A Tale of Two Sectors: Product Differentiation and Heterogeneity in Price Stickiness in a General Equilibrium Model*

Jorge Salas†

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Abstract

International trade data show that the prices of differentiated goods barely moved in recent recessions, while the declines in volumes were large. In contrast, for nondifferentiated goods, price reductions greatly contributed to the adjustment in trade values. I document similar patterns in U.S. domestic data on producer prices and industrial production. I use a two-country general equilibrium model with trade in nondifferentiated and differentiated manufactured goods to shed light on the reasons behind these sectoral differences. I focus on two mechanisms: sector-specific nominal rigidities and endogenous variable markups at the producer level. The calibration of the key parameters of the model is based on micro data and national accounts data. The impulse responses of relative prices and quantities to a monetary shock are compared with empirical vector autoregressions, showing a reasonably good match. These responses can largely be explained by heterogeneity in the frequency of price adjustment, while the variable markup channel is quantitatively less important.

JEL Codes: E3, E52, F14, F41, L11, L16

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†Department of Economics; University of Maryland, College Park. Email: salas@econ.umd.edu.
1 Introduction

A consistent finding across empirical studies in the closed and open economy literature is that relatively raw products or “nondifferentiated goods” exhibit a smaller degree of price stickiness than more processed products or “differentiated goods” (Gordon, 1990; Bils and Klenow, 2004; Gopinath and Rigobon, 2008; Nakamura and Steinsson, 2008). This evidence suggests that, all else equal, a conventional monetary model with sticky prices would predict: first, a larger decline in prices of nondifferentiated goods relative to those of differentiated goods following a contractionary demand shock, and second, a reduction in the relative quantities consumed/produced of differentiated goods to nondifferentiated goods (see e.g. Bils et al., 2003). An implication of these theoretical predictions is that, under some qualifications, the heterogeneity in price stickiness across sectors could help explain the evidence of larger declines among trade quantities of differentiated goods than among nondifferentiated goods during recent recessions (Haddad et al., 2010; Gopinath et al., 2012).1

Motivated by the above discussion, this paper emphasizes the dichotomy between sectors producing nondifferentiated and differentiated manufactured goods, and investigates how the relative prices and quantities across these sectors respond to an explicitly identified aggregate demand shock. In the empirical section, I use monthly U.S. data at the industry level to construct measures of domestic and trade prices and quantities for nondifferentiated and differentiated goods. I then estimate vector autoregressions to generate impulse response functions of relative prices and quantities across sectors to a contractionary monetary policy shock. The theoretical section of the paper uses an open economy general equilibrium model calibrated to U.S. data to shed light on the extent to which two potential sources of heterogeneity in price stickiness—sector-specific nominal

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1See e.g. Alessandria et al. (2010), Eaton et al. (2011), and the survey in Bems et al. (2013) for theories on the so-called Great Trade Collapse of 2008–2009. None of them considers the empirically relevant distinction between nondifferentiated and differentiated goods in an explicit way, nor do they model heterogeneity in price stickiness.
rigidities and endogenous variable markups—can rationalize quantitatively the sectoral differences in responses to a monetary shock observed in the data.

The empirical section of the paper exploits the Rauch (1999) classification to categorize disaggregated industries as producers of nondifferentiated and differentiated goods. Rauch (1999) distinguishes products traded on organized exchanges, goods with reference prices (quoted in trade publications), and differentiated goods or “branded” products without a reference price. Armed with this classification, I compute aggregated measures of domestic prices and quantities (using producer price indices and industrial production indices), and trade prices and quantities (using at-the-dock data on unit values and volumes of exports and imports) for nondifferentiated and differentiated goods sectors. The empirical findings validate the aforementioned intuition that quantities adjust the least in the nondifferentiated goods sector, where prices are most responsive. I show that this result holds for domestic and trade data during recessions. A similar result holds for impulse responses to a monetary policy shock. This shock is identified in vector autoregressions by means of the recursiveness assumption used by Christiano et al. (1999).

To explain the empirical evidence, I introduce heterogeneity in price stickiness in a two-country multisector model that includes an aggregate nominal demand shock. This framework features manufacturing firms operating under monopolistic competition in two tradable sectors—nondifferentiated and differentiated goods—and a nontradable goods sector. The production sectors coexist with perfectly competitive retail and distribution sectors. Nominal rigidities among the monopolistically competitive producers take the form of time-dependent Calvo price-setting. I consider differences in the frequency of price changes across sectors, in line with microeconomic evidence on price setting. In addition, following Corsetti and Dedola (2005), retail prices include a distribution-cost component. As a result, the producers’ price elasticities of demand, and hence their optimal markups, are endogenous and depend on the cost of production relative to the local distribution costs. By allowing for different intensities in the use of distribution services across sectors,
the variable markups constitute a second source of heterogeneity in price stickiness (on top of the nominal rigidities). My modeling choices for nominal rigidities and variable markups facilitate tractability and allow for a straightforward calibration.\textsuperscript{2}

Using input-output data for the U.S., I document that the share of distribution costs in the retail price is higher for differentiated goods than for nondifferentiated goods. This result is intuitive since distribution costs include, for example, marketing and advertising services. The model then implies that markups are more variable among differentiated goods relative to nondifferentiated goods; therefore, the prices of the former are less sensitive to changes in marginal costs, but more sensitive to changes in distribution costs, than those of the latter. The calibration of the distribution-cost shares also generates a lower long-run exchange rate pass-through into import prices of differentiated goods relative to nondifferentiated goods, as observed in the data (Gopinath and Itskhoki, 2010).\textsuperscript{3}

To examine the model’s quantitative predictions, I compute impulse responses of sectoral domestic and trade variables to a tightening monetary shock. The shock is correlated across countries and generates endogenous movements in the exchange rate. The main results are as follows. First, the model-based impulse responses of the relative prices and quantities across nondifferentiated and differentiated goods match the VAR-based impulse responses reasonably well, in terms of sign and magnitude. Second, from a quantitative perspective, the key mechanism that drives the model’s predictions on sectoral prices and quantities is the nominal rigidity channel, that is, the differences in durations of producer prices across sectors. This second result is driven by the calibrated differences in price durations across sectors, which imply that nondifferentiated goods producers reset prices

\textsuperscript{2}In the model, the two tradable sectors also differ in their elasticity of substitution between varieties. This difference is not quantitatively important for the main results.

\textsuperscript{3}For evidence on variable markups at the producer (or wholesale) level, see Gopinath and Itskhoki (2011). The recent empirical literature on (incomplete) exchange rate pass-through that uses firm-level data highlights the role of variable markups in offsetting the effect of exchange rate movements on trade prices at the dock (De Loecker et al., 2012; Amiti et al., 2014; Fitzgerald and Haller, 2014). My model abstracts from other factors that are typically invoked in that literature, such as decreasing returns to scale and imported inputs (see e.g. Goldberg and Hellerstein, 2008).
every 3 months (on average), whereas differentiated goods producers reset prices every 14 months. These calibration targets rely on recent micro-data evidence for the U.S. I find a relatively less important role for the variable markup channel.

**Related Literature.** This paper is related to the existing empirical literature on price rigidities and the sectoral effects of monetary policy shocks in a closed economy (see e.g. Bils et al., 2003; Balke and Wynne, 2007; Boivin et al., 2009; Baumeister et al., 2013; and Kaufmann and Lein, 2013). It is more closely related to Balke and Wynne (2007) and Boivin et al. (2009) in that they also study data on producer prices, as opposed to consumer prices. As in Boivin et al. (2009) and Baumeister et al. (2013), I find evidence that quantities react the least in sectors where prices react the most. My work differs in that I explicitly distinguish between nondifferentiated and differentiated goods, and in that I analyze international trade prices in addition to domestic prices.

Other papers have studied the sectoral implications of heterogeneity in price setting using sticky-price multi-sector monetary models (Bouakez et al., 2009, 2014; Carvalho and Lee, 2011).\(^4\) Employing a less restrictive framework than mine (with additional features such as input-output production linkages), Bouakez et al. (2014) conclude that heterogeneity in price stickiness is the primary factor explaining the differences in responses of sectoral prices and outputs to a monetary shock. In the same vein, my results indicate that heterogeneity in the sectoral frequency of price adjustment generates quantitatively important differences in the responses of prices and quantities across the analyzed sectors. None of those papers, however, focuses on the distinction between nondifferentiated and differentiated goods, nor do they consider variable markups based on additive distribution costs. Further, those papers do not consider predictions for international trade variables.

In parallel work, Bergin and Corsetti (2014) also construct and calibrate a two-country

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\(^4\)Previous studies distinguish between durable and nondurable goods (Barsky et al., 2007; Erceg and Levin, 2006). Another related strand of literature focuses on the role of price heterogeneity in *aggregate* dynamics, such as Carvalho (2006) who studies the real effects of monetary policy, and in an open economy context, Carvalho and Nechio (2011) who study the real exchange rate.
general equilibrium monetary model with trade in nondifferentiated and differentiated goods. In their model, the nondifferentiated goods sector produces commodity-type products and operates under perfect competition, while manufacturing firms are located exclusively in the (monopolistically competitive) differentiated goods sector. They also incorporate sunk entry costs in the differentiated goods sector, as well as sectoral productivity shocks. Their goal is to show that monetary policy affects the pattern of country specialization between nondifferentiated and differentiated goods.\footnote{My model also shares some features with Cravino (2014), in that he incorporates nominal rigidities and variable markups in an open economy setting. However, Cravino considers a single tradable sector and focuses on the effects of exchange rate movements on aggregate productivity.}

By modeling price rigidities à la Calvo, I do not endogenize the reason why the frequency of price adjustment is higher for nondifferentiated than for differentiated goods. Other papers offer insights on this point. Neiman (2011) considers a partial equilibrium model of trade that endogenizes the relationship between elasticity of demand and price rigidity in a menu-cost setting. Gordon (1990) suggests that the prices of more differentiated goods embody relatively large amounts of labor, implying a role for wage rigidity.\footnote{Peneva (2011) provides formal empirical support for the idea that the degree of labor intensity and the degree of price flexibility are negatively correlated across industries in the U.S.}

Finally, Boivin et al. (2009) find empirically that a higher degree of price flexibility following a monetary shock is explained by a lower degree of sectoral market power, as measured by gross profit rates, and by a larger volatility of sector-specific shocks. In the context of my paper, we would indeed expect nondifferentiated goods to exhibit both relatively low market power, due to a higher elasticity of substitution, and relatively volatile idiosyncratic shocks due to their high commodity content.

**Organization.** The remainder of the paper is organized as follows. Section 2 documents empirical evidence on the behavior of U.S. sectoral prices and quantities, both during recessions and conditional on an identified monetary shock. Section 3 describes the theoretical model. Section 4 introduces the calibration strategy, presents model-based impulse response functions to a monetary shock and compares them with VAR-based im-
pulse responses, and scrutinizes the quantitative role of nominal rigidities and variable markups. Section 5 concludes.

2 Empirical evidence: sectoral prices and quantities

This section uses monthly disaggregated data on U.S. manufacturing producer prices and industrial production to establish a number of stylized facts. I group industries into sectors producing nondifferentiated and differentiated goods. Special attention is paid to the patterns exhibited by sectoral prices and production during recessions. I also survey related evidence from international trade data. Finally, I estimate vector autoregressions (VARs) to analyze the impact of monetary policy shocks on relative prices and quantities between those two sectors.\footnote{The analysis focuses on prices that reflect business-to-business transactions, namely producer prices (wholesale prices) and prices at the dock of internationally traded goods (border prices).}

2.1 Domestic data: new facts

To compute measures of domestic production and prices of nondifferentiated and differentiated manufacturers, I first map the 2007 version of the Rauch (1999) categories, originally available at the 4-digit Standard International Trade Classification level, into the NAICS classification of manufacturing industries. Details on the concordance method are provided in Appendix A. In the Rauch (1999) classification, I treat goods traded on organized exchanges or with reference prices as nondifferentiated. For illustrative purposes, Table 1 reports an exhaustive list of 4-digit NAICS industries categorized as differentiated, nondifferentiated, or unclassified. Examples of nondifferentiated goods industries are Dairy Products; Beverages; Fibers, Yarns, and Threads; Basic Chemicals; and Nonferrous Metal (Except Aluminum) and Processing. Differentiated goods industries include Apparel; Pharmaceuticals and Medicines; Industrial Machinery; and Audio and Video
Equipment. In the analysis below, I exclude petroleum industries.

