculation. In the US, the share of imports of goods and services in the gross domestic product was 14.7% in 2016 (the last year of the Obama Administration) and 14.9% in 2021 (after 4 years of the Trump tariffs).[14] Once taking into account the input-output linkages through imported intermediate products, Hale et al. (2019) suggests that about 11% of the US consumer spending can be traced to imported goods. As is shown in panel (a) of Figure 1, 1 unit of BRW shock is associated with a 13.3% increase in the annual US import price, which then induces a 1.463% increase in the CPI when domestic prices are unchanged. To be noticed, one unit shock is quite big since one standard deviation of the annual shock is only 0.124 unit. The unexpected interest rate change is also different from the observed gross change. It is worth mentioning that, as is pointed out by Bernanke (2007), the direct effect of higher import prices on the inflation of consumer prices could understate the overall effect if higher import prices motivate domestic firms to increase their prices as well through a competition channel. As for the spillover effect to other countries, this logic is comparable.

(using UN Comtrade data), namely a tightening US monetary shock is associated with an increase in import prices for a lot of countries (see Figure B1). The country list is the same as above including 40 countries.

[14]The data source of GDP and import share is the FRED database.

As the macro-level import price index from a particular partner country is averaged across firms and products, the index variation could be induced by compositional changes in either the producer mix, product mix, or both, instead of changes in the true original prices. Moreover, even without considering the compositional shifts, we still don't know whether a price change is due to markup or marginal cost. If a price change is driven by marginal cost, then which type of cost matters? To account for these concerns and questions, it is necessary to use disaggregated data to investigate the pricing patterns of foreign exporting firms. For this purpose, as an example, we focus on the exporters from China, the largest exporting country of manufacturing goods. Our detailed firm-product-destination-month level data from the Chinese customs records and balance-sheet information of Chinese manufacturing firms allow us to finely distinguish determinants of price changes and identify the transmission mechanism of US monetary policy.

3 Data and Measurement

We draw on three main data sources. First, we use unexpected monetary policy shocks from the U.S. Federal Reserve, and, in an extension, from the European Central Bank. Second, we exploit disaggregated trade data from China's General Administration of Customs. Third, we combine these with firm-level information from the Annual Survey of Industrial Enterprises (ASIE) conducted by the National Bureau of Statistics of China (NBSC). Based on the overlap of these sources, our matched dataset spans January 2000 to December 2006.[15]This section will introduce the basic information about these datasets and briefly describe the sample construction process.

3.1 Monetary policy shocks

Our baseline measure of unanticipated US monetary policy shocks comes from Bu, Rogers and Wu (2021), who use a two-step step similar to Fama and MacBeth (1973) in empirical asset pricing literature: estimating the sensitivity of interest rates at different maturities to FOMC announcements and then regressing all outcome variables onto the corresponding estimated sensitivity index from step one.[16]This measure has several attractive features. First, it is unpredictable from the available information in the past so we can regard it as exogenous

[15]The sample period does not compromise the external validity of our analysis. As shown in the motivating facts, the puzzling price responses to monetary shocks are present over a much longer horizon (1995–2009). The shorter, highly disaggregated sample allows us to investigate the underlying mechanisms with rich firm-level information. We provide further discussion of the impact in other periods and countries in Section 6.

[16]A common approach in the literature is to use monthly or annual changes in the federal funds rate as exogenous U.S. monetary policy shocks, under the assumption that foreign economies—especially smaller ones—do not influence U.S. policy, thereby mitigating reverse causality concerns. However, given China's

for the US import prices (and for import prices into other countries, or the output prices of Chinese exporting firms);[17] Second, no evidence of a significant information effect, so we can treat it as a pure policy shock and avoid the confounding interpretation of the Federal Reserve's private information about the state of the US economy. Third, the measure bridges easily over both conventional and unconventional monetary policy periods. Consequently, we choose this as our baseline shock.

The BRW measure is standardized in such a way that a unit of positive BRW shock corresponds to an unexpected increase in the daily 2-year US treasury yield by 100 basis points. The monthly series can be seen in Figure 3. To match our monthly trade data, we mainly focus on the seven-year period from 2000-2006, which is marked with vertical red lines. Typically, there are eight scheduled FOMC meetings each year, and each meeting has a corresponding policy shock. If there is no FOMC announcement in a month, then the shock in this month is zero.[18]

As robustness checks, we will also use other measures of monetary policy surprises in the literature. Nakamura and Steinsson (2018) use three euro-dollar futures and two federal fund rate futures to extract the first principal component of these price changes within a 30-minute window around the FOMC announcement. The underlying assumption of these shocks is that, in such a tight window around the FOMC statements, most of the asset price changes are driven by monetary policy instead of other factors. Also, if the financial market is efficient, the asset prices before the announcement have already absorbed all the available information; thus, the price changes capture the unexpected component of monetary policy shocks. Using a similar method, Bauer and Swanson (2023*b*) also construct a composite shock, and we use it as a robustness check.

Another measure of Gürkaynak, Sack and Swanson (2005) also uses high-frequency approaches but decomposes the aggregate shock into a target and a path factor, representing the conventional monetary policy and forward guidance, respectively. These shock measures are updated by Acosta (2022). To further alleviate concerns about a possible information effect of the monetary policy, we also employ the "pure monetary policy shock" of Jarociński and

size as both the world's largest exporter and the second-largest economy, it is more difficult to rule out the possibility that U.S. monetary decisions may, at least indirectly, take developments in China into account. Moreover, global shocks can simultaneously affect both U.S. monetary policy and China's exports. In this context, the BRW measure helps alleviate endogeneity concerns.

[17]Recently, Bauer and Swanson (2023*a*) and Bauer and Swanson (2023*b*) suggest that monetary policy shocks identified from 30-minute futures price changes (including the alternative shocks we use later) may be predictable using past economic or financial information, raising concerns about their exogeneity. However, the robustness and interpretation of this predictability remain actively debated (see Acosta 2022 for further discussion). In light of these considerations, we use the BRW shock as our baseline while employing other measures for robustness checks.

[18]This procedure is widely used in the literature, such as Chari, Dilts Stedman and Lundblad (2021).

Karadi (2020), which is identified through the direction of movement in interest rates and stock prices.[19]

When we compare the effects of the European Central Bank and US monetary policies, we use the ECB's pure policy shock and information shock constructed by Jarociński and Karadi (2020), which is methodologically similar to the US counterpart.

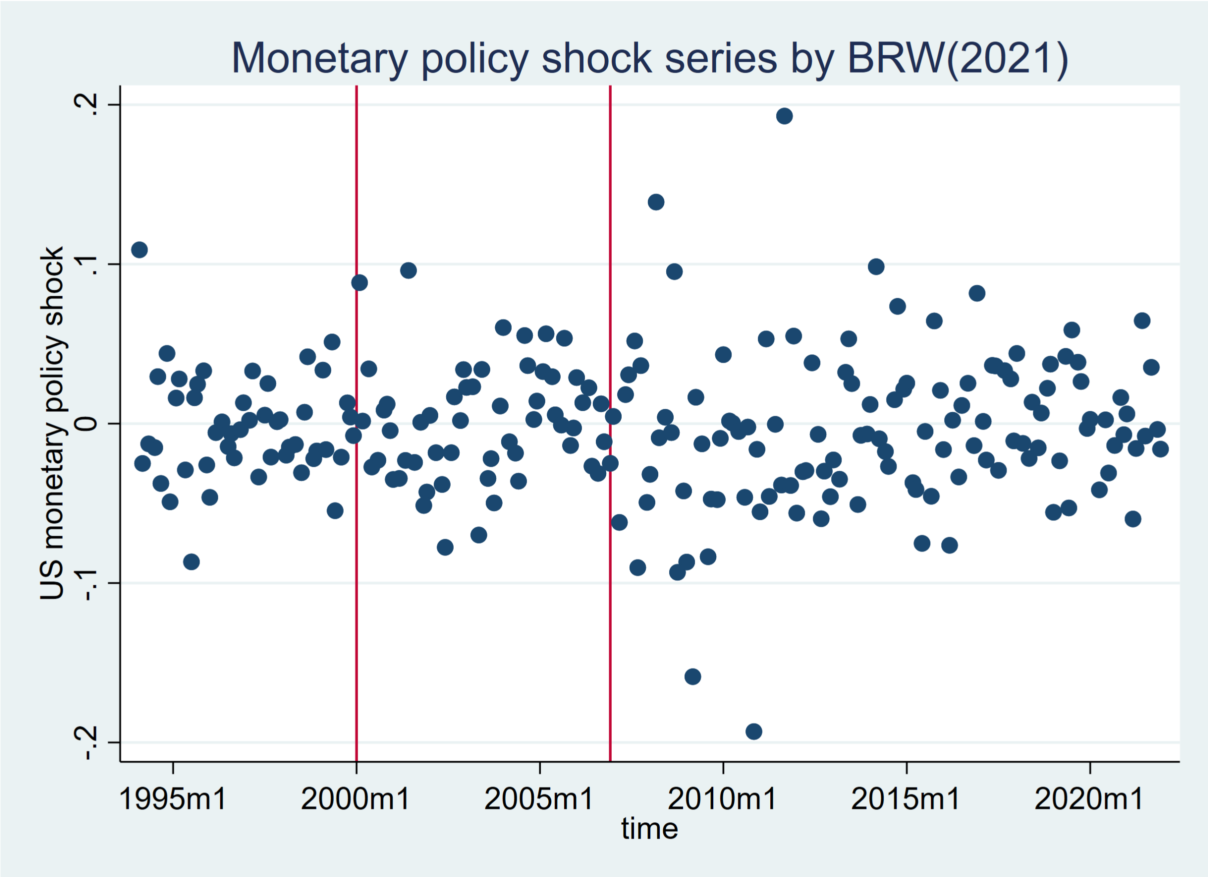


Figure 3: US monetary policy shock: Bu, Rogers and Wu (2021)

Notes: The whole period of Bu, Rogers and Wu (2021) shocks series is from 1994 to 2021. One unit of positive shock means an unexpected increase in the daily 2-year US treasury rate by 100 basis points. This paper will focus on the 84 months from 2000 to 2006, which are marked with vertical red lines. For ease of illustration, in this picture, we drop all months without any FOMC announcement, which means the shocks are zero.

3.2 Customs trade data

To investigate exporters' pricing behavior, we use the monthly transaction records from the General Administration of Customs of China (GACC) from 2000 to 2006 and annual data

[19]The existence of the information effect and its decomposition from pure policy effect is still a hot debate in the literature (see Bauer and Swanson 2023*a* and Acosta 2022). Therefore, we only use this decomposed shock by Jarociński and Karadi (2020) as a double check.

from 2000 to 2007. This dataset includes the most comprehensive information on all Chinese trade transactions, including each firm's import or export value denominated in US dollars, quantity, unit, product name and code, and source or destination country. Our measure of the export price is computed as the unit value (export value divided by quantity), similar to De Loecker and Warzynski (2012).

The product categories are coded according to the Harmonized Coding and Description System (HS) of the World Customs Organization (WCO). HS 8-digit is the most disaggregated classification for Chinese trade. Since there were two major revisions of the HS system in 2002 and 2007, mostly about re-classifying some 8-digit products, we aggregate all HS8 products to the HS6 level and then use the conversion tables from the United Nations Trade Statistics to map them into the older version of HS1996. We exclude (1) products for which we cannot compute unit values due to missing information of either value or quantity, (2) special product categories such as arms (HS2=93), antiques (HS2=97), and other categories (HS2=98 and 99), (3) product-destination level transactions that exist for only one year.

3.3 Chinese firm-level data

Firm-level production and financial information comes from the Annual Surveys of Industrial Enterprises in China (ASIE) conducted by the National Bureau of Statistics of China (NBSC). This database includes all state-owned enterprises and all "above-scale" firms with more than 5 million RMB in annual sales. There were about 160,000 firms in 2000 and around 300,000 in 2007. This data has been widely used (e.g. Khandelwal, Schott and Wei (2013) and Brandt et al. (2017)) since it contains details about firms' identification codes, ownership (e.g., state-owned, private, foreign-invested, and joint ventures), industry type, and about 80 other accounting variables on the three major accounting statements (i.e., balance sheets, profit & loss accounts, and cash flow statements). Among all that information, our research will focus on the variables related to three aspects: (1) firms' production costs and sales, including total wage payment, total operation inputs, and sales income, etc., (2) firms' financial costs, including interest payment and total financial expenses, and (3) liquidity conditions in balance sheets, such as accounts receivable and payable, net liquid assets, and cash holdings.

Manufacturing firms participating in international trade in the matched sample are uniquely identified by their FRDM (legal entity) code and the survey year. To remove potentially problematic observations, such as reporting errors in the ASIE, we follow the criteria of Fan, Lai and Li (2015) and Brooks, Kaboski and Li (2021). In particular, we keep firms that satisfy the following conditions: (1) the firm identification numbers are available and unique; (2) the key financial variables (such as total assets and sales income) are available; (3) the total assets are greater than the liquid assets and total fixed assets; (4) sales are non-negative; (5) total liability is non-negative; (6) number of employees is at least 10.

We merge firm-level survey data with customs trade data based on the firms' contact information as in Fan, Li and Yeaple (2015). In Table A1, we provide summary statistics for firm information and their export patterns in our matched sample. One notable point is that the distribution of firms' export value is skewed to the right, with large exporters accounting for a disproportionate share of the trade value (which is consistent with the empirical trade literature).

3.4 Export price index

The customs dataset contains disaggregated trade values denominated in US dollars and quantities by each firm i, for each HS6 product h, to each country c, at time t, denoted as V_{ihct}, and Q_{ihct}.[20] We compute the unit values as the proxy of export prices:

$$P_{ihct} = \frac{V_{ihct}}{Q_{ihct}}$$

Because product categories are highly subdivided, we believe that the unit value is an ideal proxy for export price.

Using the above unit value price, we construct a firm-level Tornqvist price index using detailed information about the price of each product in each destination market following Smeets and Warzynski (2013).[21] First, we aggregate the unit value to the firm-product level, which is the average price of product h produced by firm i weighted by the relative sales to each market c at time t, $P_{iht} = \sum_c s_{c,iht} P_{ihct}$, where the market-specific value share is $s_{c,ihct} = V_{ihct}/V_{iht}$.

Second, we calculate the firm-product level price growth rate $\Delta_n \ln P_{iht} = \ln P_{iht} - \ln P_{ih(t-n)}$, for all product categories across n periods and then aggregate it to firm level:

$$\Delta_n \ln P_{it} = \sum_h \frac{s_{h,i(t-n)} + s_{h,it}}{2} \Delta_n \ln P_{iht}$$

where the product-specific value share is $s_{h,it} = V_{iht}/V_{it}$, and the effective weight is the average value of product weights at time t and $t-n$. In the regressions with monthly frequency of annual price changes, we set the time gap n as 12 months. This firm-level price growth rate

[20] In robustness checks, we also use the exchange rate between the US dollar and Chinese RMB to convert all trade values into RMB denominations, that is, we compute $P^{RMB}_{ihct} = \frac{V_{ihct} \cdot NER_{US,t}}{Q_{ihct}}$ where $NER_{US,t}$ is the bilateral nominal exchange rate of US dollars in terms of RMB in month t.

[21] We aggregate the price change to the firm level to match with the firm-level channel analysis and to alleviate the data noise originating from multi-dimensions. Moreover, in other specifications, we also exploit product and market differences using more disaggregated data, such as firm-product, or firm-product-destination country level.

$\Delta_{12m} P_{it}$ describes year-over-year changes for the average export price of a certain exporter, considering all adjustments to both product and market scopes.[22] We exclude the top and bottom 1 percent of the year-over-year growth rates of the unit values. For the annual regressions, we first sum up all the values and quantities in a given year and then compute the unit value for that year. The dependent variable would be the change over adjacent years.

4 Impacts of US Monetary Shocks on Chinese Export Prices

This section documents an interesting finding, using granular firm-product level customs data, that unexpected US monetary tightening shocks induce an *increase* in Chinese export prices to both the US (a spill-back effect) and other countries (a spillover effect), which is inconsistent with the prediction of canonical models, either through a demand shrink story or exchange rate appreciation effect. Our results are robust to several checks, including using alternative measures of monetary policy shocks, alternative ways to construct the price changes, different sub-samples and econometric specifications.

4.1 Baseline specification

We first compute changes in firm-level prices by aggregating the export unit values across all product-destination levels for a given firm (see Section 3 for more details). Similar to the literature (e.g. Nakamura and Steinsson, 2018, Chari, Dilts Stedman and Lundblad, 2021, and Gürkaynak, Karasoy-Can and Lee, 2022), we use a local projection method (Jordà, 2005) to study the price responses to monetary policy shocks: [23]

$$\Delta \ln P_{it} = \alpha + \beta \cdot m_t + \Gamma \cdot \mathbf{Z}_{it-n} + \Psi \cdot \Omega_t + \xi_i + \varepsilon_{it} \tag{1}$$

[22] Year-on-year monthly price changes are preferred because they better account for strong seasonality in exporters' activities. Export prices often follow recurring seasonal patterns due to factors such as holidays, harvest cycles, or global shopping seasons. Comparing the same month across years (t vs. $t-12$) holds these seasonal effects constant, while month-on-month changes are more likely to be contaminated by seasonal noise unrelated to monetary shocks. Nevertheless, results using alternative measures (month-on-month, $t-1$ to $t+n$, approximate matching for missing months) are broadly consistent. In Figure B4, we also control for past monetary shocks within the window and obtain robust results. Our annual regressions further confirm that the findings are unaffected by the monthly-inconsecutive data structure and seasonality.

[23] For ease of illustration, we display the concurrent responses in the main text, and the dynamic impacts are shown in the appendix. Apart from this approach, we also employ a dynamic panel GMM with Arellano-Bond estimation method in the robustness part to deal with the possible dynamic panel bias.

where $\Delta \ln P_{it}$ represents the monthly year-over-year export price change of firm i at time t, m_t denotes the unexpected monetary policy shock at time t.

We use the US monetary policy shock from Bu, Rogers and Wu (2021) as our baseline measure, denoted by m_t, and will later use alternative measures for robustness checks.[24] Our parameter of interest, β, represents the average price response to monetary policy surprises. The control variables in $\mathbf{Z}_{it-n}$ are lagged values of price changes in the previous month (to control for possible autocorrelation in price adjustment) and real sales revenue (to control for firm size, n=12).[25] Ω_t represents changes in the RMB/USD exchange rate and potentially other time-varying variables such as China's inflation rate, real growth rate (of industrial production), VIX index, and commodity price movement as additional control variables. To account for unobserved firm heterogeneity, we include ξ_i, the firm-level fixed effects that capture any time-invariant factors for a given firm. We cluster the standard errors at both the firm and time levels to account for possible autocorrelation within a firm and cross-sectional correlation within a period.[26]

We also report annual results, which are more comparable to the yearly data in the motivation facts. In this case, the dependent variable is the annual price change. Following Gürkaynak, Karasoy-Can and Lee (2022) and Di Giovanni and Rogers (2023), monetary shocks are aggregated to an annual frequency. $\mathbf{Z}_{it-n}$ denotes price changes and real sales income in the previous year, and Ω_t represents the annual RMB/USD exchange rate change.

4.2 Baseline results

Table 1 displays the spill-back effect to the US market. In column (1), we only include the monthly monetary shock to estimate the average unconditional effects and find a price increase following the US unexpected tightening.[27] In column (2), we additionally control for the monthly nominal RMB/USD bilateral exchange rate change to account for the confounding effect from exchange rate movement.[28] On average, following an unexpected increase in the interest rate by 10 basis points, the export prices of the Chinese firms to the US market go up by 1.3% (note that a typical magnitude of the surprise change in the monetary policy is less than 10 basis points). The standard deviation of monthly unexpected monetary policy shocks

[24]See Section 3 for more description about the advantages of the baseline shocks and also the details on other measurements.

[25]Larger firms may exhibit more stable prices, given their potential stronger market power and ability to smooth short-term shocks.

[26]Results are robust to other ways of clustering, see robustness part in Section 4.3.

[27]This result echoes the finding in Breitenlechner, Georgiadis and Schumann (2022) that US tightening is associated with domestic inflation in emerging countries.

[28]Note that the coefficient on the exchange rate variable is negative (and near complete), suggesting that a stronger dollar tends to reduce import prices, which is consistent with the existing exchange rate channel.

in our matched sample period is 3.59 basis points, which corresponds to a 0.47% increase in price.

Although the magnitude of US monetary policy shocks does not seem to be quite large, it indeed constitutes one of the main drivers of the global financial cycle (see Miranda-Agrippino and Rey 2020). The associated magnitude of price response is also substantial in our estimation. For reference, the CPI target of the Fed is around 2%, and the average China's export price change (monthly year-on-year) to the US is 4%, suggesting our estimated price responses due to US monetary shock solely account for a substantial proportion of the overall price movements. In column (3), after controlling for other firm-level factors, the magnitude of the price response is moderated slightly to around 10 %, but is still statistically significant.

Table 1: Spill-back - Price responses of Chinese exporters to the US market

	(1)	(2)	(3)	(4)	(5)	(6)
	To the US market					
Dependent Var	Monthly ΔlnP_{it}			Annual ΔlnP_{it}		
brw_t	0.119*	0.130**	0.103*	0.079	0.142***	0.218***
	(0.068)	(0.064)	(0.060)	(0.043)	(0.028)	(0.026)
$Sales_{it-n}$			-0.010***			-0.028***
			(0.003)			(0.006)
ΔlnP_{it-1}			0.299***			-0.348***
			(0.007)			(0.033)
$\Delta lnNER_t^{US}$		-1.080***	-0.888***		-1.026***	-1.791***
		(0.225)	(0.224)		(0.209)	(0.230)
Firm FE	Yes	Yes	Yes	Yes	Yes	Yes
Observations	319773	319773	247028	59695	59695	35639
Number of Firms		40292			46529	

Notes: The dependent variables in columns (1)-(3) are year-over-year changes in monthly prices, while those in columns (4)-(6) are changes in annual prices. All regressions only include exports to the US. Robust standard errors are based on two-way clustering at both the firm and time levels (year-month for monthly regression and year for annual regression). *, **, and *** indicate significance at 10%, 5%, and 1% levels, respectively.

Columns (4)-(6) show the annual results, where we aggregate the monthly changes to the annual level. The dependent variable becomes annual price changes, and the monetary shock represents the sum of all the shocks in that year. It turns out that the elasticity of the export price change to the monetary shock is somewhat bigger than the corresponding one in monthly regressions. For example, based on the estimates in column (6), an unexpected tightening of US monetary policy by 10 basis points raises Chinese firms' export prices to the US by 2.18%.[29]

Table 2: Spillover - Price responses of Chinese exporters to non-US markets and all countries

Panel A: non-US markets	(1)	(2)	(3)	(4)	(5)	(6)
Dependent Var		Monthly ΔlnP_{it}			Annual ΔlnP_{it}	
brw_t	0.177**	0.184**	0.151**	0.136**	0.175***	0.248***
	(0.076)	(0.077)	(0.068)	(0.041)	(0.040)	(0.053)
$Sales_{it-n}$			-0.005*			-0.015**
			(0.005)			(0.003)
ΔlnP_{it-1}			0.299***			-0.303***
			(0.030)			(0.006)
$\Delta lnNER_t^{US}$		-0.777***	-0.654***		-0.717**	-1.119**
		(0.260)	(0.297)		(0.195)	(0.201)
Observations	1016974	1016974	834168	146735	146735	92927
Number of Firms		75523			86882	
Panel B: all countries	(1)	(2)	(3)	(4)	(5)	(6)
Dependent Var		Monthly ΔlnP_{it}			Annual ΔlnP_{it}	
brw_t	0.173**	0.180**	0.150**	0.136**	0.177***	0.244***
	(0.075)	(0.075)	(0.064)	(0.041)	(0.038)	(0.048)
$Sales_{it-n}$			-0.003			-0.014*
			(0.006)			(0.003)
ΔlnP_{it-1}			0.281***			-0.313***
			(0.031)			(0.006)
$\Delta lnNER_t^{US}$		-0.740***	-0.629***		-0.688**	-1.103**
		(0.269)	(0.312)		(0.201)	(0.208)
Observations	1100400	1100400	917419	151542	151542	96296
Number of Firms		76811			88425	
Firm FE	Yes	Yes	Yes	Yes	Yes	Yes

Notes: The dependent variables in columns (1)-(3) are changes in monthly prices, while columns (4)-(6) are changes in annual prices. Panels A and B include exports to the non-US markets and all countries, respectively. Robust standard errors are based on two-way clustering at both the firm and time levels (year-month for monthly regression and year for annual regression). *, **, and *** indicate significance at 10%, 5%, and 1% levels, respectively.

In panel A of Table 2, we show the responses of Chinese firms' export prices in non-US markets, namely the spillover effects. The slope coefficients are both positive and significant

[29]The coefficients of lagged price change are opposite in the two versions: price changes display inertia in the short term and show a pattern of mean reverting in a relatively longer run.

in both the monthly and annual regressions. The point estimates are moderately bigger than those for the export prices to the US market. In panel B, we examine the export prices to all markets (both US and non-US) and find similar results.

To check for possible roles of outliers, we present a scatter plot of monthly export price changes, averaged across Chinese firms, on the vertical axis, against US monetary policy shocks on the horizontal axis in Figure B2. The graph suggests that our conclusion is unlikely to be driven by a small number of outliers. As another way to validate our results, we conduct separate regressions of Chinese export prices in each of the country's top 20 trading partners. A histogram of the 20 slope coefficients is plotted in Figure B3. In an overwhelming majority of the cases (18 out of 20), the export price response to an unexpected US tightening shock is positive.

These findings are noteworthy as they differ from the prediction of the standard open economy macroeconomic models, where a tightening monetary shock should decrease global import demand due to an expenditure decline effect and hence reduce export prices. Furthermore, the exchange rate depreciation effect also contributes to a decrease in the price of the US dollar.

It is also worth mentioning that at the macro level, using time series models such as VAR, people usually observe an ambiguous impact of monetary shocks on price indexes, either the CPI price index or the import price index. Sometimes monetary tightening leads to a price increase (a price puzzle, e.g., Sims 1992, Bernanke and Mihov 1998, Wu and Xia 2016, and Breitenlechner, Georgiadis and Schumann 2022), while other times a decrease (consistent with canonical models' prediction). These results are quite sensitive to the method, such as OLS VAR, Bayesian VAR, to identification assumptions, to instrumental variables, and to specifications, like lag orders and inclusion of different endogenous variables.[30] Instead of taking a stance on which macro estimation is right, we use micro-level prices to offer a new perspective to this question.

We display that the average impact in the current period is significant, which does not mean that prices are not sticky (as Gopinath et al. 2020). As long as a proportion of firms can adjust, the average prices will respond to shocks, and less sticky firms may adjust more.[31]

[30]See Rudebusch (1998) and Miranda-Agrippino and Ricco (2021) for more discussion on the limitations of those methodology issues. We also verified this point using different methods and specifications with the aggregate price index and yield mixed results, which are available upon request.

[31]Regarding stickiness, there are two popular theories: (1) only firms with a low menu cost could reset price immediately (e.g. Golosov and Lucas Jr (2007)); (2) each firm has a probability to change the price in a given period (e.g. Calvo (1983)). Both suggest that with the existence of stickiness, some firms can adjust prices to the current shock. This is different from completely rigid prices where all the firms could not change prices. Nakamura and Steinsson (2008) show that the US firms have a probability of around 20% to change the price in a month on average and this frequency could range from 7% to 87% across different products. Moreover, according to Zhang (2022), homogeneous goods are usually more flexible in price adjustment than

Likewise, we do not rule out that other firms may adjust only with a lag (e.g., due to contract renewal). In addition to the concurrent reaction, regarding dynamic responses, we find that this impact is more prominent in the first 6 months, and it will gradually fade out after that, indicating a short-term effect. The detailed results are shown in Figure B4.

In Table B1, we analyze the impact of the US monetary policy shocks on the export values and quantities separately. In the annual regression, a tightening shock will reduce export quantity. This effect dominates the export price response since we see a reduction in the overall export value. This is consistent with Lin and Ye (2018*b*), who find that US tightening will reduce global trade volumes using cross-country industry-level data. In the monthly regression, the responses are insignificant. This is plausible, as quantity adjustment in the short run might be more sticky, perhaps due to relatively high capital adjustment costs.

4.3 Robustness checks

4.3.1 Alternative measures of monetary policy shocks

We explore the robustness of our results to using alternative measures of monetary policy shocks in Table 3. First, we use the composite shock from Nakamura and Steinsson (2018), which is derived from 30-minute high-frequency data on the price changes of 2 federal fund rate futures and 3 Eurodollar futures around the FOMC announcements. This unitary measure combines the information in forward guidance as well as the conventional monetary policy shock. As for our baseline BRW measure, this measure is rescaled so that a one-unit increase is equivalent to an increase in the 2-year bond yield by 100 basis points. (We will rescale all other alternative measures to be discussed below in the same way.) The regression results with this alternative measure of monetary policy shock are reported in columns 1 and 2 in Panel A (monthly change in export prices) and Panel B (annual changes in export prices) in Table 3. In all specifications, the export price responds positively to unanticipated monetary policy tightening.[32] Based on the estimate in column 2, Panel B, the Chinese export prices go up by 2.27% following a surprise tightening of US monetary policy by 10 basis points.

Our second alternative measure is a composite shock from Bauer and Swanson (2023*b*). From columns 3-4 in Table 3, we see similar results both qualitatively and quantitatively. That is, the export prices respond positively to a surprise tightening of US monetary policy, with an elasticity that is similar to those estimated by other measures of monetary policy shocks. Our third alternative, based on the idea of Gürkaynak, Sack and Swanson (2005) but extended

differentiated goods. As is displayed in Table C9, we find that the former group will increase the prices by a larger magnitude facing a tightening US monetary shock.

[32] Results are consistent without the inclusion of exchange rate change. This also applies to other robustness checks.

by Acosta (2022), is a pair of "target" measures akin to a conventional monetary shock and a "path" from forward guidance, respectively. From the regression results in columns 5-6 in Panels A and B, we see that the positive export price response to the monetary policy shock is mostly driven by the surprises in the "path" or "forward guidance" component.

The measures so far may reflect the private information that the Federal Reserve may have regarding the current and future economic fundamentals (see Nakamura and Steinsson 2018, Jarociński and Karadi 2020, Acosta 2022, etc.). In comparison, our fourth alternative measure, based on Jarociński and Karadi (2020), is supposed to have excluded the information effect. Since BRW also describes their measure as having purged the information effect, this alternative is more comparable to our baseline measure in this sense. On the other hand, this measure has two components of monetary policy shock, similar to Gürkaynak, Sack and Swanson (2005). From the last two columns in Table 3, we see that a surprise tightening in component representing pure monetary policy shock (MP_t^{JK}) also triggers an increase in the export prices.[33]

A different robustness check with regard to monetary policy shock is to account for the location of a FOMC announcement in a particular month or quarter. Similar to Ottonello and Winberry (2020), we construct a date-weighted shock (based on BRW) according to the number of days remaining in the month after an FOMC event.[34] We calculate the weighted shock in annual frequency in an analogous way. In particular, the shock in a given year is the sum of all the shock components in that year, starting with the remaining part of the last shock in the previous year, and ending with the effective part of the last shock of this year. As can be seen in Table B2, our results are unaffected by this adjustment.

To summarize, even though different measures of unanticipated monetary policy shocks may capture different aspects of the monetary policy,[35] and even though their differences may matter in other context, our key conclusion is not sensitive to which exact measure is used.

[33] There is a caveat to these results with alternative shocks. Recent studies by Bauer and Swanson (2023*a*) and Bauer and Swanson (2023*b*) challenge the exogeneity of monetary policy shocks identified from 30-minute futures price changes, arguing that these shocks are predictable based on prior economic and financial information. However, the robustness and interpretation of this predictability remain debated (see Acosta 2022). Besides, the existence of the information effect and its decomposition from pure policy effect is also under hot debate in the literature (see Bauer and Swanson 2023*a*). We address these concerns by using the BRW shock as our baseline measure and conducting robustness checks with alternative shocks.

[34] For example, if a FOMC announcement was made on March 20, 2001, and the magnitude of the monetary shock was 1, we will attribute $(31-20)/31$ of this shock to the current month and the remaining part to the next month. Therefore, the adjusted shock measure in a given month is equal to the remaining part of the shock in the last month plus the effective part of the next shock in this month.

[35] The correlation of BRW shock with the NS shock, BS shock, target shock, path shock, MP shock, and CBI shock are 0.54, 0.49, 0.28, 0.47, 0.42, and 0.15, respectively, as reported in Table A2.

Table 3: Alternative US monetary policy shocks

Panel A: monthly	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Dependent Var				Monthly $\Delta \ln P_{it}$				
NS_t	0.111***	0.105***						
	(0.041)	(0.037)						
BS_t			0.130***	0.126***				
			(0.044)	(0.043)				
$Target_t^{Acosta}$					0.047	0.044		
					(0.035)	(0.028)		
$Path_t^{Acosta}$					0.101***	0.097***		
					(0.037)	(0.034)		
MP_t^{JK}							0.062	0.068
							(0.039)	(0.047)
CBI_t^{JK}							0.137***	0.094*
							(0.047)	(0.049)
Observations	1100400	917419	1100400	917419	1100400	917419	1100400	917419
Panel B: annual	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Dependent Var				Annual $\Delta \ln P_{it}$				
NS_t	0.126***	0.227***						
	(0.023)	(0.014)						
BS_t			0.092**	0.320***				
			(0.026)	(0.019)				
$Target_t^{Acosta}$					0.044	0.077***		
					(0.030)	(0.012)		
$Path_t^{Acosta}$					0.115***	0.221***		
					(0.021)	(0.012)		
MP_t^{JK}							0.029***	0.245**
							(0.005)	(0.083)
CBI_t^{JK}							0.099***	0.148***
							(0.006)	(0.009)
Observations	151542	96296	151542	96296	151542	96296	151542	96296
Firm Controls	No	Yes	No	Yes	No	Yes	No	Yes
NER Control	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Firm FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes

Notes: In panels A and B, the dependent variables are changes in monthly and annual prices, respectively. The monetary policy shocks in columns (1)-(2), (3)-(4), (5)-(6) and (7)-(8) are from Nakamura and Steinsson (2018), Bauer and Swanson (2023*b*), Acosta (2022), and Jarociński and Karadi (2020), respectively. For ease of comparison, all shocks are rescaled so that a one-unit increase is equivalent to a rise in the daily 2-year US treasury yield by 100 basis points. Robust standard errors with two-way clustering at both the firm and time levels (year-month for monthly regression and year for annual regression) are reported. *, **, and *** indicate significance at 10%, 5%, and 1% levels, respectively.

4.3.2 Alternative ways to construct the dependent variable

We investigate the sensitivity of the results to how changes in export prices are constructed. Rather than computing firm-level export prices as we do in the baseline analysis, we examine export prices at the more disaggregated firm-product level (as reported in columns 1-3 in Table B3), or at even more disaggregated firm-product-destination level (in columns 4-6 of the same table). The more disaggregated levels allow us to add many additional controls. In particular, we include exporting firm-product fixed effects in columns 1-3, and firm-product-export market fixed effects in columns 4-6. These additional fixed effects should, in principle, absorb many other potential determinants of the export price changes not present in our baseline regressions.

In addition, in columns 4-6, we can account for changes in bilateral exchange rates between the Chinese RMB and the currency of the destination country. We also include CPI inflation and real GDP growth of the export destination countries. (In comparison, in columns 1-3, we include the changes in the RMB/dollar exchange rate.) In columns 2 and 5, we include lagged sales growth by firm-product and firm-product-market, respectively. In columns 3 and 6, we additionally include the corresponding lagged dependent variable. Across the 12 specifications reported in Table B3, the coefficients on unanticipated monetary policy shock are positive in all cases, and statistically significant in 11 out of 12 cases. Using the estimates in columns 3 and 6 of Panel B as illustration, following a 10 basis points surprise tightening of the US monetary policy, the export prices by Chinese firms tend to go up by 2.5% averaged at the firm-product level, and 2.0% averaged at the firm-product-market level, respectively.

At the disaggregated level, it is common for a given firm not to export every month as the shipments tend to be bunched. This is certainly the case at the firm-product or firm-product-market level, but it can happen at the firm level too. This means that the dependent variable in our baseline estimation has many missing values if a given firm does not export anything exactly 12 months earlier. As a robustness check, if no export value is available exactly 12 months earlier, we fill it up with the value from the most adjacent month from the 13th to the 11th (or from 14th to 10th) month before. While the number of observations has increased as a result, the regression coefficients are essentially the same as in the baseline regressions (see Table B4).