[Table 1]

Aggregate production of nondifferentiated and differentiated goods is measured as the weighted average of (seasonally-adjusted) industrial production indices corresponding to 4-digit NAICS manufacturing industries classified as nondifferentiated and differentiated, respectively. The weights are calculated using data on the relative importance weight of each individual industry in the overall Industrial Production index. The industry-level data are available since January 1972 and their source is the Federal Reserve Board of Governors.

The aggregate producer prices for nondifferentiated and differentiated goods are constructed as weighted averages across the producer price indices (PPIs) for industries classified as nondifferentiated or differentiated, respectively. Consistent with the methodology underlying the construction of the overall PPI index, the weights are based on industry shipment values taken from the 2007 Economic Census. The price data are available since January 1976 and are obtained from the Bureau of Labor Statistics (BLS). For the PPI analysis, I use data at the 6-digit NAICS level, as this allows me to maximize the sample period while collecting many disaggregated price series with a balanced number of observations. All indices are normalized to 100 in 2007.

Figure 1 depicts the percentage (log) changes of the sectoral industrial production indices for the period 1972:1 to 2014:6, along with NBER recession bars. The sectoral indices are smoothed using 9-month moving averages to remove high-frequency noise. The evidence suggests that while production in both sectors declines during recessions, the quantity produced of differentiated goods falls relatively more than that of nondifferentiated goods. The first two rows of Table 2 show the total declines in sectoral production during the six recessions in the sample period and during the acute phase of the last recession. In an average economic downturn, the decline in production of differentiated
goods is roughly twice as large as for nondifferentiated goods. For instance, the sectoral cumulated changes in production during the 2007–2009 recession were −27 percent (differentiated goods) and −13 percent (nondifferentiated goods).

[Figure 1]
[Table 2]

Figure 2 shows the percentage changes of the (smoothed) sectoral PPIs for the period 1976:1–2014:6. In the graph, the relative stability of differentiated goods prices contrasts with the much more volatile prices of nondifferentiated goods. This is consistent with evidence from micro data underlying the construction of the U.S. PPI, that raw manufactured goods tend to show higher price flexibility than other categories of more differentiated manufacturers (Nakamura and Steinsson, 2008). As observed in the last two rows of Table 2, prices of nondifferentiated goods declined in four of the last five recessions. In the critical phase of the 2007–2009 recession (2008:8–2009:6), their fall was particularly severe (−15 percent). On the other hand, differentiated goods prices either increased or remained fairly stable in all of the recessions included in the sample period.8

[Figure 2]

The empirical results imply an important role for differences in the elasticity of sectoral supply curves. To clarify this idea, consider the following argument. The majority of recent recessions are associated with demand shocks, which tend to reduce prices and quantities for all sectors. However, differentiated goods are likely to face larger shifts in demand than nondifferentiated goods. This is because differentiated goods are arguably

8I also found that the heterogeneity in the adjustment of sectoral PPIs holds controlling for durability of the goods. But in terms of production, at least in 2007–2009, the disproportionate decline in the quantity produced of differentiated goods was mostly driven by the adjustment in nondurable goods. Moreover, although my focus in on recessions, I have found evidence of symmetric sectoral quantity and price patterns during the following booms: 1978q2–1979q1, 1987q4-1989q1, and 1998q4-2000q2. In all these episodes, nondifferentiated goods prices increased relative to those of differentiated goods, and the quantity produced of differentiated goods increased at least as much as that of nondifferentiated goods.
characterized by higher income elasticity and durability than nondifferentiated goods. Although by itself a bigger demand shock could explain the relatively larger quantity drops observed among differentiated goods, it cannot rationalize the evidence of smaller price declines among these industries. In fact, a flatter supply curve for differentiated goods than for nondifferentiated goods appears to be a key element to explain the empirical findings.

2.2 International trade data: a brief survey

In their study of micro data on U.S. import and export prices at the dock, Gopinath and Rigobon (2008) document that differentiated goods display lower frequency of price adjustment than nondifferentiated goods. There is also evidence that differentiated goods prices of imports and exports were stable during the 2007–2009 recession, whereas nondifferentiated goods prices strongly declined (see e.g. Gopinath et al., 2012). Moreover, in that recession the reduction in trade volumes of differentiated goods was particularly severe, whereas nondifferentiated goods exhibited relatively smaller quantity decreases (see e.g. Haddad et al., 2010). That is, the aforementioned patterns for sectoral domestic prices and quantities appear to be present as well in international trade data.

It is worth noting that the heterogeneity across nondifferentiated and differentiated goods in international trade variables has been documented by papers that employ different sources of data. Gopinath et al. (2012) use confidential product-level data on at-the-dock prices from the BLS, Levchenko et al. (2010) employ data on trade flows and prices for end-use industries, and Haddad et al. (2010) use data on trade values and unit values at the 6-digit Harmonized System level for a group of countries including the U.S.\footnote{Gopinath et al. (2012) explicitly identify manufactured goods in their data. They also verify the robustness of the results to the consideration of different relationship structures between trading parties (i.e., market-based or related-party transactions), different durability and end-uses of the goods, and different locations of the trading partners.}
2.3 Effects of monetary policy shocks: VAR evidence

So far I have presented evidence of heterogeneous adjustment in prices and quantities across nondifferentiated and differentiated manufacturers during recessions. The previous results imply that in a typical recession both the relative prices of nondifferentiated goods to differentiated goods and the relative quantities of differentiated goods to nondifferentiated goods decline.

I now investigate whether these patterns of adjustment in relative prices and quantities hold conditional on a contractionary monetary policy shock. The results of this exercise provide an empirical benchmark to test the sectoral predictions of the model described in the next section. As anticipated in the introduction, standard monetary models with heterogeneity in price stickiness would predict declines in both the ratio of nondifferentiated-to-differentiated goods prices and the ratio of differentiated-to-nondifferentiated goods quantities in the wake of a tightening monetary shock.

I estimate VARs on monthly U.S. data to generate impulse response functions (IRFs) to contractionary monetary policy shocks. The identification method is based on the standard recursiveness assumption developed in Christiano et al. (1999)—the central bank observes the current and lagged values of all the variables in the VAR, and the monetary policy shock affects only the monetary instrument contemporaneously. I report IRFs to a positive one-standard-deviation innovation to the federal funds rate. The approximate 90% confidence bands associated with the IRFs are constructed using 500 Monte Carlo replications.

Domestic data: The VAR for domestic data includes (in Cholesky order) the relative industrial outputs of differentiated to nondifferentiated goods, the relative PPIs of nondifferentiated to differentiated goods, the Commodity Research Bureau commodity price index, and the federal funds rate. Both the output and the price ratios enter the VAR in logs, while the commodity price index is in log changes, and the federal funds rate is in
levels. The sample period is 1976:1–2012:12 and the lag order is 9 (which is sufficient to eliminate autocorrelation of the residuals).\footnote{The commodity price index is introduced to capture supply shocks and to alleviate the problem of “price puzzles” in the responses of prices. To assess the sensitivity of the results, I estimated other VARs using different specifications; e.g., introducing money (M2), including more lags, and using the Wu and Xia (2014) “shadow federal funds rate” that allows the Fed’s policy rate to go below zero during the recent “zero lower bound” period. The main results in these VARs are broadly the same as in the baseline estimation.}

As shown in Figure 3, prices of nondifferentiated goods exhibit a statistically significant decline relative to prices of differentiated goods within the first year after the shock (left panel). The peak response is nearly $-0.3$ percent and occurs after three to five months. The relative quantity produced of differentiated to nondifferentiated goods also falls following a monetary contraction in a statistically significant way (right panel), and the effect appears to be highly persistent. During the first six first months after the shock, the maximum reduction in the relative outputs is 0.16 percent.\footnote{I also estimated two additional sectoral VARs, each including sector-specific measures of PPI and industrial production, and find no evidence of price puzzles, as desired.}

\textbf{Figure 3}

My results in terms of price responses echo those reported by Balke and Wynne (2007). They study highly disaggregated U.S. producer prices, and find suggestive evidence that a contractionary monetary shock is more likely to have the standard short-term negative effect on prices of raw goods than on prices of more processed goods.

\textbf{Trade data:} To evaluate the responses of relative prices and quantities to monetary shocks in international trade data, I use disaggregated data on U.S. exports and imports from the U.S. International Trade Commission (USITC) website, readily available at the 4-digit NAICS level since 1997. In particular, export and import volume indices for nondifferentiated and differentiated goods are constructed as weighted averages across 4-digit NAICS manufacturing industries classified as nondifferentiated and differentiated, respectively.\footnote{Outliers are excluded by eliminating changes of a magnitude greater than 2 log points in the disaggregated trade volume series.} The weights are based on total trade values over the sample period. A
similar procedure is followed to construct sectoral measures of export and import prices at the dock. To proxy for these prices I use data on unit values published by the USITC. In the appendix I provide further methodological details on the construction of these variables.

In an initial analysis of the USITC-based sectoral measures of trade prices and volumes, I find that the import data conform with the evidence surveyed in section 2.2. In particular, during the 2001 and 2007–2009 recessions, both the relative import prices of nondifferentiated goods to differentiated goods and the relative import volumes of differentiated goods to nondifferentiated goods decline. However, the relative export prices of nondifferentiated goods to differentiated goods slightly increase in the 2007–2009 recession. This anomaly might indicate problems associated with the use of unit values as proxies for export prices or with the use of data at the industry-level (e.g., the impossibility of isolating price changes from shifts in quality, or composition changes within each industry). In light of this finding, I conduct the VAR analysis only for data on imports.

To conserve degrees of freedom, I estimate two separate VARs: one for relative import prices, and the other for relative import volumes. For the same reason, the VARs exclude two potentially important variables in an open economy, namely the foreign interest rate and the exchange rate. However, adding either of these variables or estimating a single VAR with the relative prices and the relative quantities does not change the main results. The variables in Cholesky order are domestic manufacturing production, a domestic price index, foreign (average of non-U.S., G-6 countries) industrial production, a commodity price index, the relative import volumes of differentiated to nondifferentiated goods or the relative import prices of nondifferentiated to differentiated goods, and the federal funds rate. The relative prices and volumes enter in logs, and the federal funds rate is in levels. All other variables enter in log changes. The sample period is 1997:1–2012:12 and the lag

\footnote{See Eichenbaum and Evans (1995) and Kim (2001) for related VAR specifications of open-economy monetary models in the spirit of Christiano et al. (1999).}
order is 9.\textsuperscript{14}

In Figure 4, the left panel shows that the import prices of nondifferentiated goods relative to prices of differentiated goods tend to decline following the monetary policy shock. The maximum statistically significant response of the relative price is a decline of 1 percent after six months. The response of the import volumes of differentiated goods relative to that of nondifferentiated goods (right panel) is imprecisely estimated, as implied by the wide confidence bounds. However, around the horizon when the relative import price is most responsive to the monetary policy shock (i.e., between the fourth and the sixth month after the shock), the relative import quantities decline too, showing a peak reduction of 0.4 percent.