As another robustness check, we define the dependent variable as the price change last export price of the previous calendar year to the last export price of this calendar year. The corresponding monetary shock is the sum of all monthly shocks in that year. While this leads to a smaller regression sample, the key coefficients, as reported in Table B5, are very similar to the baseline regressions.

Finally, while the export prices are all denominated in US dollars in the baseline regression, we convert them into Chinese RMB as a robustness check. This could make a difference if the currency invoicing behavior differs across destination markets. While our data does not provide information on the exact currency invoicing, anecdotal evidence suggests strongly that Chinese exports were overwhelmingly denominated in US dollars, regardless of the ultimate export markets. Nevertheless, our results barely change using RMB price as reported in Table B6.

4.3.3 Different sub-samples and econometric specifications

As many firms export multiple products, a change in firm-level export price could also reflect a change in the mix of products. To purge possible product-switching, we repeat our baseline regressions using a reduced sample of single-product (HS-6 digit) firms and report the results in Table B7. Although the new sample size is much smaller, we observe similar price responses to monetary policy shocks as in the baseline regressions.

We examine export price behavior separately for firms of different ownerships: state-owned enterprises (SOE), domestic private enterprises (DPE), multinational enterprises (MNE), and sino-foreign joint venture enterprises (JV). From Table B8, we see that the export prices respond positively to surprised monetary policy tightening for firms of all ownership types, with the responsiveness somewhat bigger for majority state-owned firms and domestic privately owned firms.

Some exporting firms use imported inputs in their production (“two-way traders), while others use only domestic inputs (“pure exporters). We examine how these two types of firms exhibit different responses to US monetary policy shocks. From Table B9, we see that both types of exporters raise their prices facing surprised tightening shock. While the point estimate is somewhat bigger for the pure exporters in monthly and yearly regressions, the difference with the two-way traders is small.

We explore alternative fixed effects and standard error clustering. In Table B10, we include additional year or month fixed effects in columns 1-4, Besides, we also apply alternative clusters of standard errors (firm-level, time-level, or sector-level) to panel regressions. All the results are robust and significant.

We experiment with adding more macroeconomic indicators as additional control variables, such as China’s industrial production growth rate and inflation rate, VIX index, and commodity prices. This should help to account for additional time-varying factors on export prices. As shown in Table B11, they do not alter the key conclusion that the Chinese export prices respond positively to unanticipated US monetary tightening.

While our baseline estimation uses local projection specification (as in Jordà 2005), we also utilize the dynamic panel GMM with Arellano-Bond estimation method to account for the possible dynamic panel bias. Both the difference GMM and the system GMM estimation, reported in Table B12, yield a positive coefficient on the monetary policy variable.[36]

5 International Borrowing Cost Channel

In this section, we explore the mechanism behind our findings. In a standard open economy macroeconomic model, a tightening of the US monetary policy will reduce the demand for imported goods due to an expenditure reduction effect. This tightening will also appreciate US dollar. Both forces would imply a decrease in import prices. Since the actual change in the US import prices go in the opposite direction from this prediction, we need to search for a different explanation.

We propose an "**International Cost Channel**" to explain the empirical patterns. In particular, an unexpected US monetary tightening worsens liquidity conditions for exporting firms in other countries (including the Chinese exporters), causing them to pass the higher cost of capital to their output prices, including the export prices to the US and other markets.

For foreign countries with no capital controls, it is easy to see that a contractionary US monetary shock leads to financial tightening in these foreign countries (see Miranda-Agrippino and Rey 2020). For countries with capital controls (such as China), a US tightening can also generate a deterioration of liquidity conditions for the exporting firms in these countries. For example, as their foreign sales decline, their relatively cheap internal financing may dwindle, and they may also find it more difficult to access trade credit from their international upstream or downstream client firms. In both cases, monetary tightening may motivate them to turn to the more expensive types of external financing, such as borrowing from banks. In response to a higher cost of capital, these firms raise their export prices to compensate.

5.1 Conceptual Framework

We construct a simple partial equilibrium model to show how exporting firms' pricing responds to a US monetary contraction.[37] The model is based on the heterogeneous firm trade model of Melitz (2003) and Manova (2013). We further incorporate the US monetary shocks, and a working capital constraint like Ravenna and Walsh (2006). The model's logic is that a tightening US monetary shock could motivate firms to borrow more from external financial institutions due to shrinking liquidity, thus increasing the borrowing cost and the corresponding export price.

[36] However, this approach assumes that the error terms are not serially correlated, which can be rejected in the data. Consequently, we use the local projection results as our baseline.

[37] As we focus only on the firm's behavior, we use a partial equilibrium setting for simplicity. A general equilibrium model will unlikely alter the impact of the US monetary policy shocks on exporter pricing.

The settings of consumer and preference are standard, which are shown in the Appendix D.1. We focus on exporting firms. There is a continuum of firms that ex-ante differ in their productivity level ϕ_i. We assume that there is only one input (e.g., materials or labor) for production. The production function is $y_i = \phi_i L_i$, where ϕ_i is productivity and L_i is input with a unit price of w_i. The firm in country i minimizes its cost to satisfy the demand in the country j, $Y_{ij}(\omega) = \frac{p_{ij}(\omega)^{-\sigma}}{P_j^{-\sigma}} Y_j$.

We assume a working capital constraint — a fraction δ_i of the input costs is borrowed from outside financial institutions (e.g., bank loans or issuing bonds) and paid in advance with a gross interest rate R_i.[38] Here $\delta_i \in [0, 1]$ is a decreasing function of the firm's liquidity condition: $\delta_i \equiv 1 - c_i^\gamma$, where $c_i \in [0, 1]$ is the liquidity condition (a higher value means better situation), γ is a positive constant and reflects the elasticity of borrowing fraction with respect to liquidity condition.[39] Intuitively, a firm with better liquidity conditions is assumed to have fewer external financing needs. Solving the unconstrained optimization problem will give us the optimal price (for simplicity, we suppress the subscripts):[40]

$$p = \frac{\sigma}{\sigma - 1} \frac{\tau w[c^\gamma + (1 - c^\gamma) R^\alpha]}{\phi} \tag{2}$$

From Equation 2, we can see that a monetary tightening could potentially affect the price through input price w, borrowing rate R, and liquidity condition c.[41] A US monetary tightening can raise the cost of funding for the firm and, by extension, its output price in two channels. If the interest rate R in the exporting firm's country increases, then the cost of funding will go up. Alternatively, if a deterioration of the liquidity condition causes the firm to turn to a higher proportion of external borrowing, $1 - c_i^\gamma$, which is more expensive than the firm's internal financing ($R > 1$), its cost of funding also goes up. Either way, the firm responds to the rise in cost of capital by raising its output (and export) prices.

This effect is non-linear. In particular, the output price response varies with the initial liquidity condition and borrowing cost of a firm. We summarize these results in the form of propositions in Appendix D.3 and D.4.

In the baseline model, we assume a single input factor, no credit constraints, flexible

[38] Compared with domestic sales or purchasing, the cross-border trade is riskier and more demanding, and relying more on external capital (Manova 2013).

[39] For simplicity, the liquidity conditions are assumed to be exogenously determined, and the endogenization of liquidity will not alter our main mechanisms. This simplification can help us to get an analytical solution to the pricing equation, which will show the mechanism clearly. Moreover, under this setting, our model can be easily extended to incorporate two input factors, borrowing constraints, dynamic and sticky pricing, and different invoicing currencies.

[40] More details are in Appendix D.2.

[41] Here $\tau >= 1$ is a constant iceberg cost, namely how many units of goods must be shipped for one unit to arrive at a destination.

and static pricing, and abstract from exchange rate and currency invoicing considerations. Importantly, our channel remains robust when these assumptions are relaxed (section 6.4).

Empirical strategy. To empirically test this mechanism, we will show (1) that the US contractionary monetary shock would worsen the liquidity conditions of Chinese exporters; (2) that the average borrowing costs and borrowing proportions increase in response to a tightening shock; (3) that firms with higher borrowing costs, or tighter liquidity conditions would increase their prices by a larger magnitude.

Additionally, we will demonstrate that the price increase is most driven by a change in the marginal cost and less by any change in the markup. Finally, in the sample of Chinese exporters, the borrowing cost reaction is mainly due to the increase in borrowing proportion instead of the interest rate itself. This lends support to the idea that the international cost channel works even in countries with restrictive capital controls.

5.2 Liquidity responses to monetary shocks

We check the responses of firms' liquidity conditions to foreign monetary shocks. Intuitively, an unanticipated US monetary tightening, by reducing the world demand, will reduce a Chinese exporter's exporting revenue and reduce the trade credit it otherwise may be able to access from its international upstream or downstream firms. To measure a firm's liquidity condition, we use two variables: the cash ratio *Cash* (cash holding over total assets) and the net liquid asset ratio *Liquid* (liquid asset minus current liability over total assets) following Manova and Yu (2016) and Dai et al. (2021).[42] A lower value in either variable corresponds to a tighter liquidity condition. We regress the change in a firm's annual liquidity condition on US monetary shocks:

$$\Delta Liq_{it} = \alpha + \beta \cdot m_t + \Gamma \cdot \mathbf{Z}_{it-1} + \xi_i + \varepsilon_{it} \tag{3}$$

where Liq_{it} represents a measure of liquidity condition ($\Delta Liq_{it} = Liq_{it} - Liq_{it-1}$), and Z_{it-1} is a group of firm-specific one-year lagged control variables, including log real sales income (a proxy for firm size) and the ratio of total debt to total assets. ξ_i is firm fixed effect. Following Lin and Ye (2018*a*), who also use the ASIE database to study the effects of US monetary policy on China's trade credits, we cluster the standard errors at the firm level to account for the possible correlation of error terms for a given firm.

From columns (1) and (2) of Table 4, we find that an unexpected US tightening shock significantly worsens the liquidity of firms. One possible explanation is that the contractionary

[42] All the balance sheet variables are obtained from the Annual Surveys of Industrial Enterprises in China (ASIE). Without special notice, the period of all other variables in this database is from 2000 to 2007.

US shock will reduce global demand and thus reduce the firm's operational cash flow from exporting. This is consistent with our previous finding that the US tightening shock will reduce the firm's annual export quantities and values (see Table B1). In addition to *direct* measures of exporters' liquidity conditions, we also use trade credit changes as *indirect* measures to account for firms' liquidity condition changes in columns (3)-(4). We find that the overall trade credit of Chinese exporters decreases after a tightening US shock, which could further worsen the firms' liquidity.

As is documented in the literature, the US monetary policy is a main driver of the global financial cycle, and a tightening shock would cause a decline of global asset prices, an increase in financing costs, and a contraction of loans (see Miranda-Agrippino and Rey 2020 and Di Giovanni et al. 2022). International firms or local firms borrowing abroad may reduce trade credit provisions to their trade partners to mitigate financial pressure (Alfaro, García-Santana and Moral-Benito 2021, Ding et al. 2024, Ersahin, Giannetti and Huang 2024). As a result, even those Chinese exporters that do not borrow directly from international banks or the capital market can be affected indirectly if their trading partners (upstream or downstream firms) are affected by the global liquidity retrenchment. [43]

To verify this argument, we first check how Chinese exporters' trade credit acceptance (accounts payable) responds to the US monetary shocks. Similar to Fisman and Love (2003), trade credit acceptance is measured by the ratio of accounts payable to total assets ($APay$).[44] The specification is the same as Equation 3. The results are displayed in column (3) of Table 4 that an exporter in China will get fewer trade credit provisions from its trade partners after a tightening shock. Moreover, through the responses of accounts receivable over total assets in column (4), we see that exporters also reduce trade credit provisions to other firms due to liquidity contraction. This finding is consistent with the result of Lin and Ye (2018*a*) that foreign liquidity shortage will cause the trade credit provisions of Chinese firms, especially FDI firms, to shrink. It is also in line with Love, Preve and Sarria-Allende (2007), showing trade credit collapsed in the aftermath of the 1997 Asian crisis, and Adelino et al. (2023) finding that an easing European Central Bank monetary policy increases the trade credit provisions of core countries to periphery countries.[45]

[43] Note that the effect of international connections can go both ways. When international trading partners are in a stronger financial position, they might help to ease the financial constraints of the Chinese firms with trade credit. (e.g. Nilsen 2002, Alfaro, García-Santana and Moral-Benito 2021 and Ersahin, Giannetti and Huang 2024). However, when the international partner firms themselves are affected by worsening financial liquidity due to a US monetary tightening, they may also transmit the effect to Chinese exporters.

[44] This variable $APay$ is available from 2004 to 2006. We don't have the sources of trade credits, but an overall account for an exporter. Nevertheless, on average, the exporter's total trade credits shrink and its liquidity conditions deteriorate.

[45] It is worth noting that, beyond the liquidity implication, a reduction in trade credit may also reflect an increase in the opportunity cost of providing such credit. This can directly induce exporters to raise prices. Alternatively, the lack of trade credit may impair firms' productivity, thereby exerting additional upward

Table 4: Liquidity changes of exporters

	(1)	(2)	(3)	(4)
Dependent Var	Direct measures		Indirect measures	
	$\Delta Cash_{it}$	$\Delta Liquid_{it}$	$\Delta APay_{it}$	$\Delta ARec_{it}$
brw_t	-0.018***	-0.012**	-0.025***	-0.012***
	(0.004)	(0.005)	(0.006)	(0.004)
$Sales_{it-1}$	-0.003***	-0.011***	-0.016***	-0.018***
	(0.001)	(0.001)	(0.002)	(0.001)
$Debt_{it-1}$	-0.014***	0.630***	-0.310***	-0.066***
	(0.005)	(0.007)	(0.008)	(0.004)
Firm FE	Yes	Yes	Yes	Yes
Observations	155699	155699	88076	155699

Notes: This table displays the liquidity responses of Chinese exporters to the US monetary shocks. The specification is $\Delta Liq_{it} = \alpha + \beta \cdot m_t + \Gamma \cdot \mathbf{Z}_{it-1} + \xi_i + \varepsilon_{it}$, where the dependent variables in columns (1)-(4) are changes in cash over total assets, net liquidity assets over total assets, accounts payable over total assets, and accounts receivable over total assets, respectively. Z_{it-1} is a group of firm-specific one-year lagged control variables, including log real sales income (a proxy for firm size) and the ratio of total debt to total assets. All regressions include firm fixed effects.

5.3 Borrowing cost and export price

After confirming the effect of the monetary policy on Chinese exporting firms' liquidity conditions, we now check what happens to borrowing costs. We argue that due to the worsening of liquidity conditions, firms are forced to borrow more from outside institutions, thus yielding a higher average borrowing cost. To verify this hypothesis, we use a specification similar to Equation 3, and the dependent variables are changes (first difference) in the borrowing cost or debt ratio. To measure the average borrowing cost, we use four variables: the ratio of interest rate expenditure over total debt IE/L, the ratio of interest rate expenditure over current debt IE/CL, the ratio of financial expenses over total debt FN/L, and the ratio of financial expenses over current debt FN/CL.[46] The total debt may contain funding borrowed from financial markets, payroll payable, trade account payable, etc.[47] Only the debt of financial institutions requires an explicit interest expenditure. So, an increase in the average borrowing cost may originate from the rise of the borrowing rate itself or the increase in borrowing

pressure on prices. Both mechanisms reinforce our main finding. A more thorough investigation of these channels would require detailed trade credit data, which we leave for future research.

[46] The financial expenses include both the interest rate expenditures and some other financing related costs, for example accounting and auditing fees, etc.

[47] The ASIE database doesn't provide information for each type of debt, but only the aggregate level.

proportion from the financial markets.[48]

Table 5: Borrowing cost changes of exporters

	(1)	(2)	(3)	(4)	(5)	(6)
Dependent Var		Borrowing costs			Liability	
	$\Delta\frac{IE}{L}_{it}$	$\Delta\frac{IE}{CL}_{it}$	$\Delta\frac{FN}{L}_{it}$	$\Delta\frac{FN}{CL}_{it}$	$\Delta Debt_{it}$	$\Delta CDebt_{it}$
brw_t	0.005***	0.007***	0.014***	0.015***	0.039**	0.038**
	(0.001)	(0.001)	(0.002)	(0.003)	(0.017)	(0.019)
$Sales_{it-1}$	-0.000*	-0.001	-0.001*	-0.002**	-0.144***	-0.147***
	(0.000)	(0.000)	(0.000)	(0.001)	(0.005)	(0.006)
$Debt_{it-1}$	0.033***	0.038***	0.069***	0.077***	-2.318***	-2.208***
	(0.001)	(0.002)	(0.002)	(0.003)	(0.024)	(0.025)
Firm FE	Yes	Yes	Yes	Yes	Yes	Yes
Observations	155008	153219	155008	153219	154908	153086

Notes: This table displays the borrowing cost responses of Chinese exporters to the US monetary shocks. the specification is similar to Table 4. The only difference is that the dependent variables in columns (1)-(4) are changes (first difference) in interest expense over the total liability ratio, interest expense over the current liability ratio, total financial expense over the total liability ratio, and total financial expense over the current liability ratio, respectively. The dependent variables $Debt$ and $CDebt$ in columns (5)-(6) are changes in total and current liability over total asset ratios. All regressions include firm fixed effects.

As shown in Table 5, we find that the firm's borrowing cost increases significantly in response to a US contractionary shock, which is consistent with our previous conjecture.[49] Additionally, both firms' total debt ratio (total debt over total assets) and current debt ratio (current debt over total assets) increased after a tightening shock, suggesting that firms are relying more on external financing.[50] Furthermore, we also demonstrate that the price impact is bigger if a firm faces higher borrowing costs (see Table 6). The specification is:

$$\Delta \ln P_{it} = \alpha + \beta \cdot m_t \cdot X_{st-12} + \Gamma \cdot \mathbf{Z} + \xi_i + \xi_t + \varepsilon_{it} \tag{4}$$

[48]The average borrowing cost $BC = \frac{Interest}{Debt_T} = \frac{Interest}{Debt_F} \cdot \frac{Debt_F}{Debt_T} \equiv BR \cdot BP$, where $Interest$ means interest expenditures or financial expenses, $Debt_F$ ($Debt_T$) is the financial (total) debt, BR denotes borrowing interest rate, and BP represents the borrowing portion from financial markets.

[49]The increase of average borrowing cost doesn't necessarily require Chinese exporters to borrow from the international financial market. This can also happen when firms only borrow from domestic institutions. That's because domestic external finance is also more expensive than internal funding, thus, the average borrowing cost will go up as long as the borrowing proportion increases.

[50]Although we don't have a direct measure of financial debt, we know that the accounts payable decrease in the face of a tightening shock, which indicates that the increase of total debt is very likely contributed by the rise of financial debt.

where the dependent variable is the monthly year-on-year price changes and X_{st-12} is the one-year lagged average borrowing cost.[51] $\mathbf{Z}$ is firm-specific controls, including one-year lagged log real sales income and one-month lagged price changes. ξ_i and ξ_t are firm and time-fixed effects respectively.[52] The standard errors here are clustered at the firm level and the results are robust to clustering at the time level. Besides, it is found that firms with ex-ante higher borrowing costs would get a larger rise in the borrowing costs, which reconciles with the bigger movement of prices. These results are displayed in Table C1.

Table 6: Interactions with borrowing cost

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Dependent Var				Monthly $\Delta \ln P_{it}$				
$brw_t \times \frac{IE}{L}_{st-12}$	7.645***	6.959***						
	(2.259)	(2.141)						
$brw_t \times \frac{IE}{CL}_{st-12}$			6.269***	5.614***				
			(1.902)	(1.803)				
$brw_t \times \frac{FN}{L}_{st-12}$					6.288***	3.694*		
					(2.387)	(2.245)		
$brw_t \times \frac{FN}{CL}_{st-12}$							5.153***	3.069*
							(1.953)	(1.841)
$Sales_{it-12}$		-0.017***		-0.017***		-0.017***		-0.017***
		(0.001)		(0.001)		(0.001)		(0.001)
$\Delta \ln P_{it-1}$		0.296***		0.296***		0.296***		0.296***
		(0.003)		(0.003)		(0.003)		(0.003)
Firm FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Year-month FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Observations	1072227	917419	1072227	917419	1072227	917419	1072227	917419

Notes: The specification is $\Delta \ln P_{it} = \alpha + \beta \cdot m_t \cdot X_{st-12} + \Gamma \cdot \mathbf{Z} + \xi_i + \xi_t + \varepsilon_{it}$. The interaction terms in columns (1)-(2), (3)-(4), (5)-(6), and (7)-(8) are interest expense over the total liability ratio, interest expense over the current liability ratio, total financial expense over the total liability ratio and total financial expense over the current liability ratio, respectively. $\mathbf{Z}$ is firm-specific controls, including one-year lagged log real sales income and one-month lagged price changes. ξ_i and ξ_t are firm and time-fixed effects respectively. All regressions include firm and time-fixed (year-month pair) effects.

Similarly, consistent with the borrowing cost channel, if the borrowing cost increase is due to the worsening of liquidity conditions, we would expect that the price change should be related to these factors. Firms under different liquidity states may react distinctly in response to the same tightening shock. This conjecture is validated in Table C2. The specification is

[51] To alleviate endogeneity, we aggregate the variables at the sector level and use the value of last year. Here and thereafter, we mainly use the monthly price changes for illustration and the annual results are consistent.

[52] Due to the control of time-fixed effect, all the time-varying factors, including the monetary shock, are absorbed. Our interest parameter here is β, which represents the heterogeneous responses form different firms.

similar to Equation 4, and the only difference is replacing the borrowing cost measurements with the liquidity variables. It is found that firms with fewer cash holdings and net liquid assets would experience a bigger increase in export prices.

5.4 Markup, other costs, and interest rate

To further validate the cost-driven channel, we also provide suggestive evidence about the role of markup adjustments in the export pricing by decomposing the export price into markup and marginal cost, following De Loecker and Warzynski (2012) and Brooks, Kaboski and Li (2021).[53] It is found that only marginal cost responds positively and significantly to the US monetary shock (see Table 7). Meanwhile, the reaction to markup change is quite mild both economically and statistically, which is not fully in line with the implication of the "pricing-to-market" theory. Instead, it is consistent with the finding of Li, Ma and Xu (2015) that Chinese exporters have very limited ability to absorb shocks by adjusting markups and almost completely passing exchange rate shocks to their export prices. Consistently, when we control the change of markup or cost in our baseline regression, it shows that controlling cost change will substantially reduce the coefficient of monetary shock while controlling markup change barely changes the results.

[53] We derive the firm-specific markup as the ratio of an input factor's output elasticity to its firm-specific factor payment share $\mu_t = \theta_t^X \left(\alpha_t^X\right)^{-1}$, where α_t^X is the share of expenditures on input X in total sales and θ_t^X denotes the output elasticity on input X. We apply the methodology of Ackerberg, Caves and Frazer (2015) to address the endogeneity of inputs, assuming a third-order translog gross output production function. The marginal cost, therefore, could be written as $MC_{it} = P_{it}/\mu_{it}$ and $\Delta \ln(MC)_{it} = \Delta \ln P_{it} - \Delta \ln \mu_{it}$.

Table 7: Decomposition of prices: markup vs marginal cost

	(1)	(2)	(3)	(4)	(5)	(6)
Dependent Var	$\Delta \ln \mu_{it}$	$\Delta \ln MC_{it}$	Monthly $\Delta \ln P_{it}$		Annual $\Delta \ln P_{it}$	
brw_t	-0.011*	0.168***	0.153***	0.026***	0.250***	0.094***
	(0.006)	(0.010)	(0.012)	(0.006)	(0.011)	(0.006)
$\Delta \ln \mu_{it}$			0.009**		0.014***	
			(0.003)		(0.005)	
$\Delta \ln MC_{it}$				0.788***		0.618***
				(0.003)		(0.004)
$\Delta \ln P_{it-1}$			0.279***	0.063***	-0.312***	-0.119***
			(0.003)	(0.001)	(0.005)	(0.003)
$Sales_{it-n}$	-0.019***	0.014***	-0.005**	-0.019***	-0.014***	-0.020***
	(0.002)	(0.003)	(0.002)	(0.002)	(0.003)	(0.002)
NER Control	No	No	Yes	Yes	Yes	Yes
Firm FE	Yes	Yes	Yes	Yes	Yes	Yes
Observations	110510	105098	663876	662132	81348	81098

Notes: This table shows the responses of markup and marginal cost to the US monetary shock. The specification in columns (1)-(2) is $\Delta Y_{it} = \alpha + \beta \cdot m_t + \gamma \cdot Sales_{it-1} + \xi_i + \varepsilon_{it}$, where the dependent variables are annual changes in markup and marginal cost. The specification in columns (3)-(6) is similar to the baseline, and here we additionally control the change of markup and marginal cost. The dependent variables in columns (3)-(4), (5)-(6) are monthly and annual changes in prices, respectively. The standard errors here are clustered at the firm level, and the results are robust to clustering at the time level.

In addition, we explore heterogeneity in firms' markeups. Specifically, we examine whether a firm's relative markup within a sector (within-sector markup) or the median markup at the sector level (across-sector markup) affect the export price response to US monetary policy shocks in Table C3. We find that all interaction terms are insignificant, implying that the export prices of firms with different markups do not exhibit significant differences in response to monetary policy shocks. Moreover, we also test that among all the costs, only the borrowing cost responds substantially. The responses of material input cost and labor cost are insignificant. The level of import intensity (ratio of imports to total material inputs) is also irrelevant in explaining the impact of monetary shocks on prices, which implies that the price increase is not mainly contributed by the costs of imported goods.[54] These results are shown in Table C4.

Besides, regarding why borrowing costs increase after a tightening shock, there may be another possible explanation: the US tightening increases China's interest rates. To verify this possibility, we regress the overnight return of Chinese treasury bonds and corporate bonds price index on the US monetary policy shock, and the results turned out to be relatively weak and insignificant, although the reaction direction is consistent (see Table C5). This implies

[54]As shown in Table B9, pure exporter also increase export prices.

that the average borrowing cost movement is mainly driven by the higher reliance on more expensive external financing rather than the increases in the borrowing rate itself. This is plausible as China has quite tight capital control and a highly independent monetary authority in our sample period; thus, China's financial market is not tightly exposed to US monetary policy adjustments. Similar results are also documented in previous papers such as Hausman and Wongswan (2011) and Ho, Zhang and Zhou (2018), etc. This result is interesting because it suggests that exporters' financing conditions are affected by external monetary shock, even though the Chinese financial market itself responds quite mildly.

To support this cost channel, we also provide additional cross-sectional evidence showing heterogeneous impacts across FDI and domestic firms, in destination markets with varying levels of financial development, and between processing and ordinary trade. See Section C.2. Finally, we have more discussion on several alternative stories, such as global demand shift, international competition, and exchange rate pass-through. See Section C.3.

6 Extension

6.1 China's monetary policy stance

In this extension, we examine how the US monetary policy shocks interact with the Chinese monetary condition. Because China does not have a deep and liquid derivative market in our main sample, we cannot construct the monetary policy surprise series similar to the US series. Instead, we use negative normalized M2 growth rate as a measure of China's domestic monetary tightness, and a bigger value means tighter conditions. This reflects the greater relative reliance on quantity tools by the Chinese central bank in conducting its monetary policy. [55]

[55] In practice, we construct two measures, year-on-year tightness index, $tightness_{YoY} = -(M2_t - M2_{t-12})/M2_{t-12}$, and month-on-month tightness index, $tightness_{MoM} = -(M2_t - M2_{t-1})/M2_{t-1}$, where t is an index of month.

Table 8: Domestic monetary tightness in China

	(1)	(2)	(3)	(4)
Dependent Var	Monthly $\Delta \ln P_{it}$			
	Year-on-year tightness		Month-on-month tightness	
brw_t	0.145**	0.125*	0.261***	0.204***
	(0.072)	(0.065)	(0.084)	(0.075)
$brw_t \times CN - tightness_t^{YoY}$	0.132**	0.075*		
	(0.051)	(0.045)		
$brw_t \times CN - tightness_t^{MoM}$			0.106**	0.075*
			(0.044)	(0.040)
$tightness_t^{YoY}$	0.006***	0.004**		
	(0.002)	(0.002)		
$tightness_t^{MoM}$			0.001	0.000
			(0.002)	(0.002)
$Sales_{it-12}$		-0.006**		-0.005**
		(0.002)		(0.003)
$\Delta \ln P_{it-1}$		0.298***		0.299***
		(0.006)		(0.006)
NER Control	Yes	Yes	Yes	Yes
Firm FE	Yes	Yes	Yes	Yes
Observations	1100400	917419	1100400	917419

Notes: In this table, compared with the baseline, we control the stance of Chinese monetary policy and its interaction term with BRW shock. Chinese monetary policy stance *tightness* in columns (1)-(2) is measured by the minus year-on-year M2 growth rate, while in columns (3)-(4) it is the minus month-on-month M2 growth rate. Robust standard errors are based on two-way clustering at both the firm level and time level (year-month for monthly regression); *, **, and *** indicate significance at 10%, 5%, and 1% levels.

In Table 8, the effects of the US shocks are robust and the coefficients of *tightness* are positive, which implies that domestic tightening also causes exporters to raise prices. Besides, it is worth noting that all interaction coefficients are positive, indicating that a US contractionary shock would have a larger impact conditional on a tighter domestic monetary environment. This is consistent with Lin and Ye (2018*a*) who find that the impacts of foreign liquidity shock on trade credits of Chinese firms are more profound in tighter domestic monetary conditions.

6.2 ECB monetary policy

In addition to the spillover effect of the US monetary policy shock, recent literature (see Ca'Zorzi et al. (2020), Corsetti et al. (2021), and Miranda-Agrippino and Nenova (2022), etc.) indicates that the monetary policy of the European Central Bank (ECB) also has a substantial effect on global financial conditions and real economic activities. Using a similar specification as our baseline regression, we explore how China's export prices respond to ECB monetary shocks in addition to the US Fed shock, shown in Table 9.

Table 9: Export price responses to EU monetary policy shocks

	(1)	(2)	(3)	(4)
	To ECB markets	To US market	To other countries	To all countries
Dependent Var	Monthly $\Delta \ln P_{it}$			
$US - brw_t$	0.146*	0.103*	0.163**	0.151**
	(0.077)	(0.060)	(0.068)	(0.065)
$ECB - MP_t$	0.078	-0.003	0.017	0.011
	(0.064)	(0.062)	(0.056)	(0.056)
$ECB - CBI_t$	-0.025	0.012	-0.038	-0.018
	(0.063)	(0.063)	(0.058)	(0.056)
$Sales_{it-12}$	0.003	-0.010***	-0.004	-0.005*
	(0.005)	(0.003)	(0.003)	(0.003)
$\Delta \ln P_{it-1}$	0.195***	0.299***	0.279***	0.299***
	(0.007)	(0.007)	(0.006)	(0.006)
NER Control	Yes	Yes	Yes	Yes
Firm FE	Yes	Yes	Yes	Yes
Observations	183189	247028	779883	917419

Notes: This table investigates the impact of the European Central Bank shock. The specification is similar to the baseline, and we additionally include the ECB shocks, including the pure monetary policy shock MP and the central bank information shock CBI. The ECB shocks in all columns are from Jarociński and Karadi (2020), which are rescaled so that each interest rate surprise has the standard deviations of the 1-year OIS swap rate. Columns (1)-(4) use the changes in export prices to the ECB market, the US market, other countries, and all countries, respectively. Robust standard errors are based on two-way clustering at both the firm level and time level (year-month for monthly regression); *, **, and *** indicate significance at 10%, 5%, and 1% levels.

For ease of comparison, we report the result of the US BRW shock in the first row and the ECB shocks in the second and third rows. In columns (1)-(4), we show the export price responses in the Eurozone, the US market, other countries, respectively, as well as all countries in our samples. The impacts of the US monetary policy shocks on export prices are robust after controlling for ECB shocks. However, Chinese export prices barely move in response to the ECB shocks.

Our finding is consistent with the literature (like Ca'Zorzi et al. (2020), Corsetti et al. (2021), and Miranda-Agrippino and Nenova (2022), etc.) that finds a less powerful spillover effect from ECB shocks than from the US Federal Reserve. This confirms the special role of the US monetary policy in the global financial cycle given the dominant role of the US dollar along with the intensive integration of the global financial market (Miranda-Agrippino and Rey (2020)).

6.3 Price responses in other periods and countries

As documented in the motivating facts, abnormal price responses to monetary shocks are evident over a long horizon (1995–2019). To investigate the underlying mechanism, we focus on China during January 2000–December 2006, when highly disaggregated firm-level data are available. A natural question is how these effects evolve outside this sample window and across other countries.

In theory, the cost channel should operate whenever exporters' liquidity conditions deteriorate in response to U.S. monetary tightening (a pattern widely documented across countries and periods; see Miranda-Agrippino and Rey 2020). However, whether this channel dominates—i.e., whether prices rise following a tightening—depends on the joint equilibrium outcome of demand- and supply-side forces, whose relative strength may vary with the domestic and international environment. Importantly, even an insignificant price response does not imply that supply-side channels were irrelevant. Rather, they likely continued to offset demand-side forces, preventing the decline in export prices that would be predicted by demand effects alone.

The relative importance of these forces may vary across periods and contexts. After the global financial crisis, advanced economies entered a zero lower bound environment (2009–2015), likely weakening the borrowing cost channel. At the same time, evidence of a flatter Phillips curve in advanced economies (e.g., Florio, Siena and Zago 2025) suggests that monetary policy had a diminished effect on prices through the demand side, implying a larger relative role for the cost channel. For China, several institutional changes may also have shaped the price responses. The adoption of a more flexible exchange rate regime could partially cushion adverse cost-push shocks, while the gradual relaxation of capital control policies may have exposed China more directly to global financial fluctuations, increasing the role of financing costs.

Data availability prevents us from fully assessing this question for the post-crisis period and for other economies. Extending the analysis to more recent years, including the post-COVID era, as well as to other countries, remains an important avenue for future research.[56]

6.4 More discussion on the model

In the baseline model, we assume one input factor and no credit constraint, use flexible and static pricing, and ignore the role of exchange rate and currency invoicing. However, our channel is robust even when relaxing these assumptions. (1) If capital is added as another factor, the price should be additionally affected by monetary shocks through the capital rental rate. Together with other input prices, these factor prices should be determined in a general equilibrium, and the movement depends on the specific model settings. In this case, the role of liquidity is mixed with other potential channels. (2) If we assume an exporter has a binding credit constraint (e.g. Manova and Zhang 2012, Manova 2013, Manova, Wei and Zhang 2015), the price should also be affected by the collateral rate. Nevertheless, liquidity still plays a role. Intuitively, a tightening US monetary policy shock would reduce an exporter's liquidity conditions, which accordingly increases firms' credit demand. To satisfy the requirement from credit needs, the exporters are forced to increase prices and obtain more cash flow (relative to the funds in need) to improve credit access.[57] (3) If we allow dynamic and sticky price settings, the prices will increase in response to the rise of both current and future marginal costs. A firm with lower price rigidity should increase more. Besides, in this case, the markup is no longer constant, which is implicitly affected by some variables like the total import demand and aggregate price index. (4) In a more realistic setting where there are exchange rates and differences in invoicing currency, the price should also be adjusted by the movement of the exchange rate. For example, in a Producer Currency Pricing case, if the US tightening shock appreciates the US dollar against the home currency of an exporter, the export price in the US dollar should be lower. In different invoicing currencies, the role of exchange rates should be slightly different. For more details, please refer to Appendix D.5.

7 Conclusion

This paper studies the spillback and spillover effects of US monetary policy through the import and export prices. We take China as an example to study the pricing of exporters in

[56]Nevertheless, using the cross-country import price index (adjusted for quality) from the Penn World Table (1994–2019), we find consistent suggestive evidence: the U.S. monetary tightening tends to increase import prices, with the effect becoming even more pronounced after the financial crisis. Detailed results are available upon request.

[57]This is consistent with Gilchrist et al. (2017) who show that financial distortions create an incentive for firms to raise prices in response to adverse financial or demand shocks so as to preserve internal liquidity.

response to the US monetary shocks. Using exogenous monetary shocks, disaggregated custom transaction records, and comprehensive firm-level balance sheet data, we find that exporters do not reliably lower their export prices in response to a contraction in total demand following a tightening of US monetary policy. The monetary contraction mainly affects export prices through an "international cost channel", which is related to firms' liquidity conditions. This is also not fully consistent with the prediction of the "pricing-to-market" theory that firms could adjust markups to absorb adverse shocks.