[Figure 4]

To summarize, the VAR evidence for domestic and trade data suggests that the movements in the relative prices and relative quantities conditional on a contractionary monetary policy shock are consistent with the patterns of adjustment across sectors observed during a typical recession. That is, the relative prices of nondifferentiated goods to differentiated goods and the relative quantities of differentiated goods to nondifferentiated goods tend to decline in the short run following a monetary contraction.

3 Model

This section presents a two-country multi-sector general equilibrium model with heterogeneity in price stickiness. The model features two tradable manufacturing sectors, which produce nondifferentiated ($N$) and differentiated ($D$) goods. The description of the model

\textsuperscript{14}The source for the foreign variables is the International Financial Statistics dataset. The results are robust to several model modifications, such as introducing money (M2) and changing the order of the variables.
mostly focuses on the Home country (H). Unless otherwise stated, analogous equations hold for the Foreign country (F).

In each country, infinitely-lived households consume a CES aggregate over tradable and nontradable goods, supply labor to producing firms, and invest in a complete set of freely-traded state-contingent financial assets, which without loss of generality are denominated in the Home currency. An explicit reference to the states of nature is omitted to simplify the notation. Labor is immobile across countries and mobile across sectors. To close the model, nominal aggregate spending is assumed to follow an exogenous autoregressive process.

**Households.** The representative household in the Home country maximizes expected lifetime utility:

$$E_0 \sum_{t=0}^{\infty} \beta^t \left[ \frac{C_{H,t}^{1-\gamma}}{1-\gamma} - \frac{L_{H,t}^{1+\phi}}{1+\phi} \right],$$

where $C_{H,t}$ is an overall consumption index, $L_{H,t}$ denotes labor supply, $E_0$ is the time-0 expectations operator, $\beta$ is the discount factor, $\gamma$ is the inverse of the intertemporal elasticity of substitution, and $\phi$ is the inverse of the Frisch elasticity of labor supply.

The consumption aggregators have a CES structure. Final consumption $\overline{C}_{H,t}$ is an aggregator of tradable and nontradable goods, denoted by $C_{H,t}$ and $C_{\pi H,t}$, respectively:

$$\overline{C}_{H,t} = \left[ \chi^{1/\omega} \left( C_{H,t} \right)^{(\omega-1)/\omega} + (1-\chi)^{1/\omega} \left( C_{\pi H,t} \right)^{(\omega-1)/\omega} \right]^{\omega/(\omega-1)},$$

where $\omega \geq 0$ is the elasticity of substitution between tradable and nontradable goods, and $\chi \in [0,1]$ is the steady-state share of tradable goods in final consumption. The consumption of nontradable goods $C_{\pi H,t}$ is given by an aggregator of varieties $i \in [0,1]$ with elasticity of substitution $\xi > 1$: $C_{\pi H,t} = \left( \int_0^1 C_{H,t} \left( i \right)^{(\xi-1)/\xi} di \right)^{\xi/(\xi-1)}$. The consumption of tradable goods $C_{H,t}$ is an aggregator of nondifferentiated and differentiated goods.
(denoted by $C_{HN,t}$ and $C_{HD,t}$, respectively) with elasticity of substitution $\eta \geq 0$:

$$C_{H,t} = \left[ \mu_N^{1/\eta} C_{HN,t}^{(\eta-1)/\eta} + \mu_D^{1/\eta} C_{HD,t}^{(\eta-1)/\eta} \right]^{\eta/(\eta-1)},$$

where $\mu_N = 1 - \mu_D \in [0, 1]$ controls the expenditure share of nondifferentiated goods. The tradable good from sector $s$ consumed in Home is given by a composite of varieties $i \in [0, 1]$ produced by firms in that sector both in the Home and Foreign countries, as follows:

$$C_{Hs,t} = \left[ \psi_s^{1/\sigma_s} C_{HHs,t}^{(\sigma_s-1)/\sigma_s} + (1 - \psi_s)^{1/\sigma_s} C_{FHs,t}^{(\sigma_s-1)/\sigma_s} \right]^{\sigma_s/(\sigma_s-1)},$$

where $s \in \{N,D\}$, $k \in \{H,F\}$, and the parameter $\psi_s \in [1/2, 1]$ introduces home bias. The sector-specific elasticity of substitution $\sigma_s > 1$ is assumed to be the same across the domestic and imported sectoral composites ($C_{HHs,t}$ and $C_{FHs,t}$), and across the varieties produced in a given country $k$ ($C_{kHs,t}(i)$). A reasonable assumption is that $\sigma_N > \sigma_D$. That is, it is easier to substitute nondifferentiated goods than differentiated goods both across varieties $i$ and across countries of supply $k$.\footnote{As I discuss below in the calibration of the model, there is empirical evidence in line with the assumption $\sigma_N > \sigma_D$. Using the same elasticity of substitution at the sector level for two levels of disaggregation of consumption (i.e., across countries of origin and across domestic varieties) simplifies the calibration.}

Consumer optimization yields the following set of demands:

$$C_{H,t} = \chi \left( \frac{P_{H,t}}{\bar{P}_{H,t}} \right)^{-\omega} \bar{C}_{H,t}, \quad C_{H,t}^2 = (1 - \chi) \left( \frac{P_{H,t}^2}{\bar{P}_{H,t}} \right)^{-\omega} \bar{C}_{H,t},$$

$$C_{Hs,t} = \mu_s \left( \frac{P_{Hs,t}}{\bar{P}_{H,t}} \right)^{-\eta} \bar{C}_{H,t}, \quad C_{HHs,t} = \psi_s \left( \frac{P_{HHs,t}}{\bar{P}_{Hs,t}} \right)^{-\sigma_s} \bar{C}_{Hs,t},$$

$$C_{Hs,t} = \mu_s \left( \frac{P_{Hs,t}}{\bar{P}_{H,t}} \right)^{-\eta} \bar{C}_{H,t}, \quad C_{HHs,t} = \psi_s \left( \frac{P_{HHs,t}}{\bar{P}_{Hs,t}} \right)^{-\sigma_s} \bar{C}_{Hs,t},$$
\[ C_{FHs,t} = (1 - \psi_s) \left( \frac{P_{FHs,t}}{P_{Hs,t}} \right)^{-\sigma_s} C_{Hs,t}, \quad C_{kHs,t}(i) = \left( \frac{P_{kHs,t}(i)}{P_{kHs,t}} \right)^{-\sigma_s} C_{kHs,t}, \]

where \( \mathcal{P}_{H,t} \) is the overall consumption price index, \( P_{H,t} \) is the price index of tradable goods, \( P_{H,t}^z \) is the price index of nontradable goods, \( P_{Hs,t} \) is the price index of tradable goods from sector \( s \), \( P_{kHs,t}(i) \) is the price index of tradable goods from sector \( s \) sold in the Home country and produced in country \( k \), \( P_{kHs,t}(i) \) is the price of variety \( i \) from tradable sector \( s \) sold in the Home country and produced in country \( k \), and \( P_{zH,t}(i) \) is the price of variety \( i \) of nontradable goods. The CES price indices are given by:

\[ \mathcal{P}_{H,t} = \left[ \chi \left( P_{H,t} \right)^{1-\omega} + (1 - \chi) \left( P_{zH,t} \right)^{1-\omega} \right]^{1/(1-\omega)}, \quad P_{H,t}^z = \left( \int_0^1 P_{H,t}(i)^{1-\xi} di \right)^{1/(1-\xi)}, \]

\[ P_{H,t} = \left[ \mu_N P_{HN,t}^{1-\eta} + \mu_D P_{HD,t}^{1-\eta} \right]^{1/(1-\eta)}, \quad P_{Hs,t} = \left[ \psi_s P_{HHs,t}^{1-\sigma_s} + (1 - \psi_s) P_{FHs,t}^{1-\sigma_s} \right]^{1/(1-\sigma_s)}, \]

\[ P_{kHs,t} = \left( \int_0^1 P_{kHs,t}(i)^{1-\sigma_s} di \right)^{1/(1-\sigma_s)}. \]

The household’s flow budget constraint is:

\[ \mathcal{P}_{H,t} C_{H,t} + \mathbb{E}_t \Theta_{t,t+1} B_{H,t+1} \leq W_{H,t} L_{H,t} + B_{H,t} + \Pi_{H,t}, \]

where \( B_{H,t+1} \) is the state-contingent value of the portfolio held at the beginning of period \( t+1 \) (optimally chosen for each possible state of nature), \( \Pi_{H,t} \) is the sum of profits from Home tradable and nontradable firms, and \( W_{H,t} \) is the nominal wage. The nominal stochastic discount factor \( \Theta_{t,t+1} \) that prices the financial asset portfolio in period \( t \) is the same for both countries given an assumption of no arbitrage opportunities.

Utility maximization subject to the flow budget constraint and to a standard solvency constraint yields a consumption-labor supply optimality condition:

\[ \bar{C}_{H,t}^\psi L_{H,t}^\phi = \frac{W_{H,t}}{\mathcal{P}_{H,t}}, \]
and an international risk-sharing condition:

\[
\frac{E_{HF,t} P_{F,t}}{P_{H,t}} = \left( \frac{C_{H,t}}{C_{F,t}} \right)^\gamma,
\]

where \(E_{HF,t}\) is the bilateral nominal exchange rate (price of a unit of Foreign currency in terms of units of Home currency), and \(C_{F,t}\) and \(P_{F,t}\) are the Foreign-country counterparts of the Home overall consumption and price indices, respectively. The risk-sharing condition results from combining the intertemporal first-order conditions for asset holdings in both countries, assuming symmetric initial conditions.

**Distribution and retail sectors.**

Firms in the distribution sector are perfectly competitive. They combine varieties of distribution services \(D_{H,t}^z(i)\) supplied by firms in the nontradable sector to produce a CES composite defined by

\[
D_{H,t}^z = \left( \int_0^1 D_{H,t}^z(i)(\xi - 1)/\xi di \right)^{\xi/(\xi - 1)}.
\]

For simplicity, no distinction is made between nontradable consumption goods and distribution services. Hence, the price of the composite \(D_{H,t}^z\) is given by the price index of nontradable goods \(P_{H,t}^z\), and the optimal demand from the distribution sector for a variety of nontradables is given by

\[
D_{H,t}^z(i) = \left( \frac{P_{H,t}^z(i)}{P_{H,t}} \right)^{-\xi} D_{H,t}^z.
\]

Firms in the retail sector are also perfectly competitive and their prices are flexible. They combine varieties either of nondifferentiated or differentiated goods with distribution services, in fixed proportions, before selling them to final consumers. Thus, the retail price in the Home country for a sector-\(s\) variety \(i\) produced domestically is given by:

\[
P_{HHs,t}(i) = P_{HHs,t}^P(i) + \kappa_s P_{H,t}^z,
\]

where \(P_{HHs,t}(i)\) denotes the corresponding producer price in units of the Home currency, \(\kappa_s > 0\) denotes the required number of units of distribution services to bring sector-\(s\) goods to Home consumers (or alternatively, the fixed distribution cost per sector-\(s\) good), and

\[16^\text{The specification of these sectors is based on the original contribution of Corsetti and Dedola (2005). Burstein and Gopinath (2014) discuss several models in the international macro literature that produce endogenous variable markups.}\]
$P_{H,t}^z$ is the unit price of such distribution services.\(^{17}\)

I assume an asymmetric pricing structure for cross-border transactions. Producers in the Home country set prices for export in units of the Home currency, while producers in the Foreign country set prices for export also in units of the Home currency. Therefore, the retail price in the Foreign country for a sector-$s$ variety $i$ produced in the Home country is:

$$P_{HFs,t}(i) = \frac{P_{HFs,t}(i)}{E_{HF,t}} + \kappa_s P_{F,t}^z,$$

where $P_{HFs,t}(i)$ denotes the producer price set by a Home firm $i$ in units of the Home currency, and $P_{F,t}^z$ is the unit price of distribution services in the Foreign country.\(^{18}\) The retail price in the Home country for a sector-$s$ variety $i$ produced in the Foreign country is:

$$P_{FHs,t}(i) = P_{FHs,t}(i) + \kappa_s P_{H,t}^z,$$

where $P_{FHs,t}(i)$ denotes the producer price set by a Foreign firm $i$ in units of the Home currency.