In an era characterized by the increasing integration of global trade and finance, understanding how export prices adapt to global monetary policy shocks in the presence of financial frictions is crucial for both market players and policymakers. Our paper casts new light on the special role of US monetary policy shocks in shaping international trade through their influence on exporters' liquidity conditions and financing costs. We use China, the largest exporter in the world, as an example, which has general implications for all economies. The different responses of other countries' export prices to US monetary policy shocks could be the subject of future research. Besides, the response of exporters' pricing behaviors also provides new implications on how the Fed's monetary policy could affect US and global inflation through the trade connection. Many interesting and important questions remain in this area, and we hope our paper can serve as a stepping stone for future investigation.

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Appendix

A More about data and measurements

Table A1: Summary statistics of firm information

	Mean	SD	p50	p25	p75
$\Delta \ln P^{all}$	0.03	0.42	0.01	-0.11	0.17
$\Delta \ln P^{US}$	0.04	0.34	0.02	-0.11	0.19
Number of HS6 Products	6.29	10.31	3.00	2.00	7.00
Sales (*million RMB)	160	1201	34.91	15.35	90.85
Employment (persons)	449	1210	197	96	418
ϕ^{exp} (Export/Sales)	0.46	0.38	0.36	0.07	0.89
Firm-year observations			270271		
Number of Firms			88425		

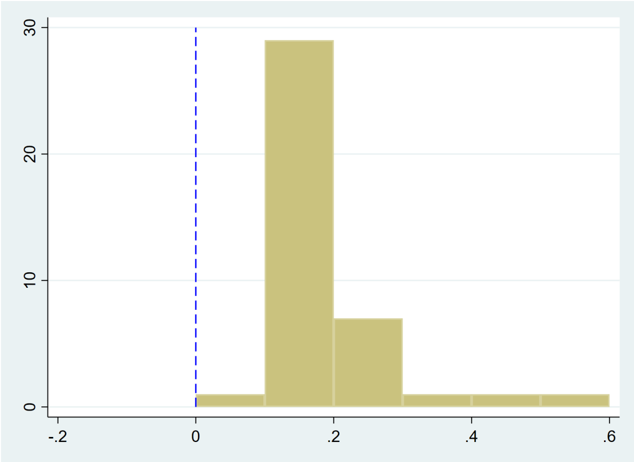
Notes: This table shows the summary statistics of firms in the matched sample. The first two rows ΔP^{all} and ΔP^{US} indicate monthly price changes exporting to all other countries and the US market, respectively, while all other rows describe annual-level firm variables. The third row denotes the number of HS6 product types a company exports in a given year. ϕ^{exp} represents the export intensity, which is the firm-level ratio of exports to total sales.

Table A2: Correlations of alternative monetary policy shock measures

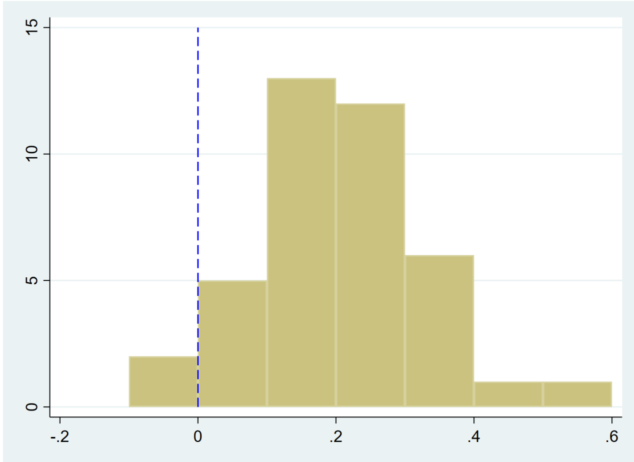
	brw	NS	BS	$Target^{Acosta}$	$Path^{Acosta}$	MP^{JK}	CBI^{JK}
brw	1						
NS	0.5398	1					
BS	0.4863	0.8636	1				
$Target^{Acosta}$	0.2793	0.6259	0.5495	1			
$Path^{Acosta}$	0.4702	0.7901	0.6768	0.0178	1		
MP^{JK}	0.4210	0.6391	0.8591	0.4361	0.4798	1	
CBI^{JK}	0.1512	0.4245	0.3768	0.3457	0.2680	-0.0897	1

Notes: The monetary policy shock measures are from Bu, Rogers and Wu (2021), Nakamura and Steinsson (2018), Bauer and Swanson (2023*b*), Acosta (2022), and Jarociński and Karadi (2020), respectively.

B More results on price responses



(a) Unconditional price responses of 40 countries (without control)

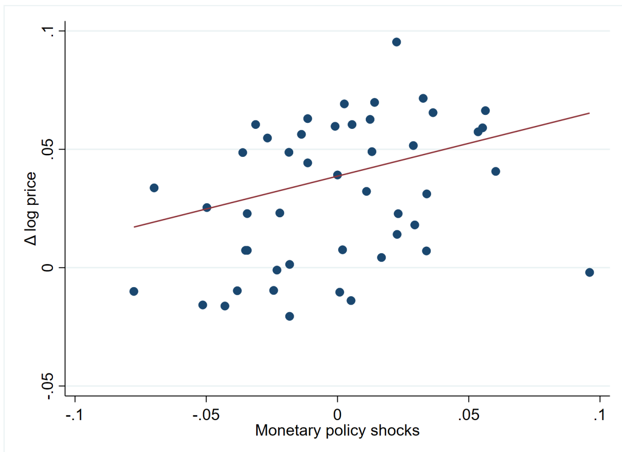


(b) Conditional price responses of 40 countries (with controls)

Figure B1: Import price responses of 40 major countries

Notes: The horizontal axis shows the regression coefficients of import price (simple average of HS6 product level unit value) changes to US monetary policy shocks (Bu, Rogers and Wu, 2021) of each source among top 40 countries (excluding the US) in terms of nominal GDP in 2006. Panel (a) shows the distribution of unconditional import price responses. Panel (b) adds control terms for price changes in the previous year and GDP growth. For display purpose, we place all observations greater than 0.6 into the rightmost bin.

(a) To the US market



(b) To non-US markets

Figure B2: The correlation between US monetary policy shock and China's export price changes

Notes: The horizontal axis represents monetary policy shocks. One unit of positive shock means an unexpected increase in the daily 2-year US treasury rate by 100 basis points. The vertical axis denotes the simple average of year-over-year price change of Chinese exporters to the US and the non-US market across all the firms.

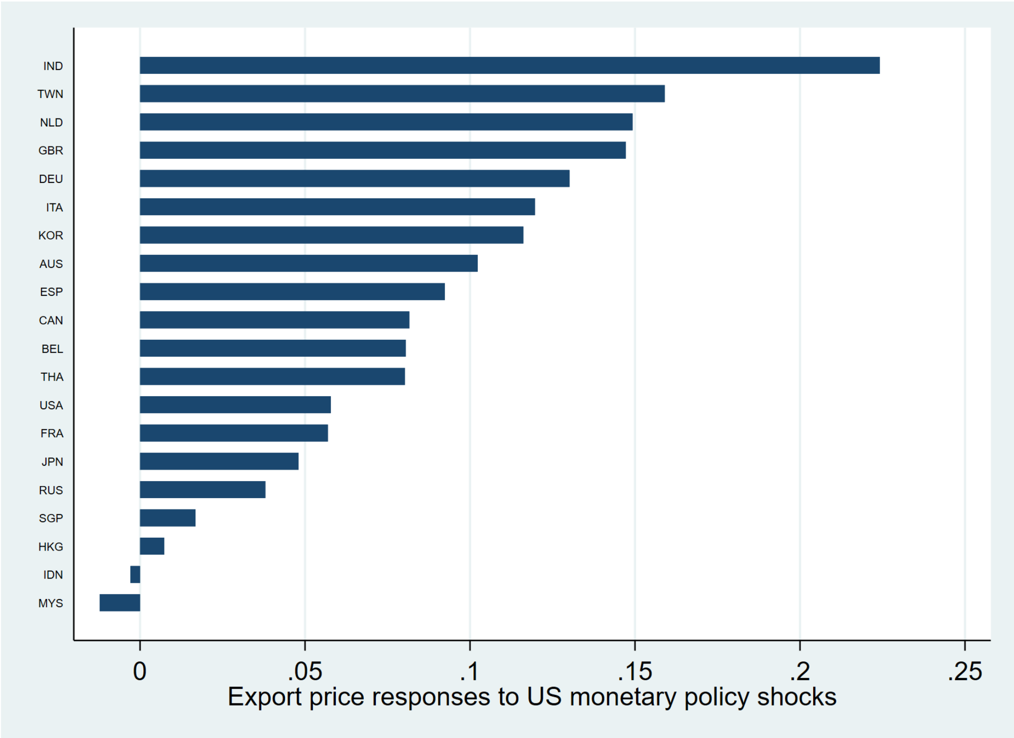
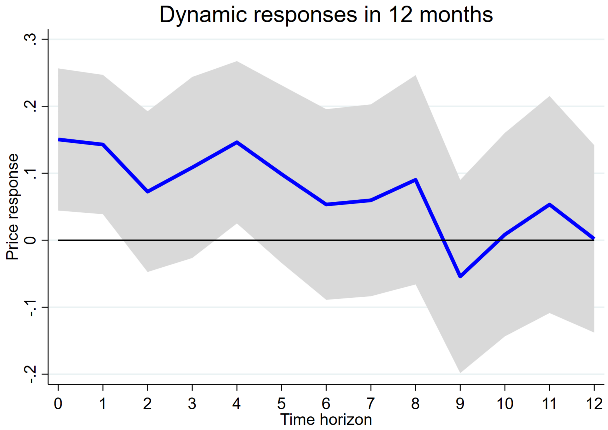
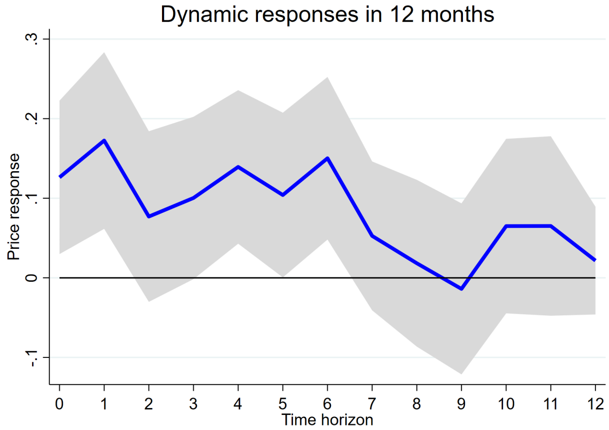


Figure B3: China's export price responses to top 20 trading partners

Notes: The estimation is at the firm-product level for each destination market sub-sample. The specification is $\Delta \ln P_{iht} = \alpha + \beta \cdot m_t + \Psi \cdot \Omega_{ct} + \xi_{ip} + \varepsilon_{iht}$, where i and h are indexes for firm and product, Ω_{ct} means country-specific time-varying controls including changes in bilateral nominal exchange rates, CPI inflation, and real GDP growth, and ξ_{ip} is firm-product fixed effects. The horizontal coordinate represents the regression coefficients for the sub-sample of each country.



(a) Dynamic responses: forward prices



(b) Dynamic responses: lagged shocks

Figure B4: Dynamic responses to monetary policy shocks

Notes: Panel (a) shows the dynamic impacts of the US monetary shocks using forward prices. The specification is $\Delta \ln P_{it+\tau} = \alpha + \beta \cdot m_t + \Gamma \cdot \mathbf{Z} + \Psi \cdot \Omega_t + \xi_i + \varepsilon_{it}^{\tau}$, where $\Delta \ln P_{it+\tau}$ are the monthly year-over-year price changes in month $t+\tau$, τ=1, 2, ..., 12, and the other controls are similar to the baseline. Panel (b) shows the dynamic impacts of the US monetary shocks using lagged shocks. The specification is $\Delta \ln P_{it} = \alpha + \sum_{j=1}^{\tau} \beta_j \cdot m_{t-j} + \Gamma \cdot \mathbf{Z} + \Psi \cdot \Omega_t + \xi_i + \varepsilon_{it}^{\tau}$, where m_{t-j} are the monetary shocks in month $t-j$, j=1, 2, ..., 12, and the other controls are similar to the baseline. In both panels, we plot β coefficients, and the shaded areas indicate the 10% confidence interval.

Table B1: Export value and quantity responses to US monetary policy shocks

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Firm level value				Firm-product level quantity			
Dependent Var	Monthly $\Delta \ln V_{it}$		Annual $\Delta \ln V_{it}$		Monthly $\Delta \ln Q_{it}$		Annual $\Delta \ln Q_{it}$	
brw_t	0.133	0.211	-0.628**	-0.184	-0.018	0.036	-1.930*	-2.011
	(0.372)	(0.366)	(0.221)	(0.242)	(0.398)	(0.333)	(0.960)	(1.084)
$Sales_{it-n}$		-0.254***		-0.245***		-0.264***		-0.059
		(0.014)		(0.038)		(0.016)		(0.188)
$\Delta \ln P_{it-1}$		0.210***		-0.456***		0.203***		-0.387***
		(0.005)		(0.093)		(0.005)		(0.025)
NER control	Yes	Yes	Yes	Yes	No	No	No	No
Firm FE	Yes	Yes	Yes	Yes	No	No	No	No
Firm-Product FE	No	No	No	No	Yes	Yes	Yes	Yes
Observations	1140624	986757	154732	99751	2359502	1751828	571830	314287

Notes: Here we investigate the value and quantity responses to the US monetary shocks using samples of all countries. The specification is similar to the baseline. The only difference lies in the dependent variable. Columns (1)-(4) show results of firm-level value, while columns (5)-(8) show results of firm-product-level quantity. All regressions include firm fixed effects. Robust standard errors are based on two-way clustering at both the firm and time levels (year-month for monthly regression and year for annual regression). *, **, and *** indicate significance at 10%, 5%, and 1% levels.

Table B2: Weighted shocks using announcement dates

	(1)	(2)	(3)	(4)	(5)	(6)
Dependent Var	Monthly $\Delta \ln P_{it}$			Annual $\Delta \ln P_{it}$		
$brw_t^{weighted}$	0.210**	0.212**	0.133*	0.159***	0.167***	0.265***
	(0.093)	(0.095)	(0.072)	(0.033)	(0.035)	(0.054)
$Sales_{it-12}$		-0.004	-0.005*		-0.015**	-0.020*
		(0.003)	(0.003)		(0.004)	(0.009)
$\Delta \ln P_{it-1}$			0.299***			-0.318***
			(0.006)			(0.033)
NER Control	Yes	Yes	Yes	Yes	Yes	Yes
Firm FE	Yes	Yes	Yes	Yes	Yes	Yes
Observations	1100400	1072227	917419	151542	147471	97987

Notes: The specification is similar to the baseline. Here we use samples of all countries and replace the original shocks with the weighted shocks, calculated according to the exact announcement dates. The shocks in columns (1)-(3) and (4)-(6) are in monthly and annual frequencies, respectively. Please refer to the main text for more details on the construction. All regressions include firm fixed effects. Robust standard errors are based on two-way clustering at both the firm and time levels (year-month for monthly regression and year for annual regression). *, **, and *** indicate significance at 10%, 5%, and 1% levels.

Table B3: Alternative aggregation levels of export prices

Panel A: monthly	(1)	(2)	(3)	(4)	(5)	(6)
Dependent Var	Firm-product level monthly $\Delta \ln P_{iht}$			Firm-product-country level monthly $\Delta \ln P_{ihct}$		
brw_t	0.140**	0.147**	0.127*	0.099*	0.104*	0.091
	(0.067)	(0.070)	(0.064)	(0.058)	(0.060)	(0.058)
$Sales_{it-12}$		-0.009***	-0.009***		-0.010***	-0.010***
		(0.003)	(0.003)		(0.003)	(0.003)
$\Delta \ln P_{ih(c)t-1}$			0.273***			0.274***
			(0.006)			(0.006)
Observations	2420018	2360154	1758341	3478000	3478000	2140247
Panel B: annual	(1)	(2)	(3)	(4)	(5)	(6)
Dependent Var	Firm-product level annual $\Delta \ln P_{iht}$			Firm-product-country level annual $\Delta \ln P_{ihct}$		
brw_t	0.168***	0.175***	0.247***	0.152***	0.164***	0.200***
	(0.039)	(0.043)	(0.051)	(0.026)	(0.029)	(0.041)
$Sales_{it-12}$		-0.016***	-0.008		-0.022***	-0.011*
		(0.002)	(0.004)		(0.003)	(0.005)
$\Delta \ln P_{ih(c)t-1}$			-0.426***			-0.449***
			(0.025)			(0.017)
Observations	573904	559749	315161	1138465	1086596	473955
NER Control	Yes	Yes	Yes	Yes	Yes	Yes
Country-time Controls	Yes	Yes	Yes	Yes	Yes	Yes
Firm-product FE	Yes	Yes	Yes	No	No	No
Firm-product-country FE	No	No	No	Yes	Yes	Yes

Notes: The specification is similar to the baseline. In panel A, the dependent variables in columns (1)-(3) are year-over-year changes in monthly firm-product level price, while in columns (4)-(6) are year-over-year changes in monthly firm-product-country level price. For the latter columns, we additionally control changes in bilateral nominal exchange rates, CPI inflation, and real GDP growth for the destination countries. In panel B, we report the annual version. All regressions include firm fixed effects. Robust standard errors are based on two-way clustering at both the firm and time levels (year-month for monthly regression and year for annual regression). *, **, and *** indicate significance at 10%, 5%, and 1% levels.