** Tradable sectors.** Firms in the tradable sectors produce nondifferentiated and differentiated goods using linear technologies with labor as the only input: $Y_{Hs,t}(i) = A_{Hs,t}L_{Hs,t}(i)$, where $A_{Hs,t}$ denotes productivity in sector $s$. (In the numerical exercises below, productivity shocks are ignored, that is, $A_{ks,t}$ is set to 1.)

Prices in the tradable sectors are sticky as a result of two mechanisms. First, I assume nominal price rigidities of the Calvo (1983) type. Second, the retail technology described above gives rise to endogenous variable markups at the producer level.

As mentioned above, I adopt a hybrid specification of price setting, by which both Foreign and Home firms set their export prices in the Home currency. That is, Foreign

\(^{17}\)Note that an elasticity of substitution between tradable goods and distribution services below one is assumed in the constant returns to scale retail technology. Also, the sectoral fixed proportion of distribution costs in the retail price, $\kappa_s$, is assumed to be constant across countries.

\(^{18}\)As shown below, due to the local distribution costs, $P_{HHs,t}(i)$ is different from $P_{HFs,t}(i)$. In other words, a given producer discriminates between markets.
firms follow “local currency pricing” (LCP) and Home firms engage in “producer currency pricing” (PCP) when selling abroad (see e.g. Devereux et al., 2007). Following Corsetti and Pesenti (2009), I label this specification as “dollar pricing”. This modelling choice is consistent with our analysis of the model as a two-country world economy with the U.S. and the rest of the world, as well as with evidence that for the U.S. there is PCP in exports and LCP in imports (Gopinath and Rigobon, 2008).

To analyze the variable markup channel, I first ignore Calvo pricing and assume that producer prices are flexible. The problem of a sector- \( s \) firm \( i \) from country \( H \) selling to country \( k \) is:

\[
\max_{P_{Hks,t}^P(i)} \left( P_{Hks,t}^P(i) - \frac{W_{H,t}}{A_{Hs,t}} \right) \left( \frac{P_{Hks,t}^P(i)}{P_{Hks,t}^P} \right)^{-\sigma_s} C_{Hks,t}.
\]

Note that this maximization problem is subject to the relationship between producer and consumer prices shown in equations (1) and (2). By symmetry, the optimal sectoral producer price under flexible prices, denoted by \( \tilde{P}_{Hks,t}^P \), is common to all firms \( i \) in a given sector \( s \), and can be written as an optimal markup over the marginal cost:

\[
\tilde{P}_{Hks,t}^P = \begin{cases} 
\zeta_{HHs,t} \frac{W_{H,t}}{A_{Hs,t}}, & \text{if } k = H \\
\zeta_{HF_{s,t}} \frac{W_{H,t}}{A_{Hs,t}}, & \text{if } k = F
\end{cases}
\]

where \( \zeta_{HHs,t} \) and \( \zeta_{HF_{s,t}} \) are the sector- and destination country-specific optimal markups, given by:

\[
\zeta_{HHs,t} = \frac{\sigma_s}{\sigma_s - 1} \left[ 1 + \frac{1}{\sigma_s} \frac{A_{Hs,t}}{W_{H,t}} P_z^z \right], \quad \zeta_{HF_{s,t}} = \frac{\sigma_s}{\sigma_s - 1} \left[ 1 + \frac{1}{\sigma_s} \frac{A_{Hs,t}}{W_{H,t}} E_{HF,t} P_z^z \right].
\]

The familiar result that a higher elasticity of substitution \( \sigma_s \) reduces the sectoral markup holds. More interestingly, the optimal markups are decreasing in the marginal cost of the origin country and increasing in the distribution cost of the destination country.\(^{19}\) The

\(^{19}\)Analogous maximization problems for the Foreign firm yield the following expressions for the optimal producer prices in units of the producer’s currency: \( \tilde{P}_{FFs,t}^P = \zeta_{FFs,t} \frac{W_{F,t}}{A_{Fs,t}} \) and \( \tilde{P}_{FHs,t}^P = \zeta_{FHs,t} \frac{W_{F,t}}{E_{HF,t}} \), where the optimal markups \( \zeta_{FFs,t} \) and \( \zeta_{FHs,t} \) are given by:

\[
\zeta_{FFs,t} = \frac{\sigma_s}{\sigma_s - 1} \left[ 1 + \frac{1}{\sigma_s} \frac{A_{Fs,t}}{W_{F,t}} P_z^z \right], \quad \zeta_{FHs,t} = \frac{\sigma_s}{\sigma_s - 1} \left[ 1 + \frac{1}{\sigma_s} \frac{A_{Fs,t}}{W_{F,t}} E_{HF,t} P_z^z \right].
\]
intuition behind these results is clarified by analyzing the producer’s price elasticity of demand, which in the case of sales abroad, for example, is given by:

$$\epsilon_{HF}^{s}(i) \equiv -\frac{\partial \log C_{HF}^{s,t}(i)}{\partial \log \left(\frac{P_{HF}^{s,t}(i)}{E_{HF,t}}\right)} = \sigma_{s} \left(1 - x_{HF}^{s,t}(i)\right),$$

where $$x_{HF}^{s,t}(i) \equiv \kappa_{s} P_{z}^{s,t} / \left(1 + \kappa_{s} P_{z}^{s,t} / E_{HF,t}\right)$$ is the sectoral share of the distribution cost in the retail price $$P_{HF}^{s,t}(i)$$—or alternatively, the sectoral “distribution margin” for final consumption of exported goods sold in the Foreign country. A higher producer price relative to the local distribution cost (measured in the local currency) reduces the distribution margin, which in turn increases the price elasticity of demand. Since the monopolistic competitor’s optimal markup is negatively related to the elasticity of demand (i.e., $$\zeta_{HF}^{s,t} = \frac{\epsilon_{HF}^{s}(i)}{\epsilon_{HF}^{s}(i) - 1}$$), a negative relationship between the markup and the producer’s price relative to the distribution cost is therefore likely to emerge. The sector-specific absolute elasticity of the markup $$\zeta_{HF}^{s,t}$$ with respect to the relative price $$\frac{P_{HF}^{s,t}(i)}{E_{HF,t} P_{z}^{s,t}}$$ is:

$$\Gamma_{HF}^{s,t}(i) \equiv -\frac{\partial \log \zeta_{HF}^{s,t}}{\partial \log \left(\frac{P_{HF}^{s,t}(i)}{E_{HF,t} P_{z}^{s,t}}\right)} = \left[\left(\sigma_{s} - 1\right) - \frac{x_{HF}^{s,t}(i)}{x_{HF}^{s,t}(i) - 1}\right]^{-1}.$$  \(3\)

The elasticity of the markup $$\Gamma_{HF}^{s,t}(i)$$ is greater (in absolute value) in a sector $$s$$ characterized by relatively high distribution margin $$x_{HF}^{s,t}(i)$$ and low elasticity of substitution $$\sigma_{s}$$. More specifically, equation (3) implies that $$\Gamma_{HF}^{s,t}(i) = 0$$ if $$x_{HF}^{s,t}(i) = 0$$. Below I show that for realistic values of $$x_{HF}^{s,t}(i)$$ and $$\sigma_{s}$$, the case in which $$x_{HF}^{s,t}(i) > 0$$ implies $$\Gamma_{HF}^{s,t}(i) > 0$$. Since the optimal producer prices are the same for all firms $$i$$ in a given sector $$s$$, I hereafter drop the firm index $$i$$ to refer to the (equilibrium) sector-specific distribution margin ($$x_{HF}^{s,t}$$), the price elasticity of demand ($$\epsilon_{HF}^{s}$$), and the elasticity of the markup ($$\Gamma_{HF}^{s,t}$$).\(^{20}\)

$$\zeta_{FH}^{s,t} \equiv \frac{\sigma_{s} - 1}{\sigma_{s}} \left[1 + \frac{A_{z}^{s,t}}{\sigma_{s} W_{z}^{s,t} P_{HF}^{s,t}} + \frac{P_{HF}^{s,t}}{E_{HF,t}}\right].$$

\(^{20}\)The producer’s elasticity of the markup for domestic sales is given by $$\Gamma_{HH}^{s,t} = \ldots$$
I now reintroduce the assumption that nominal prices are rigid in the currency of the producer/exporter according to Calvo pricing. That is, in each period a firm from sector $s$ adjusts prices with constant probability $1 - \theta_s$. For a zero-inflation steady state, the log-linear optimal sectoral reset price for sales of a Home firm in country $k$ is given by:

$$\tilde{p}^P_{Hks,t} = (1 - \beta \theta_s) \sum_{j=0}^{\infty} (\beta \theta_s)^j E_t \tilde{p}^P_{Hks,t+j},$$

(4)

where lower-case letters denote log-deviations from the steady state, so $\tilde{p}^P_{Hks,t}$ denotes the optimal log-linear price for sales of a Home firm in country $k$ if prices were flexible. Our previous results under flexible prices imply the following pricing equation for $\tilde{p}^P_{Hks,t}$:

$$\tilde{p}^P_{Hks,t+j} = \begin{cases} 
\frac{1}{1+\Gamma_{Hks}} \left( w_{H,t+j} - a_{H,s,t+j} \right) + \frac{\Gamma_{Hks}}{1+\Gamma_{Hks}} \tilde{p}_{H,t+j}, & \text{if } k = H \\
\frac{1}{1+\Gamma_{Fks}} \left( w_{H,t+j} - a_{H,s,t+j} \right) + \frac{\Gamma_{Fks}}{1+\Gamma_{Fks}} \left( \tilde{p}_{F,t+j} + e_{HF,t+j} \right), & \text{if } k = F \end{cases},$$

(5)

For either $k = H$ or $k = F$, the first terms in (5) indicate that with variable markups ($\Gamma_{Hks} > 0$), movements in the marginal cost are incompletely passed through into $\tilde{p}^P_{Hks,t}$. The pass-through rate is given by $\frac{1}{1+\Gamma_{Hks}}$, which declines with the steady-state sectoral elasticity of the markup $\Gamma_{Hks}$. Equations (4) and (5) thus imply that, in the face of changes in marginal costs, variable markups extend non-adjustments in producer prices beyond the period implied by the nominal rigidity mechanism, and increasingly so for a higher markup elasticity.

The second terms in (5) reflect the dependence of optimal markups on local distribution costs, which are determined by the price of nontradables in the destination country $\tilde{p}_{F,t}$. For the price $\tilde{p}^P_{HFs,t}$ set for sales in country $F$ (denominated in the Home currency), distribution costs are also affected by the nominal exchange rate $e_{HF,t}$. An appreciation of the Home currency (i.e., a decline in $e_{HF,t}$) reduces distribution costs, which lowers the optimal markup and hence the desired price. Movements in local distribution costs are

$$\left( \sigma_s - 1 \right)^{-1} \left( x_{H,Hs,t+1} - 1 \right)^{-1},$$

where $x_{H,Hs,t} \equiv \kappa_s \frac{\tilde{p}_{H,t}^P}{\tilde{p}_{H,Hs,t}^{(i)} + \kappa_s \tilde{p}_{H,t}^P}$ is the sectoral distribution margin for final consumption of domestically produced goods sold in the Home country.
passed through into the optimal producer prices with a coefficient \( \frac{\Gamma_{Hks}}{1+\Gamma_{Hks}} \), which increases with the sensitivity of the markup \( \Gamma_{Hks} \).

Analogous expressions can be derived for the Foreign firm. In this case, however, the optimal price under flexible prices set for sales in country \( H \) (\( \tilde{p}_{FHs,t+j}^P \)) is denominated in units of the Home currency due to the LCP assumption for Home imports. This price is given by:

\[
\tilde{p}_{FHs,t+j}^P = \frac{1}{1 + \Gamma_{FHs}} \left( w_{F,t+j} + e_{HF,t+j} - a_{Fs,t+j} \right) + \frac{\Gamma_{FHs}}{1 + \Gamma_{FHs}} \tilde{p}_{H,t+j}^z,
\]

where the presence of the exchange rate \( e_{HF,t+j} \) in the first term of equation (6) implies that an appreciation of the Home currency causes a reduction in the marginal cost expressed in units of the Home currency, and hence in the optimal price. As implied by the coefficient \( \frac{1}{1+\Gamma_{FHs}} \), this exchange rate pass-through effect into \( \tilde{p}_{FHs,t+j}^P \) is incomplete, and the degree of incompleteness is larger for a sector with higher markup elasticity \( \Gamma_{FHs} \).

Meanwhile, changes in the distribution cost of the destination country, given by \( p_{z}^{H,t+j} \), are passed through with a coefficient \( \frac{\Gamma_{FHs}}{1+\Gamma_{FHs}} \).

Overall, the variable markup channel implies that a higher elasticity of the markup reduces the response of prices to movements in the marginal cost, but increases their response to changes in the distribution costs.

**Nontradable sector.** Firms in the nontradable sector produce using linear technologies in labor, \( Y_{H,t}(i) = L_{H,t}(i) \), and face Calvo price rigidities with probability of non-adjustment denoted by \( \theta^z \). Markups in this sector are constant since distribution services are assumed to be unnecessary for nontradable goods (see Goldberg and Campa, 2010 for supporting evidence). The log-linear price-setting condition for all firms \( i \) in the nontradable sector is given by:

\[
\tilde{p}_{H,t}^z = (1 - \beta \theta^z) \sum_{j=0}^{\infty} (\beta \theta^z)^j E_t \tilde{p}_{H,t+j}^z,
\]

where the optimal price in the absence of nominal rigidity is \( \tilde{p}_{H,t+j}^z = w_{H,t+j} \).
**Exogenous nominal spending.** To close the model I assume that nominal spending (or equivalently, nominal aggregate consumption) is driven by an exogenous money supply rule: \( \mathcal{P}_{H,t} \mathcal{C}_{H,t} = M_{H,t} \), where \( \log M_{H,t} \) follows an autoregressive process with shock \( \varepsilon_{MH,t} \sim N(0,\sigma_{MH}) \). Other papers make a similar assumption (see e.g. Chari et al., 2000).

**Equilibrium.** An equilibrium is an allocation and set of prices for all states of nature and periods, such that households and firms optimize, and assets, goods, and labor markets clear in both the Home and the Foreign country. The market clearing conditions are relegated to Appendix B. The equilibrium is symmetric within each country, but asymmetric across countries due to the asymmetric pricing structure.

**Additional definitions.** I now define price indices at the industry level for the Home country that will be analyzed in our quantitative exercise. The sectoral export price index, \( EPI_{Hs} \), captures the prices for sector-\( s \) goods exported by the Home country, measured in the Home currency. The individual firms’ export prices at the dock measured in the producer’s currency are equivalent to \( P_{HFs,t}^P(i) \). Under the assumptions of symmetric firms and time-dependent pricing, \( EPI_{Hs,t} \) is then given by:

\[
EPI_{Hs,t} = \left( \int_0^1 P_{HFs,t}^P(i)^{1-\sigma_s} di \right)^{1/(1-\sigma_s)} = \left[ \theta_s \left( EPI_{Hs,t-1} \right)^{1-\sigma_s} + (1-\theta_s) \left( \hat{P}_{HFs,t}^P \right)^{1-\sigma_s} \right]^{1/(1-\sigma_s)},
\]  

where \( \hat{P}_{HFs,t}^P \) denotes the level of the optimal reset export price chosen by a Home firm.

Analogously, the sectoral import price index, \( IPI_{Hs} \), captures the industry-level prices for sector-\( s \) goods produced in the Foreign country and sold in Home, measured in the Home currency:

\[
IPI_{Hs,t} = \left( \int_0^1 P_{FHs,t}^P(i)^{1-\sigma_s} di \right)^{1/(1-\sigma_s)} = \left[ \theta_s \left( IPI_{Hs,t-1} \right)^{1-\sigma_s} + (1-\theta_s) \left( \hat{P}_{FHs,t}^P \right)^{1-\sigma_s} \right]^{1/(1-\sigma_s)},
\]  

where \( \hat{P}_{FHs,t}^P \) denotes the level of the optimal reset export price chosen by a Foreign firm.
As observed in equations (7) and (8), neither the export price indices nor the import price indices for the Home country are directly affected by movements in the exchange rate. This is because, following our dollar-pricing assumption, the optimal firm-level export and import prices for the Home country are originally invoiced in the Home currency.

Finally, the definition of the sectoral producer price index, $PPI_{Hs}$, takes into account that PPIs in the U.S. data include changes in prices for exported goods. Thus I define:

$$PPI_{Hs,t} = (DPI_{Hs,t})^{\psi_s} \left( EPI_{Hs,t} \right)^{1-\psi_s},$$

where $\psi_s$ is the domestic share of spending in sector $s$ and $DPI_{Hs,t}$ is the domestic component of the sectoral producer price index (or simply, the sectoral domestic price index), defined as:

$$DPI_{Hs,t} = \left[ \theta_s (DPI_{Hs,t-1})^{1-\sigma_s} + (1-\theta_s) \left( \hat{P}_{HHs,t}^P \right)^{1-\sigma_s} \right]^{1/(1-\sigma_s)},$$

where $\hat{P}_{HHs,t}^P$ is the level of the optimal reset price for domestic sales set by a Home firm.

### 4 Quantitative results

This section first describes the benchmark calibration. I then show the impulse responses of sectoral prices and quantities to an adverse nominal spending shock. Next, I analyze the relevance of the sector-specific nominal rigidities and variable markups in explaining the results. Lastly, I compare the theoretical impulse responses of relative prices and quantities across sectors with the empirical impulse responses shown in section 2.3.

#### 4.1 Calibration

The model is calibrated to monthly data. Table 3 shows the benchmark calibration. Some of the parameters are set to values that are standard in the literature. To match an annual discount rate of 4 percent, I set the discount factor $\beta = 0.96^{1/12}$. The risk aversion
parameter and the inverse of the Frisch elasticity of labor supply take values of $\gamma = 1$ and $\phi = 0.4$, respectively. I calibrate the weight of nontradable goods in total consumption $\chi = 0.53$ to match the share observed in U.S. data, following the calculations in Corsetti et al. (2008). The elasticity of substitution between tradables and nontradables $\omega$ is set to 0.74, as estimated in Mendoza (1991) based on data from industrialized countries.

[Table 3]

For both countries, the persistence of the money supply shock is set to $\rho_M = 0.97$, and the standard deviation to $\sigma_M = 0.33$ percent. These values are estimated by fitting an AR(1) process to seasonally adjusted, HP filtered (log) M2 supply data for the U.S over the sample period 1970:1–2014:6. I assume that the shocks are partially and positively correlated across countries ($\rho_{MH, MF} = 0.5$). This assumption is supported by the evidence of monetary policy synchronization between the U.S. and other advanced economies (see e.g. Scotti, 2011 and Arouri et al., 2013). I checked that using alternative moderate values for $\rho_{MH, MF}$ does not change the main results.

A second set of parameters are chosen to match evidence from the international trade literature that uses the Rauch (1999) classification. I use elasticities of substitution $\sigma_N = 3.2$ and $\sigma_D = 2.3$, which correspond to averages of the median sectoral elasticity estimates for the 1972–1988 and 1990–2001 periods documented in Broda and Weinstein (2006). The value of $\sigma_N$ relies upon Broda and Weinstein’s estimates for goods with reference prices.

Gopinath and Rigobon (2008) calculate the monthly frequency of price adjustment using micro data on U.S. export and import prices at the dock for the period 1994–2005. Since they break down their results by Rauch (1999) categories, I use an average of their median sectoral estimates for export and import prices to set $\theta_N = 0.60$ and $\theta_D = 0.93$. These values imply median price durations of 2.5 and 14.3 months, respectively.\footnote{Since Gopinath and Rigobon (2008) report the number of goods in each Rauch category, I construct a weighted average of their median estimates for organized exchanges and reference priced goods to}
If the sectoral frequencies of price adjustment were instead measured using data for producer prices, the results would be very similar. Using disaggregated data underlying the U.S. PPI for 1994–2005, Nakamura and Steinsson (2008) compute the frequency of price changes by two-digit major groups (see Table VI in their paper). These groups map reasonably well into 3-digit NAICS industries, which allows me to classify them as nondifferentiated or differentiated goods.\footnote{22} The median frequencies of producer price persistence across the groups classified as nondifferentiated and differentiated are 0.63 and 0.96, respectively.

To set the values of a third group of parameters, I rely on new calculations based on several sources of U.S. data, such as disaggregated trade data, input-output (I-O) accounts, the national income and product accounts (NIPAs), and industry concordances from the U.S. Census and the Bureau of Economic Analysis (BEA). Methodological details are provided in Appendix A. Using annual data from the NIPAs for Personal Consumption Expenditure (PCE) by type of product and my own classification of PCE categories of nondurable consumption goods as nondifferentiated and differentiated, I measure the real expenditure share in these two types of goods ($\mu_N$ and $\mu_D$, respectively). For the period 1999–2012, the average expenditure shares are $\mu_N = 0.46$ and $\mu_D = 0.54$. As is standard in the literature (see e.g. Stockman and Tesar, 1995), an implicit assumption is that all goods in the PCE basket for nondurable goods, as opposed to services, are tradable. The use of data on nondurable consumption goods is consistent with the nature of the goods in the model.

Relying on the same classification for PCE categories, I pin down the domestic shares in sectoral consumption $\psi_s$ and the steady-state distribution margin for final consumption at the sector level, which is denoted by $x_s$. (In the steady state, distribution margins for calculate $\theta_N$.)

\footnote{22}The following PPI major groups are considered as nondifferentiated: ‘Farm products’, ‘Processed foods and feeds’, ‘Fuels and related products and power’, and ‘Pulp, paper, and allied products’. The remaining major groups listed by Nakamura and Steinsson (2008) are labeled as differentiated, with the exception of ‘Chemicals and allied products’ and ‘Metals and metal products’, which are left unclassified.
final consumption of domestic and imported goods are equalized within sectors: \( x_s \equiv x_{HHs} = x_{HFs} = x_{FFs} = x_{FHs}. \) Since imports by PCE categories are not published in the NIPAs, I use the import share estimates of McCulley (2011) for the latest year in his analysis, namely 2009. Based on this evidence, I calculate values for \( \psi_N \) and \( \psi_D \) of 0.92 and 0.81, respectively.\(^{23}\)

To compute the sectoral distribution margins \( (x_s = \frac{\kappa_s P_z^s}{\xi_s}) \) I use the BEA’s bridge tables that map PCE categories into the final use categories in the I-O accounts at the most disaggregated level. Following Goldberg and Campa (2010), the distribution margin for sector \( s \) is then calculated as the sum of (wholesale and retail) distribution margins and transportation costs, divided by the sum of all output valued at purchaser’s prices, across all final use categories in the I-O accounts from sector \( s \). Using data from 2007, I obtain aggregate sectoral margins of \( x_N = 0.42 \) and \( x_D = 0.49 \).\(^{24}\) These constitute calibration targets, which conditional on the values for \( \sigma_N \) and \( \sigma_D \) specified above imply sectoral inputs of distribution services of \( \kappa_N = 1.32 \) and \( \kappa_D = 5.67 \).