Table B4: Approximate year-over-year time match

	(1)	(2)	(3)	(4)	(5)	(6)
Dependent Var	Monthly $\Delta \ln P_{it}$					
	YoY + − 1 month			YoY + − 2 months		
brw_t	0.176**	0.197**	0.167**	0.187**	0.202**	0.172**
	(0.079)	(0.083)	(0.072)	(0.080)	(0.085)	(0.073)
$Sales_{it-12}$		-0.010***	-0.007**		-0.011***	-0.008**
		(0.004)	(0.003)		(0.004)	(0.003)
$\Delta \ln P_{it-1}$			0.342***			0.342***
			(0.007)			(0.007)
NER Control	Yes	Yes	Yes	Yes	Yes	Yes
Firm FE	Yes	Yes	Yes	Yes	Yes	Yes
Observations	1275434	1121510	943499	1358899	1130947	945449

Notes: The specification is similar to the baseline. The dependent variables in columns (1)-(3) are approximate year-on-year changes in monthly prices with time gaps from 11 to 13 months, while columns (4)-(6) are approximate year-on-year changes in monthly prices with time gaps from 10 to 14 months. All regressions include firm fixed effects. Robust standard errors are based on two-way clustering at both the firm and time levels (year-month for monthly regression). *, **, and *** indicate significance at 10%, 5%, and 1% levels.

Table B5: End-of-year export price responses

	(1)	(2)	(3)	(4)	(5)	(6)
Dependent Var	$\Delta \ln P_{it}$					
	To the US market		To non-US market		To all countries	
brw_t	0.118*	0.205**	0.147*	0.186*	0.144*	0.192**
	(0.051)	(0.049)	(0.064)	(0.068)	(0.062)	(0.066)
$Sales_{it-12}$		-0.041**		-0.010		-0.012
		(0.012)		(0.008)		(0.009)
$\Delta \ln P_{it-1}$		-0.403***		-0.372***		-0.372***
		(0.024)		(0.035)		(0.033)
NER Control	Yes	Yes	Yes	Yes	Yes	Yes
Firm FE	Yes	Yes	Yes	Yes	Yes	Yes
Observations	25087	12282	82226	42072	89141	46061

Notes: The dependent variables in all columns are December-to-December changes in monthly prices (except for those in 2005 due to missing data, which are replaced by approximate time matches). Columns (1)-(2), (3)-(4), and (5)-(6) include exports to the US, non-US markets, and all countries, respectively. All regressions include firm fixed effects. Robust standard errors are based on two-way clustering at both the firm and time levels (year-month for monthly regression). *, **, and *** indicate significance at 10%, 5%, and 1% levels.

Table B6: RMB price responses to monetary policy shocks

	(1)	(2)	(3)	(4)	(5)	(6)
Dependent Var	Monthly $\Delta \ln P_{it}^{RMB}$			Annual $\Delta \ln P_{it}^{RMB}$		
brw_t	0.180**	0.183**	0.150**	0.180***	0.195***	0.263***
	(0.075)	(0.077)	(0.065)	(0.040)	(0.054)	(0.044)
$Sales_{it-n}$		-0.004	-0.005*		-0.021*	-0.024***
		(0.003)	(0.003)		(0.009)	(0.006)
$\Delta \ln P_{it-1}$			0.299***			-0.317***
			(0.006)			(0.032)
NER Control	Yes	Yes	Yes	Yes	Yes	Yes
Firm FE	Yes	Yes	Yes	Yes	Yes	Yes
Observations	1100399	1072223	917424	155049	150863	97987

Notes: The specification is similar to the baseline. The dependent variables in columns (1)-(3) are changes in monthly prices denominated in the Chinese RMB, while columns (4)-(6) are changes in annual prices denominated in the Chinese RMB. All regressions include firm fixed effects. Robust standard errors are based on two-way clustering at both the firm and time levels (year-month for monthly regression and year for annual regression). *, **, and *** indicate significance at 10%, 5%, and 1% levels.

Table B7: Alternative sample: only single-product firms

	(1)	(2)	(3)	(4)	(5)	(6)
Dependent Var	Monthly $\Delta \ln P_{it}$			Annual $\Delta \ln P_{it}$		
brw_t	0.233***	0.219**	0.177**	0.210***	0.212***	0.312***
	(0.083)	(0.086)	(0.073)	(0.042)	(0.048)	(0.053)
$Sales_{it-n}$		-0.003	-0.006		-0.019*	-0.025**
		(0.005)	(0.004)		(0.008)	(0.008)
$\Delta \ln P_{it-1}$			0.272***			-0.344***
			(0.008)			(0.027)
NER Control	Yes	Yes	Yes	Yes	Yes	Yes
Firm FE	Yes	Yes	Yes	Yes	Yes	Yes
Observations	359864	265249	187491	21567	14675	8690

Notes: The specification is similar to the baseline using the samples of single-product firms. The dependent variables in columns (1)-(3) are changes in monthly prices, while columns (4)-(6) are changes in annual prices. All regressions include firm fixed effects. Robust standard errors are based on two-way clustering at both the firm and time levels (year-month for monthly regression and year for annual regression). *, **, and *** indicate significance at 10%, 5%, and 1% levels.

Table B8: Alternative sample: different ownership

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Dependent Var	Monthly $\Delta \ln P_{it}$				Annual $\Delta \ln P_{it}$			
	SOE	DPE	MNE	JV	SOE	DPE	MNE	JV
brw_t	0.215***	0.222***	0.136**	0.129**	0.201	0.274**	0.231***	0.248***
	(0.099)	(0.083)	(0.060)	(0.064)	(0.117)	(0.070)	(0.045)	(0.047)
$Sales_{it-n}$	0.015	0.008*	-0.012***	0.001	0.015	-0.005	-0.026***	-0.009
	(0.011)	(0.005)	(0.003)	(0.003)	(0.020)	(0.006)	(0.006)	(0.007)
$\Delta \ln P_{it-1}$	0.167***	0.186***	0.378***	0.286***	-0.314***	-0.346***	-0.280***	-0.290***
	(0.024)	(0.007)	(0.007)	(0.007)	(0.057)	(0.029)	(0.033)	(0.029)
NER Control	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Firm FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Observations	13429	197037	390138	316814	1613	25069	36565	33049

Notes: The specification is similar to the baseline using the samples of different ownerships. The ownership types of firms in columns (1)-(4) are state-owned enterprises, domestic private enterprises, multinational enterprises, and joint ventures, respectively. Columns (5)-(8) report the annual results. All regressions include firm fixed effects. Robust standard errors are based on two-way clustering at both the firm and time levels (year-month for monthly regression). *, **, and *** indicate significance at 10%, 5%, and 1% levels.

Table B9: Alternative sample: two-way traders vs pure exporters

Panel A: monthly	(1)	(2)	(3)	(4)	(5)	(6)
Dependent Var	Monthly $\Delta \ln P_{it}$					
	Two-way traders			Pure exporters		
brw_t	0.163**	0.165**	0.136**	0.189**	0.192**	0.173**
	(0.073)	(0.076)	(0.063)	(0.073)	(0.073)	(0.067)
$Sales_{it-12}$		-0.004	-0.004		-0.001	-0.004
		(0.003)	(0.003)		(0.004)	(0.004)
$\Delta \ln P_{it-1}$			0.320***			0.186***
			(0.006)			(0.008)
Observations	840092	817078	718544	259669	254523	198297
Panel B: annual	(1)	(2)	(3)	(4)	(5)	(6)
Dependent Var	Annual $\Delta \ln P_{it}$					
	Two-way traders			Pure exporters		
brw_t	0.171***	0.185***	0.232***	0.196***	0.198***	0.272***
	(0.036)	(0.038)	(0.045)	(0.046)	(0.049)	(0.051)
$Sales_{it-12}$		-0.017***	-0.013**		-0.013**	-0.015*
		(0.003)	(0.005)		(0.004)	(0.007)
$\Delta \ln P_{it-1}$			-0.278***			-0.368***
			(0.028)			(0.034)
Observations	101007	97980	66073	41899	40982	24084
NER Control	Yes	Yes	Yes	Yes	Yes	Yes
Firm FE	Yes	Yes	Yes	Yes	Yes	Yes

Notes: The specification is similar to the baseline. In panel A, the dependent variables are year-over-year changes in monthly price while in panel B, the dependent variables are changes in annual price. Columns (1)-(3) cover the sub-sample with two-way traders (both export and import), while columns (4)-(6) cover the sub-sample with pure exporters (only export). All regressions include firm fixed effects. Robust standard errors are based on two-way clustering at both the firm and time levels (year-month for monthly regression). *, *,*, and *** indicate significance at 10%, 5%, and 1% levels.

Table B10: Alternative standard error clusters and fixed effects

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Dependent Var	Monthly $\Delta \ln P_{it}$							
	FE 1		FE 2		Cluster 1		Cluster 2	
brw_t	0.034***	0.054***	0.219***	0.181***	0.180**	0.150**	0.180***	0.150***
	(0.010)	(0.010)	(0.012)	(0.012)	(0.076)	(0.066)	(0.021)	(0.022)
$Sales_{it-12}$		-0.017***		-0.005***		-0.005*		-0.005**
		(0.001)		(0.001)		(0.003)		(0.002)
$\Delta \ln P_{it-1}$		0.296***		0.299***		0.299***		0.299***
		(0.003)		(0.006)		(0.006)		(0.019)
NER Control	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Firm FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	No	No	No	No	No	No
Month FE	No	No	Yes	Yes	No	No	No	No
Cluster	Firm	Firm	Firm	Firm	Time	Time	Sector	Sector
Observations	1100400	917419	1100400	917419	1100400	917419	1100400	917419

Notes: The specification is similar to the baseline. Robust standard errors are clustered at the firm level for columns (1)-(4) and the time (year-month) level for columns (5)-(6), and industry level for columns (7)-(8); *, **, and *** indicate significance at 10%, 5%, and 1% levels. Regressions for columns (1)-(2) include firm fixed effects and year fixed effects, while those for columns (3)-(4) include firm fixed effects and month fixed effects, and only the firm level for columns (5)-(8).

Table B11: Additional control variables

	(1)	(2)	(3)	(4)	(5)
Dependent Var			Monthly $\Delta \ln P_{it}$		
	CN CPI	CN Value Added	VIX	Input Price	All
brw_t	0.152***	0.155**	0.157**	0.157**	0.156**
	(0.057)	(0.068)	(0.072)	(0.061)	(0.065)
CPI_{t-1}^{China}	0.221**				-0.043
	(0.099)				(0.143)
IVA_{t-1}^{China}		-0.014			-0.013
		(0.038)			(0.035)
$ln(VIX)_{t-1}^{US}$			-0.013**		-0.012**
			(0.006)		(0.005)
$\Delta \ln(P)_{t-1}^{input}$				0.067***	0.072***
				(0.011)	(0.021)
$Sales_{it-12}$	-0.006*	-0.005	-0.005*	-0.009***	-0.010***
	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)
$\Delta \ln P_{it-1}$	0.287***	0.287***	0.287***	0.286***	0.286***
	(0.006)	(0.006)	(0.006)	(0.006)	(0.006)
NER Control	Yes	Yes	Yes	Yes	Yes
Firm FE	Yes	Yes	Yes	Yes	Yes
Observations	815538	815538	815538	815538	815538

Notes: The specification is similar to the baseline. The control variables in columns (1)-(5) are CPI inflation in China, industrial value-added growth in China, the log of CBOE volatility index (VIX), and global industrial input (agriculture and mineral goods) price change. All the variables have a one-month lag. The control variables in column (5) are all above. All regressions include firm fixed effects. Robust standard errors are based on two-way clustering at both the firm and time levels (year-month for monthly regression). *, **, and *** indicate significance at 10%, 5%, and 1% levels.

Table B12: Dynamic panel GMM estimations

	(1)	(2)	(3)	(4)
Dependent Var	Monthly $\Delta \ln P_{it}$			
	Difference GMM		System GMM	
brw_t	0.057***	0.057***	0.057***	0.057***
	(0.014)	(0.015)	(0.014)	(0.014)
$Sales_{it-12}$		-0.029***		-0.020***
		(0.006)		(0.005)
$\Delta \ln P_{it-1}$	0.639***	0.639***	0.646***	0.637***
	(0.024)	(0.007)	(0.007)	(0.007)
NER Control	Yes	Yes	Yes	Yes
Firm FE	Yes	Yes	Yes	Yes
Observations	836117	816461	942113	919483

Notes: In this table, we use Arellano-Bond estimation to account for possible biases in the dynamic panel regressions. Columns (1)-(2) and (3)-(4) show the results with difference GMM and system GMM, respectively. Robust standard errors are reported in parentheses. *, **, and *** indicate significance at 10%, 5%, and 1% levels.

C More results on mechanism

C.1 More results on costs, liquidity, and markup

Table C1: Borrowing cost changes with lag interaction

	(1)	(2)	(3)	(4)
	Borrowing cost measures			
Dependent Var	$\Delta\frac{IE}{L}_{it}$	$\Delta\frac{IE}{CL}_{it}$	$\Delta\frac{FN}{L}_{it}$	$\Delta\frac{FN}{CL}_{it}$
$brw_t \times \frac{IE}{L}_{it-1}$	0.716***			
	(0.190)			
$brw_t \times \frac{IE}{CL}_{it-1}$		0.866***		
		(0.221)		
$brw_t \times \frac{FN}{L}_{it-1}$			1.076***	
			(0.201)	
$brw_t \times \frac{FN}{CL}_{it-1}$				1.156***
				(0.226)
$Sales_{it-1}$	-0.001***	-0.002***	-0.004***	-0.006***
	(0.000)	(0.000)	(0.001)	(0.001)
$Debt_{it-1}$	0.033***	0.038***	0.069***	0.076***
	(0.001)	(0.002)	(0.002)	(0.003)
Firm FE	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes
Observations	155008	153219	155008	153219

Notes: This table displays the heterogeneous responses of borrowing costs across exporters. The specification is $\Delta Y_{it} = \alpha + \beta \cdot m_t \cdot Y_{it-1} + \Gamma \cdot \mathbf{Z} + \xi_i + \xi_t + \varepsilon_{it}$, where m is monetary shock, Y in columns (1)-(4) are interest expense over the total liability ratio, interest expense over the current liability ratio, total financial expense over the total liability ratio, and total financial expense over the current liability ratio, Z is firm-level control including lagged sales income and debt ratio. All regressions include firm and year fixed effects.

Table C2: Interactions with liquidity

	(1)	(2)	(3)	(4)
Dependent Var		Monthly $\Delta \ln P_{it}$		
$brw_t \times Cash_{st-12}$	-1.765***	-2.181***		
	(0.505)	(0.476)		
$brw_t \times Liquid_{st-12}$			-1.133***	-1.062***
			(0.259)	(0.242)
$Sales_{it-12}$		-0.017***		-0.017***
		(0.001)		(0.001)
$\Delta \ln P_{it-1}$		0.296***		0.296***
		(0.003)		(0.003)
Firm FE	Yes	Yes	Yes	Yes
Year-month FE	Yes	Yes	Yes	Yes
Observations	1072227	917419	1072227	917419

Notes: The specification is similar to Table 6. The interaction terms in columns (1)-(2), (3)-(4) are the lag of cash over total asset ratio and net liquidity asset over total asset ratio, respectively. All regressions include firm and time (year-month pair) fixed effects.

Table C3: Within-sector and across-sector markup

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Dependent Var				Monthly $\Delta \ln P_{it}$			
$brw_t \times \mu_{it_0}$	0.072						
	(0.047)						
$brw_t \times \mathbf{1}\{\mu_{it_0} > \bar{\mu}_{cic4,t_0}\}$		0.004					
		(0.021)					
$brw_t \times \mathbf{1}\{\mu_{it_0} > \bar{\mu}_{cic2,t_0}\}$			0.006				
			(0.021)				
$brw_t \times \mu_{cic2,t-12}$				0.154			
				(0.191)			
$brw_t \times \mu_{cic2,t_0}$					0.280		
					(0.200)		
$brw_t \times \mu_{cic4,t-12}$						0.156	
						(0.178)	
$brw_t \times \mu_{cic4,t_0}$							0.274
							(0.190)
$Sales_{it-12}$	-0.017***	-0.017***	-0.017***	-0.017***	-0.017***	-0.017***	-0.017***
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
$\Delta \ln P_{it-1}$	0.295***	0.295***	0.295***	0.296***	0.296***	0.296***	0.296***
	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)
Firm FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Time FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Observations	901462	901462	901462	917419	917419	917410	917419

Notes: The specification is similar to Table 6. The interaction terms in columns (1)-(7) are firm-level markup at its initial export year, firms' above-median dummy within the CIC 2-digit and 4-digit sector, the median markup of each CIC 2-digit and 4-digit sector in which the firm operates, in the last year or its initial year, respectively. All regressions include firm and time fixed effects.

Table C4: Discussion about other production costs

Dependent Var	(1) $\Delta \frac{Input}{Sales}_{it}$	(2) $\Delta \frac{Wage}{Sales}_{it}$	(3)	(4) Monthly $\Delta \ln P_{it}$	(5)
brw_t	0.075	0.003	0.145***	0.162***	0.161***
	(0.055)	(0.007)	(0.011)	(0.013)	(0.014)
$brw_t \times \frac{Input}{Sales}_{it}$			0.007		
			(0.005)		
$brw_t \times \frac{Wage}{Sales}_{it}$				-0.104	
				(0.076)	
$brw_t \times \phi_{it}^{imp}$					-0.039
					(0.030)
$Debt_{it-n}$	-0.044***	-0.014**			
	(0.180)	(0.162)			
$\Delta \ln P_{it-1}$			0.299***	0.299***	0.299***
			(0.003)	(0.003)	(0.003)
$Sales_{it-n}$	0.081***	0.037***	-0.005***	-0.005***	-0.005***
	(0.262)	(0.187)	(0.001)	(0.001)	(0.001)
NER Control	No	No	Yes	Yes	Yes
Firm FE	Yes	Yes	Yes	Yes	Yes
Observations	155699	155699	917419	917419	917419

Notes: The specification in columns (1)-(2) is similar to Table 4. The specification in columns (3)-(5) is similar to Table 6. The dependent variables in columns (1)-(2) are changes in intermediate input cost over sales ratio and wage expense over sales ratio, respectively. ϕ^{imp} represents the import intensity, which is the firm-level ratio of imports to total material inputs. The standard errors here are clustered at the firm level, and the results are robust to clustering at the time level.

Table C5: China's bond index responses

	(1)	(2)	(3)	(4)
Period	2003-2006		2003-2022	
Price index	treasury bond	corporate bond	treasury bond	corporate bond
brw_t	-0.070	-0.381	-0.031*	-0.052
	(0.093)	(0.364)	(0.018)	(0.037)
Constant	Yes	Yes	Yes	Yes
Observations	27	25	137	135

Notes: The specification is $y_t = \alpha + \beta \cdot m_t + \varepsilon_t$, where y_t is the bond index overnight return (from last day's close price to today's open price), m_t is the daily BRW monetary policy shock, and t is the Fed FOMC announcement date. The heteroskedasticity-adjusted robust standard errors are used here. *, **, and *** indicate significance at 10%, 5%, and 1% levels.

C.2 Additional evidence on the borrowing cost channel

FDI VS non-FDI firms. Foreign direct investment (FDI) firms usually have better supply chain management capabilities and relative advantages in global risk hedging compared to purely domestic firms in China (e.g. Desai, Foley and Forbes (2008), Manova, Wei and Zhang (2015), Ding et al. (2024)); therefore, their liquidity conditions should be less affected by external adverse shocks. Accordingly, it is expected that their export prices will also be less influenced. This is verified in Table C6.

Table C6: FDI VS non-FDI firms

	(1)	(2)	(3)	(4)	(5)	(6)
Dependent Var	Monthly $\Delta \ln P_{it}$					
	Domestic		FDI		Comparison	
brw_t	0.220***	0.222***	0.166***	0.131***	0.219***	
	(0.024)	(0.025)	(0.012)	(0.011)	(0.025)	
$brw_t \times FDI$					-0.086***	-0.107***
					(0.027)	(0.027)
$Sales_{it-12}$		0.009***		-0.007***	-0.005***	-0.017***
		(0.003)		(0.001)	(0.001)	(0.001)
$\Delta \ln P_{it-1}$		0.185***		0.336***	0.299***	0.296***
		(0.005)		(0.003)	(0.003)	(0.003)
NER Control	Yes	Yes	Yes	Yes	Yes	No
Firm FE	Yes	Yes	Yes	Yes	Yes	Yes
Year-month FE	No	No	No	No	No	Yes
Observations	269743	210467	830657	706952	917419	917419

Notes: The samples in columns (1)-(2) and (3)-(4) include domestic firms and FDI firms, respectively. The interaction term in columns (5)-(6) is the FDI dummy variable, which takes a value of 1 for multinational firms or joint ventures and 0 for domestic Chinese firms, identified one year ago. All regressions include firm fixed effects. Column (6) additionally incorporates time-fixed (year-month pair) effects. Robust standard errors are clustered at the firm level. *, **, and *** indicate significance at 10%, 5%, and 1% levels.

Financial development of destinations. If a firm exports more to financially developed countries, it should have a smaller price change because these countries have better capacities to absorb liquidity contraction shocks and are less likely to pass through the adverse impacts to their trade partners. To test this hypothesis, we use the ratio of private credit to GDP (following Beck, Demirgüç-Kunt and Levine (2009) and Lin and Ye (2018*b*)) as an indicator of market-specific financial development and aggregate it to the firm level. The results are consistent with our guess, see Table C7.

Table C7: Financial development of export markets

	(1)	(2)	(3)	(4)	(5)	(6)
Dependent Var			Monthly $\Delta \ln P_{it}$			
	Exporters selling more to undeveloped markets		Exporters selling more to developed markets		Comparison	
brw_t	0.194***	0.181***	0.149***	0.122***	0.182***	
	(0.017)	(0.017)	(0.014)	(0.013)	(0.017)	
$brw_t \times \mathbf{1}\{fd_{it} > \bar{fd}_t\}$					-0.052**	-0.060***
					(0.021)	(0.021)
$Sales_{it-12}$		0.002		-0.009***	-0.005***	-0.017***
		(0.002)		(0.002)	(0.001)	(0.001)
$\Delta \ln P_{it-1}$		0.227***		0.338***	0.298***	0.295***
		(0.004)		(0.004)	(0.003)	(0.003)
NER Control	Yes	Yes	Yes	Yes	Yes	No
Firm FE	Yes	Yes	Yes	Yes	Yes	Yes
Year-month FE	No	No	No	No	No	Yes
Observations	484334	392014	610852	520009	912476	912476

Notes: We define the firm-level financial development indicator, which takes 1 if $fd_{it} > \bar{fd}_t$ (the median) and 0 otherwise. In columns (1)-(2), we limit our sample to firms with $fd_{it} \leq \bar{fd}_t$ (selling more to financially undeveloped markets). In columns (3)-(4), we limit our sample to firms with $fd_{it} > \bar{fd}_t$ (selling more to financially developed markets). In columns (5)-(6), we use the whole sample but additionally include the interaction term of monetary shock and the median dummy of the firm-level financial development indicator. All regressions include firm fixed effects. Column (6) additionally incorporates time-fixed (year-month pair) effects. Robust standard errors are clustered at the firm level. *, **, and *** indicate significance at 10%, 5%, and 1% levels.

Ordinary VS processing trade. Firms that participate in more processing trade usually have less borrowing needs and are less affected by credit conditions (Manova and Yu (2016)). A processing trader imports raw materials and intermediate inputs from a foreign firm for domestic processing and re-exports to the same firm as its customer. So, suppose the borrowing cost channel holds, we expect that, facing the same monetary policy shock, the price responses for processing traders should be smaller than those of ordinary traders. The conjecture is proved in Table C8.

Table C8: Ordinary trade vs processing trade

	(1)	(2)	(3)	(4)	(5)	(6)
Dependent Var			Monthly $\Delta \ln P_{it}$			
	Only ordinary trade		Only processing trade		Comparison	
brw_t	0.194***	0.181***	0.100***	0.071***	0.190***	
	(0.018)	(0.019)	(0.019)	(0.016)	(0.016)	
$brw_t \times process$					-0.088***	-0.102***
					(0.023)	(0.024)
$Sales_{it-12}$		-0.001		-0.011***	-0.005***	-0.017***
		(0.002)		(0.002)	(0.001)	(0.001)
$\Delta \ln P_{it-1}$		0.189***		0.473***	0.299***	0.296***
		(0.003)		(0.005)	(0.003)	(0.003)
NER Control	Yes	Yes	Yes	Yes	Yes	No
Firm FE	Yes	Yes	Yes	Yes	Yes	Yes
Year-month FE	No	No	No	No	No	Yes
Observations	499448	391356	283934	242572	917419	917419

Notes: In columns (1)-(2), we limit our sample to firms doing only ordinary trade. In columns (3)-(4), we limit our sample to firms doing only processing trade. In columns (5)-(6), we use the whole sample but additionally include the interaction term of monetary shock and the processing trade intensity. A higher value of *process* means a firm is more involved in processing trade. All regressions include firm fixed effects. Column (6) additionally incorporates time-fixed (year-month pair) effects. Robust standard errors are clustered at the firm level. *, **, and *** indicate significance at 10%, 5%, and 1% levels.

C.3 Alternative stories

Global demand shift. A tightening US monetary shock will induce a global recession, which might shift demand toward Chinese exports because some Chinese products are usually cheaper and of lower quality than goods from developed countries. The idea is similar to the concept of Giffen goods: prices of inferior goods increase when incomes drop. However, this could not explain why US import prices from many other countries also increase (see Picture 1). Besides, in Table C9, we find that the prices of homogeneous goods also increase, which have little differences in quality across producers and can not be treated as Giffen goods. Prices of homogeneous goods have risen even more, probably because these goods are generally more price elastic, so their prices are more likely to adjust in response to cost increases, as in Zhang (2022). Moreover, the quality of homogeneous goods is relatively stable over time, suggesting that our results are not driven by quality adjustment but by price change itself. Also, if monetary tightening causes a quality drop, this would reduce the price, which implies that our cost-side effect is even underestimated.

Table C9: Homogeneous good vs differentiated good

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Dependent Var				Monthly $\Delta \ln P_{it}$				
	Conservative classification				Liberal classification			
brw_t	0.177***	0.149***	0.156***	0.137***	0.175***	0.147***	0.155***	0.139***
	(0.011)	(0.011)	(0.012)	(0.012)	(0.011)	(0.011)	(0.013)	(0.012)
$brw_t \times ToE$	0.154	0.117			0.265***	0.243***		
	(0.129)	(0.126)			(0.086)	(0.082)		
$brw_t \times Ref$			0.209***	0.125***			0.167***	0.083***
			(0.033)	(0.032)			(0.031)	(0.030)
$Sales_{it-12}$		-0.005***		-0.005***		-0.005***		-0.005***
		(0.001)		(0.001)		(0.001)		(0.001)
$\Delta \ln P_{it-1}$		0.298***		0.298***		0.298***		0.298***
		(0.003)		(0.003)		(0.003)		(0.003)
NER Control	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Firm FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Observations	1014106	850165	1014106	850165	1014106	850165	1014106	850165

Notes: The specification is $\Delta \ln P_{it} = \alpha + \beta_1 \cdot m_t + \beta_2 \cdot m_t \cdot X_{it} + \Gamma \cdot \mathbf{Z} + \xi_i + \varepsilon_{it}$. The variables ToE and Ref represent the value share of goods traded on an organized exchange and the value share of reference-priced goods of firm i. Columns (1)-(4) use the "conservative" classification, while columns (5)-(8) use the "liberal" classification, both referring to Rauch (1999). $\mathbf{Z}$ denotes lagged controls of firm-level time-variant variables, including price changes in the previous month and real sales income in the previous year. All regressions include firm fixed effects. Robust standard errors are clustered at the firm level; *, **, and *** indicate significance at 10%, 5%, and 1% levels.

International competition. One may argue that the US monetary policy shocks will deteriorate international markets more than China, where there is capital control. In this way, Chinese exporters would be more competitive than their foreign competitors and thus have greater market power to raise prices. However, the markup responses in Table 7 rule out this argument because we find that exporters' markups even fall rather than rise. This suggests that Chinese exporters do not have greater bargaining power in pricing their products after the US tightening shock.

Exchange rate pass-through. Some may think that the US tightening could depreciate the Chinese currency and cause an increase in the Chinese RMB price. However, this can't explain why the price denominated in the US dollar also increases. Besides, during most of the sample period (from January 2000 to June 2005), China's exchange rate regime was fixed to the US dollar, so a US tightening shock will cause the RMB to appreciate against other currencies. This story means the RMB-denominated price will fall, contrary to our finding in Table B6. In addition, we have controlled the change of bilateral real exchange rate in the firm-product-country level regression (see Table B3). All results are robust and significant, which implies that exchange rate pass-through can not fully explain the impact of US monetary policy shocks on export prices.

D Model appendix

D.1 Preferences and demand

For ease of illustration and to maintain the generality of the results, we introduce a general multiple-country setting for the model to explain the documented empirical evidence. Source and destination countries are denoted by i and j, respectively. In this paper, i is China, and j denotes another country. It is assumed that a representative consumer in country j has preferences over consumption of locally produced goods Y_j^h and foreign products Y_j, and $U = U(Y_j^h, Y_j)$, where $U(\cdot)$ satisfies the standard regularity conditions. The import bundle aggregates products from all countries:

$$Y_j = \left(\int Y_{ij}^{\frac{\sigma-1}{\sigma}} di \right)^{\frac{\sigma}{\sigma-1}}$$

while each bilateral import flow Y_{ij} includes a continuum of unique products $\omega \in [0, 1]$:

$$Y_{ij} = \left(\int Y_{ij}(\omega)^{\frac{\sigma-1}{\sigma}} d\omega \right)^{\frac{\sigma}{\sigma-1}}$$

where $Y_{ij}(\omega)$ is country j's consumed quantity of variety ω originated from country i, and $\sigma > 1$ is the elasticity of substitution between varieties. Therefore, consumer optimization yields the following demand function for variety ω:

$$Y_{ij}(\omega) = \frac{p_{ij}(\omega)^{-\sigma}}{P_j^{-\sigma}} Y_j$$

where $p_{ij}(\omega)$ is the price of the variety ω, $P_j = (\int p_{ij}^{1-\sigma} di)^{\frac{1}{1-\sigma}}$ is the import price index of country j, which is the aggregate of import prices $P_{ij} = (\int p_{ij}(\omega)^{1-\sigma} d\omega)^{\frac{1}{1-\sigma}}$ across all other countries.

D.2 Firm's problem

The firm in country i minimizes its cost to satisfy the demand in the country j, $Y_{ij}(\omega) = \frac{p_{ij}(\omega)^{-\sigma}}{P_j^{-\sigma}} Y_j$, which yields the cost function $C_{ij} = \frac{w_i(1-\delta_i+\delta_i R_i)}{\phi_i} \frac{p_{ij}(\omega)^{-\sigma}}{P_j^{-\sigma}} Y_j$, where $R_i > 1$ is the gross borrowing interest rate in country i. We explicitly assume $R_i = \bar{R}_i + \rho_R^i m + \epsilon_R^i$ where $\bar{R}_i$ is a trend component of R_i, $\rho_R^i >= 0$ and ϵ_R^i is a random error.[58] Also, to allow the non-linear elasticity of cost with respect to interest rate (which can be generated by other financial costs associated with borrowing), we replace R_i with R_i^α, where α is a constant and represents the elasticity of financial costs with respect to the interest rate. Following the convention, we also add an iceberg trade cost such that $\tau_{ij} \geq 1$ units of good must be shipped from country i for one unit to arrive at j. For simplicity of notation, the subscripts for source and destination and the index for variety are omitted. Thus, the new cost function is:

$$C = \frac{\tau w(1 - \delta + \delta R^\alpha)}{\phi} \frac{p^{-\sigma}}{P^{-\sigma}} Y$$

The optimization problem of the firm is:

$$\max_p \; (p - \frac{\tau w(1 - \delta + \delta R^\alpha)}{\phi}) \frac{p^{-\sigma}}{P^{-\sigma}} Y$$

Solving the unconstrained optimization problem will give us the optimal price:

$$p = \frac{\sigma}{\sigma - 1} \frac{\tau w[c^\gamma + (1 - c^\gamma) R^\alpha]}{\phi}$$

According to the price equation, we can see that the optimal price is jointly determined by both the markup and the marginal cost. The marginal cost is affected by the input price w, liquidity condition c, and the borrowing rate R.

[58] The literature (like Georgiadis (2016) and Miranda-Agrippino and Rey (2020), etc.) has reached an agreement that the borrowing interest rate in many countries will, in general, increase after a contractionary US monetary policy shock. Nevertheless, the response in China is insignificant due to capital control. Here, we assume that R_i is exogenous, and the firm takes it as given. The endogenization of interest rates will not change our main results.

D.3 Propositions

In the partial equilibrium model, the firm-level optimal export prices are affected by input prices, liquidity conditions, and borrowing interest rates, which in turn are affected by monetary policy shocks. With the expressions of export prices in hand, we could derive three propositions:

Proposition 1. *The export price decreases with liquidity conditions and increases with the borrowing interest rates:* $\frac{\partial p}{\partial c} < 0$, $\frac{\partial p}{\partial R} > 0$.

This proposition represents the relationship between export prices and the characteristics of the firms. Consequently, the impact of global monetary policy on export pricing is determined by how the shock could affect these intermediate variables. A tighter liquidity condition and a higher borrowing rate would cause a larger borrowing cost, thus implying a higher price.

Proposition 2. *The export price would increase in response to a tightening US monetary policy shock (that is,* $\frac{\partial p}{\partial m} > 0$*) if the supply side effect dominates.*

This is what we have revealed in the baseline findings in Section 4. Theoretically, a tightening US monetary shock would deteriorate firms' liquidity conditions and increase the borrowing rate. Therefore, according to Proposition 1, it is straightforward to understand why the export price would increase following a tightening monetary policy shock. Empirically, in Section 5, we show that exporters' liquidity conditions worsen significantly in reaction to a tightening shock. Meanwhile, the interest rate responses to US monetary shocks are not significant in China. Nevertheless, this impact is theoretically possible, especially for countries with less rigorous capital control.

On the demand side, a tightening shock may depress aggregate demand and decrease input prices, which would contribute to a negative effect on export prices. However, as long as this force is less powerful than that of the supply side, the net impact will be positive, which is consistent with our empirical finding. In theory, in the general equilibrium, according to international macro literature (e.g. Clarida, Galı and Gertler (2002)), the domestic input price can be written as a function of domestic and foreign output. The elasticity of the input price to output depends on some structural parameters, such as the intertemporal substitution of consumption and the elasticity of substitution between home and foreign goods. In the data, people find that the US tightening usually harms global output (e.g. Kim (2001), Maćkowiak (2007), Bluedorn and Bowdler (2011), and Georgiadis (2016)).

It is worth noting that although we emphasize the finding of a cost channel, we don't deny the existence of the demand side impact. Sometimes the demand side effect will likely be dominant. The determinants of the relative importance of demand-side versus supply-side impact are beyond the scope of this paper, which is left for future quantitative investigation.

Proposition 3. *The impact of the US monetary shock on export price (i.e., $\frac{\partial p}{\partial m}$) depends on the financial conditions of the firms. If supply-side factors dominate, it is greater when the firms' liquidity conditions (c) are worse and their average borrowing costs (δR) are higher, given some sets of parameters.*

This proposition sheds light on the role of firms' financing conditions in the transmission of the US monetary shocks to firms' export prices and reinforces the cost-driven channel. Our model shows that the impact of monetary shocks is heterogeneous at the firm level: firms with tighter liquidity conditions and higher average borrowing costs experience greater price increases in response to tightening shocks. The empirical supporting evidence is displayed in Section 5. We don't have a direct measure of the borrowing rate R for each firm in the data, and our measurements of average borrowing cost are proxies for δR.

The proof of the above proposition is shown in the Appendix D.4. Our conclusion in the benchmark model is robust to incorporate two input factors, credit constraints, dynamic and sticky price settings, and different ways of currency invoicing. All extensions are displayed in the Appendix D.5.

D.4 Proof of propositions

Proposition 1. The export price decreases with liquidity conditions and increases with the borrowing interest rates: $\frac{\partial p}{\partial c} < 0$, $\frac{\partial p}{\partial R} > 0$.

Proof

$$\frac{\partial p}{\partial c} = \frac{\sigma}{\sigma - 1}\frac{\tau w}{\phi}\gamma(1 - R^{\alpha})c^{\gamma - 1} < 0$$
$$\frac{\partial p}{\partial R} = \frac{\sigma}{\sigma - 1}\frac{\tau w}{\phi}[\alpha(1 - c^{\gamma})R^{\alpha - 1}] > 0$$

Proposition 2. The export price would increase in response to a tightening monetary shock (i.e.,$\frac{\partial p}{\partial m} > 0$) if the supply side effect dominates.