Figure 5 shows the sector-\( s \) steady-state elasticity of the markup, denoted by \( \Gamma_s \), as a function of steady-state values for the distribution margins \( x_s \). Following equation (3), \( \Gamma_s \) is given by: \( \Gamma_s = \left[ (\sigma_s - 1) \frac{1-x_s}{x_s} - 1 \right]^{-1} \), which is increasing in \( x_s \) and decreasing in the elasticity of substitution \( \sigma_s \). As observed, given that our benchmark calibration considers \( \sigma_N > \sigma_D \), for any given value of the distribution margin \( x_s \) the depicted elasticity of the markup in the nondifferentiated goods sector is smaller than that in the differentiated goods sector, i.e. \( \Gamma_N < \Gamma_D \). Since the evidence on distribution margins suggests that values above 60 percent are rare (Burstein et al., 2003; Goldberg and Campa, 2010), the figure indicates that for a plausible calibration, the model may only predict potentially

\(^{23}\)I use McCulley’s (2011) estimates for nondurable consumption goods categories displayed in Table 1 (shares of PCE by major type of product) and Table 4 (import shares of PCE by major type of product) of his paper—see footnote 22 in McCulley (2011) for methodological details.

\(^{24}\)The baseline values for the distribution margins are robust to the use of other methodologies. Considering average margins across each sector, I obtain \( x_N = 0.40 \) and \( x_D = 0.53 \). As I detail in Appendix A, using more aggregated data from the I-O accounts it is also possible to calculate aggregate sectoral margins for the period 1997–2012, obtaining annual averages of \( x_N = 0.41 \) and \( x_D = 0.49 \).
high values for $\Gamma_D$ but not for $\Gamma_N$.

As indicated by asterisks in the graph, the baseline values for $\Gamma_N$ and $\Gamma_D$ are 0.48 and 2.88, respectively. These numbers are within the range of values that have been used in the literature (see Gopinath and Itskhoki, 2011). This parameterization implies that prices in the differentiated goods sector respond less to movements in the marginal cost but more to changes in the distribution costs relative to prices in the nondifferentiated goods sector. It also implies a lower long-run exchange rate pass-through into import prices of sector $D$ relative to sector $N$ (0.26 and 0.68, respectively), which is in line with available evidence based upon micro data for the U.S. (Gopinath and Itskhoki, 2010). These measures of pass-through, calculated as $1 - \frac{\Gamma_i}{\Gamma_i + \Gamma_s}$, eliminate the effect of nominal rigidities.

The calibration of the model is completed as follows. The frequency of price non-adjustment for nontradable goods $\theta^z$ is set equal to 0.94, based on the median regular price duration for ‘Services (excluding travel)’ reported by Nakamura and Steinsson (2008) for micro data underlying the U.S. CPI. The elasticity of substitution between differentiated and nondifferentiated goods in tradable consumption is set to $\eta = 1.5$.

### 4.2 Responses to a monetary shock

To solve the model, I log-linearize the equilibrium conditions around a zero-inflation steady state and use perturbation methods. I simulate the effect of a demand-driven recession in both countries. I specifically hit the Home country with a negative one-standard-deviation shock to nominal aggregate spending, which is partially transmitted to the Foreign country because of the assumed cross-country shock correlation. The tighter monetary contraction in Home relative to Foreign causes an endogenous appreciation of the Home currency. An
international economic slowdown accompanied by an appreciation of the U.S. dollar have characterized the recent recessions of 2001 and 2007–09.25

The recession scenario for both countries is detailed in Figure 6, which depicts impulse responses of aggregate variables. With sticky prices, the negative shock to nominal spending implies a reduction in overall real consumption. This leads to a decline in the demand for labor, which explains the decline in real and nominal wages. Therefore, marginal costs and distribution costs (not shown) fall. Since the shock originates in Home and is only partially transmitted to the Foreign economy, the magnitude of these effects is greater in the Home country. The appreciation of the Home currency is apparent in the decline of the nominal exchange rate.

[Figure 6]

Figure 7 displays the responses of sectoral PPIs and outputs in the Home country. Given that the U.S. is a relatively closed economy, the PPIs (which embody a component of export price changes, as observed in equation (9)) mostly reflect movements in prices set for local sales. Prices decrease due to the generalized decline in marginal costs and distribution costs, but while the PPI of nondifferentiated goods shows a sizeable decline, the PPI of differentiated goods exhibits a relatively small reduction (see left panel). Given our calibration, this result is explained at least in part by the fact that nominal rigidities are more stringent for differentiated goods producers \((\theta_N < \theta_D)\). Below I scrutinize the role of the variable markup channel.

[Figure 7]

Production in both sectors falls on impact as a consequence of the negative income effect caused by the shock, but the decline in differentiated goods output is larger and

25Mussa (1986) is a classic reference for evidence that fluctuations in nominal and real exchange rates are driven by monetary shocks. For more recent evidence, see e.g. Bouakez and Normandin (2010). Using a VAR analysis, Eichenbaum and Evans (1995) find empirically that contractionary shocks to U.S. monetary policy lead to persistent and significant appreciations in U.S. nominal and real exchange rates.
more protracted than that of nondifferentiated goods (right panel in Figure 7). This heterogeneity in the output responses across sectors arises because the price ratio of differentiated goods to nondifferentiated goods increases after the shock, leading to a substitution effect that biases consumption, and hence production, towards the latter.

The responses of sectoral trade variables for the Home country to a contractionary monetary shock are depicted in Figure 8. Similar to the patterns exhibited by the domestic variables, the export and import prices (quantities) fall disproportionately in the nondifferentiated (differentiated) goods sector. The trade price indices decline as a result of the reductions in local marginal costs and distribution costs associated with the declines in wages. Moreover, the appreciation of the Home currency contributes both to the reduction in distribution costs faced by Home exporters and to the decline in production costs (in units of Home currency) faced by Foreign exporters. Meanwhile, as in the case of domestic production discussed above, the observed responses of sectoral trade volumes reflect a negative income effect as well as a substitution of nondifferentiated goods for differentiated goods in the wake of the adjustment of the relative trade prices.

[Figure 8]

In sum, the model predicts that a demand-driven recession in the U.S. features declines in the relative prices of nondifferentiated goods to differentiated goods and in the relative quantities of differentiated goods to nondifferentiated goods. These results suggest that price-setting frictions are indeed relevant to explain the patterns observed in the data.

4.3 The key mechanisms: nominal rigidities and variable markups

To examine the quantitative role of the key mechanisms of the model, Table 4 displays the cumulated impulse responses of several price measures to a contractionary monetary shock for three alternative model specifications. The responses are accumulated over a 36-month horizon, but the results are not affected if I consider a different horizon. In
addition to the baseline model that assumes the benchmark calibration, I consider a model with “constant markups”, in which the variable markup channel is shut down by assuming calibration targets for the distribution margins very close to zero \((x_N = x_D \approx 0)\), so that markup elasticities are zero as well \((\Gamma_N = \Gamma_N = 0)\). I also compute a model with a uniform degree of nominal rigidity, in which the same frequency of price adjustment is used for both sectors (dubbed as “Same Calvo parameters” in the Table). This model assumes \(\theta_N = \theta_D = 0.93\), the value for the Calvo parameter in the differentiated goods sector employed in the benchmark calibration.

For each of these models, Table 4 shows the cumulated sectoral responses of firm-level prices—optimal producer prices under flexible prices and optimal reset prices—and industry-level price indices corresponding to domestic and international transactions. The columns labeled as ‘\(\text{Gap} \ N - D\)’ report the cumulated responses of the relative prices of nondifferentiated goods to differentiated goods.

A first result that stands out is that the model with constant markups consistently shows larger sectoral price declines than the baseline model, which is in line with the intuition that, by offsetting the effects of movements in marginal costs, variable markups generate price stickiness above and beyond the nominal rigidity channel. Table 4 also reveals that with constant markups, the responses of export prices are the same as those of domestic prices. This is because with PCP in exports, producers do not “price to market” in the absence of local distribution costs.

Furthermore, focusing on the industry-level price indices (bottom part of the Table), the relative prices of nondifferentiated goods to differentiated goods exhibit similar magnitudes of adjustment in the baseline model as in the model with constant markups. In contrast, the model with a unique Calvo parameter shows only small changes in these relative prices. For example, the PPI of nondifferentiated goods relative to that of differ-
entiated goods decreases around 3 percent both in the baseline model and in the model with constant markups, but it only declines by 0.6 percent in the model with a single Calvo parameter. Importantly, this evidence implies that the bulk of heterogeneity in the baseline impulse responses of sectoral prices to a monetary shock is accounted for by the nominal rigidity channel. In other words, a model with variable markups and homogenous frequencies of price adjustment across sectors only produces a small degree of heterogeneity in the responses of sectoral prices and quantities.  

Although the variable markup channel by itself fails to generate significant movements in the relative prices between sectors at the industry level, this mechanism does generate a higher degree of asymmetry in the sectoral price responses at the firm level—particularly among optimal prices in a flexible-price environment. To illustrate this, observe that in the top part of Table 4, the cumulated declines in the relative optimal prices (under flexible prices) are roughly between 1 and 2 percent both in the baseline model and in the model with a unique Calvo parameter. In line with our previous discussion, this result is attributable to the higher elasticity of the markup in sector $D$, which makes optimal prices in this sector less sensitive to the reduction in marginal costs than those in sector $N$.

But as the results for the model with a single Calvo parameter indicate, the variable markup channel generates smaller movements in the relative prices between sectors if we focus instead on the optimal reset prices (middle part of Table 4) and the industry-level price indices (bottom part). This is explained by the interaction of two mechanisms. First, nominal rigidities have compounding effects, which reduce the sensitivity of all sectoral prices to changes in the economic environment (e.g., to the decline in marginal costs).  

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26 Other papers have also found that the quantitative macro effects of real rigidities in the form of variable markups are modest. For instance, Gopinath and Itskhoki (2011) use a model with strategic complementarities in price setting at the producer level, and arrive at this conclusion by simulating the dynamics of a closed economy model featuring idiosyncratic and aggregate shocks.

27 The compounding effects of nominal rigidities arise because, as observed in equations (4), (7), and (8), the optimal reset prices and the industry-level price indices are functions of the degree of nominal rigidity measured by the Calvo parameter $\theta_s$. 

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Second, a higher elasticity of the markup in sector D increases the sensitivity of prices in this sector to the reduction in distribution costs relative to sector N.

To complement the analysis, Figure 9 considers the same three alternative model specifications, and depicts the dynamic responses of relative prices (at the sectoral level) and relative quantities to an adverse nominal shock. Consistent with the results in Table 4, we see that after the shock, the three models predict that the relative prices of nondifferentiated goods to differentiated goods (in terms of the PPIs, export price indices, and import price indices) decrease in the short run. Figure 9 also shows that all of the models predict declines in the relative quantities produced and relative trade volumes of differentiated goods to nondifferentiated goods.

It is apparent from Figure 9 that a model with variable markups and homogenous frequencies of price adjustment across sectors (dashed green lines) only produces a small degree of heterogeneity in the responses of sectoral prices and quantities. Thus, this graphical evidence confirms that the nominal rigidity channel matters quantitatively more than the variable markup channel in explaining the results of the baseline model. Meanwhile, a model with constant markups (dotted black lines) slightly magnifies the responses of the relative prices and quantities as compared to the baseline model (solid blue lines). As mentioned before, this is due to the compounding effects of the cross-sector nominal rigidities and to the differences in the sectoral markup elasticities to changes in distribution costs.