Proof

$$\begin{aligned}\frac{\partial p}{\partial m} =& \frac{\partial p}{\partial c}\frac{\partial c}{\partial m} + \frac{\partial p}{\partial R}\frac{\partial R}{\partial m} + \frac{\partial p}{\partial w}\frac{\partial w}{\partial m} \\ =& \frac{\sigma}{\sigma - 1}\frac{\tau w}{\phi}\gamma(1 - R^{\alpha})c^{\gamma - 1}\rho_c + \frac{\sigma}{\sigma - 1}\frac{\tau w}{\phi}[\alpha(1 - c^{\gamma})R^{\alpha - 1}]\rho_R + \\ & \frac{\sigma}{\sigma - 1}\frac{\tau}{\phi}[c^{\gamma} + (1 - c^{\gamma})R^{\alpha}]\rho_w\end{aligned}$$

The former two parts $\frac{\partial p}{\partial c}\frac{\partial c}{\partial m}$ are related to the supply-side effect, and the last part $\frac{\partial p}{\partial R}\frac{\partial R}{\partial m}$ reflects the power of demand shrink. When the supply-side cost-push effect dominates the demand effect, the net impact of the US monetary policy shock should be positive. This prediction is verified in the empirical part.

Proposition 3. The impact of the US monetary shock on export price (i.e., $\frac{\partial p}{\partial m}$) depends on the financial conditions of the firms. If supply-side factors dominate, it is greater when the firms' liquidity conditions (c) are worse and their average borrowing costs (δR) are higher, given some sets of parameters.

Proof

From Proposition 2, we can know that $\frac{\partial p}{\partial m}$ is a function of c and R. Given some sets of parameters (e.g., when $\gamma < 1$ and $\alpha > 1$), this value will decrease with c and increase with δR. The existence of this effect has been verified empirically in the mechanism part.

D.5 Model extension

Our conclusion in the benchmark model is robust to use two factors, binding credit constraint, sticky price setting, and different currency invoicing.

1. Two factors

Suppose we additionally include capital as an input factor. Then, we assume the production function of the firm is a Cobb–Douglas type $y = \phi K^{\chi} L^{1-\chi}$, where K is capital with a rental rate of r, χ is the share of income for capital.

In this case, the associated marginal cost becomes:

$$MC = (\frac{1}{\chi})^{\chi}(\frac{1}{1-\chi})^{1-\chi}\frac{\tau r^{\chi}\left[w(1-\delta+\delta R)\right]^{1-\chi}}{\phi}$$

The form of the optimization problem is unchanged with the one-factor model, and the optimal price is:

$$p = \frac{\sigma}{\sigma-1}MC = \frac{\sigma}{\sigma-1}(\frac{1}{\chi})^{\chi}(\frac{1}{1-\chi})^{1-\chi}\frac{\tau r^{\chi}\left[w(1-\delta+\delta R)\right]^{1-\chi}}{\phi}$$

In this case, monetary shocks can also affect the price through the rental rate r. In the general equilibrium, r and w will be jointly determined to clear the corresponding factor markets. Facing a monetary tightening from the US, the role of liquidity in affecting borrowing proportion δ is mixed with other factors, and the net effects on export price depend on the relative strength of each force. As long as other effects are dominated by the cost-side impact, the export price will increase.

2. Credit constraints

As is documented in the literature (see Manova and Zhang (2012), Manova (2013), Manova, Wei and Zhang (2015)), the choice of export prices sometimes depends on firms' credit constraints.[59] For those firms with unbinding credit constraints, export prices are equal to the marginal cost times markup, and the responses of prices are determined by these two components, as we discussed in Section 5.4. Then, how about the reactions of those binding firms?

[59] Manova and Zhang (2012) finds that many exporters are constrained by credit limits. This is even more prominent than for non-exporters, as exporting activities are usually more demanding than domestic business.

The worsening credit conditions induced by the tightening of the US monetary shock would also cause an increase in export prices, and this is consistent with the efficiency sorting theory (regarding efficiency sorting, see Manova and Zhang (2012) for more details). More specifically, a tightening US monetary policy shock would reduce an exporter's liquidity conditions (as is proved in Section 5.2), which will accordingly increase firms' credit demand. To satisfy the requirement for credit needs, the exporters are forced to increase prices and obtain more cash flow (relative to the funds in need) to improve credit access.[60] That's because firms with better cash flow are less likely to default and thus could obtain more credit access from the financial markets.

This intuition can be illustrated in a simple model. In addition to the working capital constraint, here we also assume that firms cannot borrow more than a fraction θ of the expected cash flow from exporting. The optimization problem of the firm is:

$$\max_{p} \ (p - \frac{\tau w(1-\delta+\delta R^{\alpha})}{\phi})\frac{p^{-\sigma}}{P^{-\sigma}}Y$$

$$\text{s.t. } \theta\frac{p^{1-\sigma}}{P^{-\sigma}}Y \geq (1-c^{\gamma})\frac{\tau w}{\phi}\frac{p^{-\sigma}}{P^{-\sigma}}Y$$

If the credit constraint is binding, we can rewrite the credit constraint as:

$$p = \frac{1-c^{\gamma}}{\theta}\frac{\tau w}{\phi}$$

In this case, the firm-level optimal export prices are affected by both liquidity condition c and collateral ratio θ, which in turn are impacted by monetary policy shocks. Intuitively, monetary policy shocks could increase firms' credit needs by worsening the liquidity conditions c and harm their credit access by reducing the collateral ratio θ, thus motivating them to raise prices to get more cash flow to meet the credit requirements.

It is worth noting that empirically identifying which firms have binding credit constraints is quite hard, and the constraints may be occasionally binding for a firm. In this paper, we only qualitatively explore how the binding credit constraint would affect firms' pricing decisions and leave the quantitative analysis for future study.

3. Dynamic optimization and sticky price

[60] The price increase may reduce the quantity, thus, the impact on total cash flow is uncertain. However, the drop in quantity will also decrease the money for the input costs, which is symmetric to the sales income. So, raising the price only increases the relative cash flow instead of the total cash flow.

In the benchmark model, the optimization problem is static, and the prices are assumed to be flexible. In this part, we are going to illustrate that the main mechanisms still hold under dynamic optimization and sticky prices. We use the classical Calvo (1983) sticky price setting, and the firm's problem is to maximize its expected real profits:

$$\max_{p_t} \; \mathbb{E}_t \sum_{i=0}^{\infty} \lambda^i \Omega_{t,t+i} \left[\frac{p_t}{P_{t+i}} - \frac{\tau_{t+i} w_{t+i} (1 - \delta_{t+i} + \delta_{t+i} R_{t+i}^{\alpha})}{\phi_{t+i} P_{t+i}} \right] \frac{p_t^{-\sigma}}{P_{t+i}^{-\sigma}} Y_{t+i}$$

where $\Omega_{t,t+i}$ is the real stochastic discount factor, and λ is the probability of a firm keeping its price unchanged in each period. We solve the unconstrained problem and get the first-order condition:

$$\mathbb{E}_t \sum_{i=0}^{\infty} \lambda^i \Omega_{t,t+i} \left[(1-\sigma) \frac{p_t}{P_{t+i}} + \sigma \varphi_{t+i} \right] \frac{1}{p_t} \frac{p_t^{-\sigma}}{P_{t+i}^{-\sigma}} Y_{t+i} = 0$$

where $\varphi_{t+i} \equiv \frac{\tau_{t+i} w_{t+i} (1-\delta_{t+i}+\delta_{t+i} R_{t+i}^{\alpha})}{\phi_{t+i} P_{t+i}}$ is the real marginal cost in $t+i$. The optimal price can be expressed as:

$$p_t = \frac{\sigma}{\sigma - 1} \frac{\mathbb{E}_t \sum_{i=0}^{\infty} \lambda^i \Omega_{t,t+i} \frac{P_t^{-\sigma}}{P_{t+i}^{-\sigma}} Y_{t+i} \varphi_{t+i}}{\mathbb{E}_t \sum_{i=0}^{\infty} \lambda^i \Omega_{t,t+i} \frac{P_t^{-\sigma}}{P_{t+i}^{1-\sigma}} Y_{t+i}}$$

We can see that a tightening shock can increase the borrowing proportion δ_{t+i}, the borrowing interest rate R_{t+i}, and hence the marginal cost φ_{t+i}. The channel is similar to the static problem we discussed previously, and now the impact is a weighted sum of the effect on current and future marginal costs. However, in this case, the impact of the monetary shock on the discount factor $\Omega_{t,t+i}$, aggregate expenditure Y_{t+i} and price index P_{t+i} will also play a role, which reflects the power of general equilibrium. Besides, this effect is higher if the price stickiness is lower. When $\lambda = 0$, $p_t = \frac{\sigma}{\sigma-1} \frac{\tau_t w_t (1-\delta_t+\delta_t R_t^{\alpha})}{\phi_t}$, which is exactly the same as the static version.

If we consider the role of credit constraint, the firm's problem is to maximize its expected real profits:

$$\max_{p_t} \; \mathbb{E}_t \sum_{i=0}^{\infty} \lambda^i \Omega_{t,t+i} \left[\frac{p_t}{P_{t+i}} - \frac{\tau_{t+i} w_{t+i} (1 - \delta_{t+i} + \delta_{t+i} R_{t+i}^{\alpha})}{\phi_{t+i} P_{t+i}} \right] \frac{p_t^{-\sigma}}{P_{t+i}^{-\sigma}} Y_{t+i}$$

$$\text{s.t. } \mathbb{E}_t \sum_{i=0}^{\infty} \lambda^i \Omega_{t,t+i} \frac{P_t}{P_{t+i}} \theta_{t+i} \frac{p_t^{1-\sigma}}{P_{t+i}^{-\sigma}} Y_{t+i} \geq \mathbb{E}_t \sum_{i=0}^{\infty} \lambda^i \Omega_{t,t+i} \frac{P_t}{P_{t+i}} \left[(1 - c_{t+i}^{\gamma}) \frac{\tau_{t+i} w_{t+i}}{\phi_{t+i}} \frac{p_t^{-\sigma}}{P_{t+i}^{-\sigma}} Y_{t+i} \right]$$

where θ_{t+i} is the collateral ratio in period $t+i$. The left-hand side of the borrowing constraint is the weighted sum of credit access, and the right-hand side reflects the corresponding external credit demands. When the borrowing constraint is binding, we can rearrange the constraint and obtain the expression of the export price:

$$p_t = \frac{\mathbb{E}_t \sum_{i=0}^{\infty} \lambda^i \Omega_{t,t+i} \frac{Y_{t+i}}{P_{t+i}^{1-\sigma}} \frac{\tau_{t+i} w_{t+i}}{\phi_{t+i}} \delta_{t+i}}{\mathbb{E}_t \sum_{i=0}^{\infty} \lambda^i \Omega_{t,t+i} \frac{Y_{t+i}}{P_{t+i}^{1-\sigma}} \theta_{t+i}}$$

It is seen that a tightening monetary shock can raise prices by increasing the borrowing proportion δ_{t+i} and reducing the ratio of credit access θ_{t+i}. The channel is similar to the static problem we discussed before. However, in this case, the price is also affected by the expectation of future liquidity and collateral conditions, and the impact of the monetary shock on the discount factor $\Omega_{t,t+i}$, aggregate expenditure Y_{t+i} and price index P_{t+i} will also play a role, which reflects the power of general equilibrium. If $\lambda = 0$, $p_t = \frac{1-c_t^{\gamma}}{\theta_t} \frac{\tau_t w_t}{\phi_t}$, which is identical to the static solution.

4. Invoicing currency

Our benchmark model doesn't explicitly consider the role of exchange rates and invoicing currency. In this part, we will explain that our mechanisms are robust in three different cases: producer currency pricing (PCP), US dollar currency pricing (DCP), and local currency pricing (LCP).

(1) *Producer Currency Pricing (PCP)*

The firm's problem is:

$$\max_{p_t} \ \mathbb{E}_t \sum_{i=0}^{\infty} \lambda^i \Omega_{t,t+i} \left[\frac{p_t}{P_{t+i}} - \frac{\tau_{t+i} w_{t+i} (1 - \delta_{t+i} + \delta_{t+i} R_{t+i}^{\alpha})}{\phi_{t+i} P_{t+i}} \right] \left(\frac{p_t}{e_{t+i}^j P_{t+i}^j} \right)^{-\sigma} Y_{t+i}^j$$

where p is the price in the producer currency, e_j is the nominal exchange rate, defined as the price of the destination country j's currency in terms of the producer currency, P and P^j is the price index in the producer country and country j respectively, and Y^j is the total import in country j. The first order condition and optimal price are:

$$\mathbb{E}_t \sum_{i=0}^{\infty} \lambda^i \Omega_{t,t+i} \left[(1-\sigma) \frac{p_t}{P_{t+i}} + \sigma \varphi_{t+i} \right] \frac{1}{p_t} \frac{p_t^{-\sigma}}{(P_{t+i}^j)^{-\sigma}} Y_{t+i}^j (e_{t+i}^j)^{\sigma} = 0$$

$$p_t = \frac{\sigma}{\sigma-1}\frac{\mathbb{E}_t \sum_{i=0}^{\infty} \lambda^i \Omega_{t,t+i} \frac{P_t^{-\sigma}}{(P_{t+i}^j)^{-\sigma}} Y_{t+i}^j \varphi_{t+i} (e_{t+i}^j)^{\sigma}}{\mathbb{E}_t \sum_{i=0}^{\infty} \lambda^i \Omega_{t,t+i} \frac{1}{P_{t+i}} \frac{P_t^{-\sigma}}{(P_{t+i}^j)^{-\sigma}} Y_{t+i}^j (e_{t+i}^j)^{\sigma}}$$

We can see that a US monetary policy shock can still affect export prices through its impact on current and future real marginal costs φ_{t+i}. However, in this case, the export price is also affected by the bilateral exchange rate e^j and the price index in both the sourcing and destination countries. If $\lambda = 0$, $p_t = \frac{\sigma}{\sigma-1}\frac{\tau_t w_t (1-\delta_t+\delta_t R_t^{\alpha})}{\phi_t}$, which is exactly the same as the static version.

(2) ***US Dollar Currency Pricing (DCP)***

The firm's problem is:

$$\max_{p_t} \; \mathbb{E}_t \sum_{i=0}^{\infty} \lambda^i \Omega_{t,t+i} \left[\frac{p_t e_{t+j}^{us}}{P_{t+i}} - \frac{\tau_{t+i} w_{t+i} (1-\delta_{t+i}+\delta_{t+i} R_{t+i}^{\alpha})}{\phi_{t+i} P_{t+i}} \right] \left(\frac{p_t e_{t+i}^{us}}{e_{t+i}^j P_{t+i}^j} \right)^{-\sigma} Y_{t+i}^j$$

where p is the price in the US dollar, e^{us} is the nominal exchange rate against the US, defined as the price of the US dollar in terms of the producer currency. The first order condition and optimal price are:

$$\mathbb{E}_t \sum_{i=0}^{\infty} \lambda^i \Omega_{t,t+i} \left[(1-\sigma) \frac{p_t e_{t+i}^{us}}{P_{t+i}} + \sigma \varphi_{t+i} \right] \frac{1}{p_t} \frac{p_t^{-\sigma}}{(P_{t+i}^j)^{-\sigma}} Y_{t+i}^j (e_{t+i}^j / e_{t+i}^{us})^{\sigma} = 0$$

$$p_t = \frac{\sigma}{\sigma-1}\frac{\mathbb{E}_t \sum_{i=0}^{\infty} \lambda^i \Omega_{t,t+i} \frac{P_t^{-\sigma}}{(P_{t+i}^j)^{-\sigma}} Y_{t+i}^j \varphi_{t+i} (e_{t+i}^j / e_{t+i}^{us})^{\sigma}}{\mathbb{E}_t \sum_{i=0}^{\infty} \lambda^i \Omega_{t,t+i} \frac{1}{P_{t+i}} \frac{P_t^{-\sigma}}{(P_{t+i}^j)^{-\sigma}} Y_{t+i}^j (e_{t+i}^j / e_{t+i}^{us})^{\sigma} e_{t+i}^{us}}$$

The export price here is affected by both the bilateral exchange rate e^j and the US exchange rate e^{us}. If $\lambda = 0$, $p_t e_t^{us} = \frac{\sigma}{\sigma-1}\frac{\tau_t w_t (1-\delta_t+\delta_t R_t^{\alpha})}{\phi_t}$, the price in terms of home currency (here RMB) is identical to the PCP version.

(3) ***Local Currency Pricing (LCP)***

The firm's problem is:

$$\max_{p_t} \; \mathbb{E}_t \sum_{i=0}^{\infty} \lambda^i \Omega_{t,t+i} \left[\frac{p_t e_{t+j}^{j}}{P_{t+i}} - \frac{\tau_{t+i} w_{t+i} (1-\delta_{t+i}+\delta_{t+i} R_{t+i}^{\alpha})}{\phi_{t+i} P_{t+i}} \right] \left(\frac{p_t}{P_{t+i}^j} \right)^{-\sigma} Y_{t+i}^j$$

The first order condition and optimal price are:

$$\mathbb{E}_t \sum_{i=0}^{\infty} \lambda^i \Omega_{t,t+i} \left[(1-\sigma) \frac{p_t e^j_{t+i}}{P_{t+i}} + \sigma \varphi_{t+i} \right] \frac{1}{p_t} \frac{p_t^{-\sigma}}{(P^j_{t+i})^{-\sigma}} Y^j_{t+i} = 0$$

$$p_t = \frac{\sigma}{\sigma - 1} \frac{\mathbb{E}_t \sum_{i=0}^{\infty} \lambda^i \Omega_{t,t+i} \frac{P_t^{-\sigma}}{(P^j_{t+i})^{-\sigma}} Y^j_{t+i} \varphi_{t+i}}{\mathbb{E}_t \sum_{i=0}^{\infty} \lambda^i \Omega_{t,t+i} \frac{1}{P_{t+i}} \frac{P_t^{-\sigma}}{(P^j_{t+i})^{-\sigma}} Y^j_{t+i} e^j_{t+i}}$$

The export price is also affected by the bilateral exchange rate e^j, but slightly different from the PCP and DCP case. when $\lambda = 0$, $p_t e^j_t = \frac{\sigma}{\sigma-1} \frac{\tau_t w_t (1-\delta_t+\delta_t R^\alpha_t)}{\phi_t}$, the price in terms of home currency (here RMB) is identical to the PCP version.