4.4 Model-based vs. empirical impulse response functions

Finally, I evaluate the model in terms of its ability to fit the VAR evidence. Figure 10 compares the empirical impulse responses of relative prices and quantities to a tightening monetary shock, as reported in section 2.3, with their model-based counterparts. The
latter are calculated from the model’s IRFs corresponding to sectoral prices (PPIs and import price indices) and quantities (outputs and import volumes). As observed, in the first few months after the shock, the theoretical IRFs are either within or very close to the 90% confidence regions estimated in the VARs, so the match in terms of magnitude is reasonably good. Naturally, though, the model cannot replicate some of the delayed and hump-shaped empirical IRFs, as we ignore mechanisms which have proved useful in improving the empirical fit of monetary models (e.g., habit formation, capital accumulation and adjustment costs).

[Figure 10]

To analyze the quantitative effects of a monetary shock in more detail, Table 5 displays three summary measures for the empirical and theoretical impulse responses: the maximum monthly response with expected sign observed within the first year after the shock, the average response in the first 12 months after the shock, and the response recorded 6 months after the shock. Again, I report model-based IRFs using three versions of the model: the baseline model, a model with constant markups, and a model with the same frequency of price adjustment across sectors. Importantly, the main conclusions from this comparative analysis are robust to the use of other values for the Calvo parameters and the distribution-margin targets (see additional experiments in Appendix C).

[Table 5]

From Table 5, it is evident that regardless of the measure that is utilized, the model with a unique Calvo parameter severely underestimates the magnitudes of the empirical IRFs. In contrast, both the baseline model and the model with constant markups make considerable progress in terms of matching the empirical impulse responses. Focusing on the short-run responses with expected signs, the model is particularly successful in approximating the quantitative responses of the relative PPIs and the relative import
volumes. Thus, for example, 6 months after the shock, the relative producer price declines by 0.18 percentage points in the model and by 0.25 percentage points in the VAR. At the same horizon, the relative import volume falls by 0.34 percentage points in the model and by 0.38 percentage points in the VAR. That said, the model shows some limitations in replicating the subdued reduction in relative quantities produced and, especially, the sizeable decline in relative import prices.

5 Conclusions

This paper studies the effects of an aggregate nominal demand shock on prices and quantities of nondifferentiated and differentiated manufactured goods. I use U.S. data on producer prices and industrial production, as well as data on unit values and volumes of trade, to construct measures of prices and quantities for these sectors. A key empirical finding is that during recessions and in the wake of a contractionary monetary policy shock, both the relative price of nondifferentiated goods to differentiated goods and the relative quantity of differentiated goods to nondifferentiated goods decline.

I show that a two-country model featuring heterogeneity in price stickiness generates meaningful differences in the adjustment of sectoral prices and quantities following a demand disturbance. The model-based impulse responses of the relative prices and quantities across sectors to a tightening monetary shock are in line with the empirical evidence based on an estimated VAR. The numerical exercises also show that the differences in sectoral price durations, which are calibrated using micro data, constitute the central mechanism behind the quantitative findings. Although the model includes variable markups based on distribution costs, this mechanism by itself matters relatively little to the main results.

Overall, these findings add to the existing evidence in the closed economy literature that heterogeneity in price stickiness can explain significant differences in the responses of
sectoral prices and output to a monetary shock. I extend this conclusion to the analysis of nondifferentiated and differentiated goods sectors and to an open economy context. The results also imply that price-adjustment frictions help explain why domestic and trade quantities of differentiated goods declined relatively more than those of nondifferentiated goods in recent recessions. More generally, I conclude that the incorporation of sectoral differences in nominal rigidities in the quantitative analysis of macroeconomic fluctuations and international trade, within the context of multi-sector frameworks, is a promising avenue for further research.

To improve the empirical fit of the model developed in this paper, several extensions may be considered—e.g., the incorporation of features such as trade in intermediate and capital goods, input-output production linkages, and sector-specific shocks. Another topic for future work is the explicit modeling of a rationale for the different frequencies of price adjustment exhibited by nondifferentiated and differentiated goods, rather than assuming such differences exogenously, as I have done in this study.

References


## Tables and Figures

### Table 1

**Classification of manufacturing industries by product differentiation at the 4-digit NAICS level**

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<th>Code</th>
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</tr>
<tr>
<td>3112</td>
<td>N</td>
<td>Grain and Oilseed Milling Products</td>
</tr>
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<td>3113</td>
<td>U</td>
<td>Sugar and Confectionery Products</td>
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<td>3114</td>
<td>U</td>
<td>Fruit and Vegetable Preserves and Specialty Foods</td>
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<tr>
<td>3115</td>
<td>N</td>
<td>Dairy Products</td>
</tr>
<tr>
<td>3116</td>
<td>N</td>
<td>Meat Products and Meat Packaging Products</td>
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<td>3117</td>
<td>N</td>
<td>Seafood Products Prepared, Canned and Packaged</td>
</tr>
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<td>3118</td>
<td>D</td>
<td>Bakery and Tortilla Products</td>
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<tr>
<td>3119</td>
<td>D</td>
<td>Foods, NESOI</td>
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<tr>
<td>3121</td>
<td>N</td>
<td>Beverages</td>
</tr>
<tr>
<td>3122</td>
<td>N</td>
<td>Tobacco Products</td>
</tr>
<tr>
<td>3131</td>
<td>N</td>
<td>Fibers, Yarns, and Threads</td>
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<tr>
<td>3132</td>
<td>D</td>
<td>Fabrics</td>
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<td>Veneer, Plywood, and Engineered Wood Products</td>
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<td>3219</td>
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<td>3221</td>
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<td>Pulp, Paper, and Paperboard Mill Products</td>
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<td>Converted Paper Products</td>
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<td>D</td>
<td>Printed Matter and Related Product, NESOI</td>
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<td>3241</td>
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<td>3251</td>
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<td>Basic Chemicals</td>
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<td>Resin, Synthetic Rubber, &amp; Artificial &amp; Synthetic Fibers &amp; Filiment</td>
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<td>3253</td>
<td>N</td>
<td>Pesticides, Fertilizers and Other Agricultural Chemicals</td>
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<td>3256</td>
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<td>Soaps, Cleaning Compounds, and Toilet Preparations</td>
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<td>3311</td>
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Table 1 (continued)

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<td>N</td>
<td>Nonferrous Metal (Except Aluminum) and Processing</td>
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<td>3315</td>
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<td>3321</td>
<td>D</td>
<td>Crowns, Closures, Seals and Other Packing Accessories</td>
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<td>3322</td>
<td>D</td>
<td>Cutlery and Handtools</td>
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<td>3366</td>
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<td>3399</td>
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<td>Miscellaneous Manufactured Commodities</td>
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**Notes:** Calculations based on U.S. trade data, the conservative classification of goods by Rauch (1999), and the HS-NAICS-SITC concordance from Feenstra et al. (2002). The first column is the 4-digit NAICS code of the industry. The second column indicates the classification of the industry: nondifferentiated goods (“N”), differentiated goods (“D”), or unclassified (“U”). See further details in the methodological appendix.
# Table 2

Cumulated changes in industrial production indices and producer price indices during recessions and selected sub-periods, by sector (%)

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<td>Diff.</td>
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<td></td>
<td></td>
</tr>
<tr>
<td>Nondiff.</td>
<td>n/a</td>
<td>3.1</td>
<td>-5.0</td>
<td>-3.8</td>
<td>-2.7</td>
<td>-1.9</td>
<td>-15.0</td>
</tr>
<tr>
<td>Diff.</td>
<td>n/a</td>
<td>6.9</td>
<td>4.0</td>
<td>1.8</td>
<td>0.7</td>
<td>6.9</td>
<td>0.8</td>
</tr>
</tbody>
</table>

*Notes:* The table reports the percentage (log) changes cumulated during NBER recession dates and the acute phase of the 2007–2009 recession (August 2008 to June 2009). The underlying monthly data for production corresponds to the sectoral industrial production indices for nondifferentiated and differentiated goods. The underlying monthly data for prices corresponds to the sectoral producer price indices for nondifferentiated and differentiated goods.
### Table 3

**Benchmark calibration**

<table>
<thead>
<tr>
<th>Parameter</th>
<th>Value</th>
<th>Description/Source/Target</th>
</tr>
</thead>
<tbody>
<tr>
<td>$\beta$</td>
<td>0.997</td>
<td>Discount factor.</td>
</tr>
<tr>
<td>$\gamma$</td>
<td>1</td>
<td>CRRA.</td>
</tr>
<tr>
<td>$\phi$</td>
<td>0.40</td>
<td>Inverse Frisch elasticity.</td>
</tr>
<tr>
<td>$\chi$</td>
<td>0.53</td>
<td>Share of tradable goods in consumption. Source: Corsetti et al. (2008).</td>
</tr>
<tr>
<td>$\omega$</td>
<td>0.74</td>
<td>Elast. of subst. between tradables and nontradables. Source: Mendoza (1991).</td>
</tr>
<tr>
<td>$\eta$</td>
<td>1.50</td>
<td>Elast. of subst. between diff. and nondiff. goods.</td>
</tr>
<tr>
<td>$\mu_D$</td>
<td>0.54</td>
<td>Share of diff. goods in tradable consumption. Source: NIPAs.</td>
</tr>
<tr>
<td>$\psi_D$</td>
<td>0.81</td>
<td>Domestic share in diff. goods consumption. Source: NIPAs and McCulley (2011).</td>
</tr>
<tr>
<td>$\psi_N$</td>
<td>0.92</td>
<td>Domestic share in nondiff. goods consumption. Source: NIPAs and McCulley (2011).</td>
</tr>
<tr>
<td>$\sigma_D$</td>
<td>2.3</td>
<td>Elast. of subst. across diff. varieties. Source: Broda and Weinstein (2006).</td>
</tr>
<tr>
<td>$\sigma_N$</td>
<td>3.2</td>
<td>Elast. of subst. across nondiff. varieties. Source: Broda and Weinstein (2006).</td>
</tr>
<tr>
<td>$\kappa_D$</td>
<td>5.67</td>
<td>Units of distrib. services in retail diff. goods. Calibration target: aggregate distrib. margin in final consumption, diff. goods ($x_D = 0.49$; source: 2007 input-output tables).</td>
</tr>
<tr>
<td>$\kappa_N$</td>
<td>1.32</td>
<td>Units of distrib. services in retail nondiff. goods. Calibration target: aggregate distrib. margin in final consumption, nondiff. goods ($x_N = 0.42$; source: 2007 input-output tables).</td>
</tr>
<tr>
<td>$\theta_D$</td>
<td>0.93</td>
<td>Frequency of price non-adjustment, diff. goods. Source: Gopinath and Rigobon (2008) and Nakamura and Steinsson (2008).</td>
</tr>
<tr>
<td>$\theta_N$</td>
<td>0.60</td>
<td>Frequency of price non-adjustment, nondiff. goods. Source: Gopinath and Rigobon (2008) and Nakamura and Steinsson (2008).</td>
</tr>
<tr>
<td>$\theta^z$</td>
<td>0.94</td>
<td>Frequency of price non-adjustment, nontradable goods. Source: Nakamura and Steinsson (2008).</td>
</tr>
<tr>
<td>$\rho_{MH}$, $\rho_{MF}$</td>
<td>0.97</td>
<td>Persistence monetary shock. Source: M2 data.</td>
</tr>
<tr>
<td>$\sigma_{MH}$, $\sigma_{MF}$</td>
<td>0.003</td>
<td>Std. dev. monetary shock. Source: M2 data.</td>
</tr>
<tr>
<td>$\rho_{MH, MF}$</td>
<td>0.5</td>
<td>Cross-country correlation between monetary shocks.</td>
</tr>
</tbody>
</table>

*Note*: See section 4.1 for further details.
Table 4
Cumulated responses of sectoral and relative prices to contractionary nominal spending shock (Home country), under alternative model specifications (%)

<table>
<thead>
<tr>
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</thead>
<tbody>
<tr>
<td><strong>Optimal prices under flexible prices</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
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<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Domestic ($\tilde{p}_{HHs}$)</td>
<td>-6.4</td>
<td>-4.6</td>
<td>-1.8</td>
<td>-7.9</td>
<td>-7.9</td>
<td>0.0</td>
<td>-6.8</td>
<td>-4.9</td>
<td>-2.0</td>
</tr>
<tr>
<td>Export ($\tilde{p}_{HFs}$)</td>
<td>-7.3</td>
<td>-6.7</td>
<td>-0.7</td>
<td>-7.9</td>
<td>-7.9</td>
<td>0.0</td>
<td>-7.6</td>
<td>-6.6</td>
<td>-1.0</td>
</tr>
<tr>
<td>Import ($\tilde{p}_{FHs}$)</td>
<td>-6.8</td>
<td>-4.7</td>
<td>-2.1</td>
<td>-8.8</td>
<td>-8.8</td>
<td>0.0</td>
<td>-6.8</td>
<td>-4.9</td>
<td>-2.0</td>
</tr>
<tr>
<td><strong>Optimal reset prices</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Domestic ($\hat{p}<em>{HH</em>{s,t}}$)</td>
<td>-6.1</td>
<td>-3.8</td>
<td>-2.3</td>
<td>-7.5</td>
<td>-5.3</td>
<td>-2.1</td>
<td>-4.8</td>
<td>-4.0</td>
<td>-0.8</td>
</tr>
<tr>
<td>Export ($\hat{p}<em>{H</em>{Fs,t}}$)</td>
<td>-7.0</td>
<td>-5.2</td>
<td>-1.8</td>
<td>-7.5</td>
<td>-5.3</td>
<td>-2.1</td>
<td>-5.3</td>
<td>-5.2</td>
<td>-0.2</td>
</tr>
<tr>
<td>Import ($\hat{p}<em>{FH</em>{s,t}}$)</td>
<td>-6.6</td>
<td>-4.1</td>
<td>-2.5</td>
<td>-8.4</td>
<td>-6.4</td>
<td>-2.1</td>
<td>-5.2</td>
<td>-4.1</td>
<td>-1.1</td>
</tr>
<tr>
<td><strong>Industry-level price indices</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Domestic (dpi$_{Hs}$)</td>
<td>-6.0</td>
<td>-2.7</td>
<td>-3.3</td>
<td>-7.3</td>
<td>-4.0</td>
<td>-3.4</td>
<td>-3.5</td>
<td>-2.8</td>
<td>-0.7</td>
</tr>
<tr>
<td>Producer (ppi$_{Hs}$)</td>
<td>-6.0</td>
<td>-2.9</td>
<td>-3.2</td>
<td>-7.3</td>
<td>-4.0</td>
<td>-3.4</td>
<td>-3.6</td>
<td>-3.0</td>
<td>-0.6</td>
</tr>
<tr>
<td>Export (epi$_{Hs}$)</td>
<td>-6.8</td>
<td>-3.7</td>
<td>-3.1</td>
<td>-7.3</td>
<td>-4.0</td>
<td>-3.4</td>
<td>-3.9</td>
<td>-3.7</td>
<td>-0.2</td>
</tr>
<tr>
<td>Import (ipi$_{Hs}$)</td>
<td>-6.4</td>
<td>-2.8</td>
<td>-3.6</td>
<td>-8.2</td>
<td>-4.6</td>
<td>-3.6</td>
<td>-3.7</td>
<td>-2.9</td>
<td>-0.8</td>
</tr>
</tbody>
</table>

Notes: Cumulated responses at a 36-month horizon after the shock. The model with constant markups sets the distribution margins to a number very close to zero ($x_N = x_D \approx 0$), so that the elasticities of the markups are zero ($\Gamma_N = \Gamma_N = 0$). The model with homogeneous Calvo parameters assumes the same frequency of price adjustment in both sectors: $\theta_N = \theta_D = 0.93$. (In the baseline calibration, 0.93 is the value of the Calvo parameter for sector $D$.)
### Table 5
Summary measures of theoretical and empirical impulse responses to contractionary nominal spending shock: relative prices and quantities, Home country (%)

<table>
<thead>
<tr>
<th></th>
<th>Empirical</th>
<th>Baseline model</th>
<th>Constant markups</th>
<th>Same Calvo parameter</th>
</tr>
</thead>
<tbody>
<tr>
<td><strong>Maximum response with expected sign (1 year after shock)</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Rel. PPI</td>
<td>-0.27</td>
<td>-0.19</td>
<td>-0.25</td>
<td>-0.03</td>
</tr>
<tr>
<td>Rel. Production</td>
<td>-0.29</td>
<td>-0.32</td>
<td>-0.39</td>
<td>-0.04</td>
</tr>
<tr>
<td>Rel. Import price</td>
<td>-0.97</td>
<td>-0.19</td>
<td>-0.25</td>
<td>-0.03</td>
</tr>
<tr>
<td>Rel. Import volume</td>
<td>-0.38</td>
<td>-0.34</td>
<td>-0.43</td>
<td>-0.04</td>
</tr>
<tr>
<td><strong>Average response 1 year after shock</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Rel. PPI</td>
<td>-0.14</td>
<td>-0.16</td>
<td>-0.19</td>
<td>-0.02</td>
</tr>
<tr>
<td>Rel. Production</td>
<td>-0.15</td>
<td>-0.28</td>
<td>-0.32</td>
<td>-0.02</td>
</tr>
<tr>
<td>Rel. Import price</td>
<td>-0.32</td>
<td>-0.16</td>
<td>-0.19</td>
<td>-0.02</td>
</tr>
<tr>
<td>Rel. Import volume</td>
<td>0.08</td>
<td>-0.30</td>
<td>-0.36</td>
<td>-0.03</td>
</tr>
<tr>
<td><strong>Response 6 months after shock</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Rel. PPI</td>
<td>-0.25</td>
<td>-0.18</td>
<td>-0.21</td>
<td>-0.02</td>
</tr>
<tr>
<td>Rel. Production</td>
<td>-0.13</td>
<td>-0.31</td>
<td>-0.36</td>
<td>-0.03</td>
</tr>
<tr>
<td>Rel. Import price</td>
<td>-0.97</td>
<td>-0.18</td>
<td>-0.22</td>
<td>-0.02</td>
</tr>
<tr>
<td>Rel. Import volume</td>
<td>-0.38</td>
<td>-0.34</td>
<td>-0.42</td>
<td>-0.03</td>
</tr>
</tbody>
</table>

**Notes:** Relative prices (‘Rel. PPI’ and ‘Rel. Import price’) are defined as prices of nondifferentiated goods relative to differentiated goods. Relative quantities (‘Rel. production’ and ‘Rel. Import volume’) are defined as quantities of differentiated goods relative to nondifferentiated goods. The values in second column (‘Empirical’) are based on central estimates of the VAR-based impulse response functions (IRFs) depicted in Figures 3 and 4. The remaining three columns correspond to theoretical IRFs for different model specifications. ‘Baseline model’ uses the benchmark calibration. The model with ‘Constant markups’ sets the distribution margins to a number close to zero ($x_N = x_D \approx 0$), so the elasticities of the markups are zero ($\Gamma_N = \Gamma_N = 0$). The model with ‘Same Calvo parameter’ assumes the same frequency of price adjustment for both sectors: $\theta_N = \theta_D = 0.93$. See text for further details.
Figure 1
Industrial production (IP) index by sector:
Percentage (log) changes in smoothed indices, 1972:1–2014:6

Notes: The graph shows changes in the logs of sectoral industrial production indices, smoothed using 9-month centered moving averages, multiplied by 100. The shaded areas indicate NBER recession dates. See text for further details.
Figure 2
Producer price index (PPI) by sector:

Notes: The graph shows changes in the logs of sectoral PPIs, smoothed using 9-month centered moving averages, multiplied by 100. The shaded areas indicate NBER recession dates. See text for further details.
Figure 3
Empirical impulse responses to contractionary monetary policy shock: VAR with domestic data

Notes: The horizontal axis indicates the months after the shock. The variables are measured in percent. The gray area represents the corresponding 90% probability bands.
Figure 4
Empirical impulse responses to contractionary monetary policy shock: VAR with import data

Notes: The horizontal axis indicates the months after the shock. The variables are measured in percent. The gray area represents the corresponding 90% probability bands.
Figure 5
Sensitivity of elasticities of markups to distribution margins

Notes: The depicted values assume $\sigma_N = 3.2$ and $\sigma_D = 2.3$. The asterisk on each curve indicates the parameter value implied by the benchmark calibration targets for the distribution margins $x_N$ and $x_D$. 
Figure 6
Theoretical impulse responses to contractionary nominal spending shock: aggregate variables

Notes: The horizontal axis indicates the months after the shock. The variables are measured in percent.
Figure 7
Theoretical impulse responses to contractionary nominal spending shock: sectoral domestic variables, Home country

Notes: The horizontal axis indicates the months after the shock. The variables are measured in percent.
Figure 8
Theoretical impulse responses to contractionary nominal spending shock: sectoral trade variables, Home country

Notes: The horizontal axis indicates the months after the shock. The variables are measured in percent.
Figure 9
Theoretical impulse responses to contractionary nominal spending shock: relative prices and quantities under alternative models, Home country

Notes: The horizontal axis indicates the months after the shock. The variables are measured in percent. The model with constant markups sets the distribution margins to a number very close to zero ($x_N = x_D \approx 0$), so that the elasticities of the markups are zero ($\Gamma_N = \Gamma_N = 0$). The model with homogeneous Calvo parameters assumes the same frequency of price adjustment in both sectors: $\theta_N = \theta_D = 0.93$. (In the baseline calibration, 0.93 is the value of the Calvo parameter for sector $D$.)
Figure 10

Theoretical and empirical impulse responses to contractionary nominal spending shock: relative prices and quantities, Home country

Notes: The horizontal axis indicates the months after the shock. The variables are measured in percent. The empirical IRFs (solid lines) and their 90% probability bands (grey areas) are the same as in Figures 3 and 4. They correspond to the responses to tightening monetary shocks in VAR models estimated on U.S. data; see details in the text. Following the notation of the model, the theoretical IRFs (dashed lines) are calculated as: \( ppi_{HN} - ppi_{HD} \) (relative producer prices), \( y_{HD} - y_{HN} \) (relative quantities produced), \( ipi_{HN} - ipi_{HD} \) (relative import prices), and \( c_{FHD} - c_{FHN} \) (relative import volumes).
Appendix

A Methodological details

Classification of NAICS industries as nondifferentiated and differentiated. To map the Rauch (1999) categories into several levels of aggregation in the NAICS classification of manufacturing industries (see Table 1), I use the 4-digit Standard International Trade Classification Rev.2 (SITC) to 10-digit Harmonized System (HS10) concordance and the HS10 to 6-digit NAICS concordance from Feenstra et al. (2002). The Rauch classification is obtained from Jon Haveman’s International Trade Data web page. Rauch classifies goods within three types of product categories depending on their dominant method of sale: products traded on organized exchanges, which are essentially commodities; goods that have a reference price (i.e., they are listed in trade catalogues); and differentiated products, which cannot be priced by either of the former means. He also provides a “conservative” and a “liberal” classification—the former minimizes the number of goods classified as organized exchanges or reference prices, while the latter maximizes their number. I use the conservative classification, but my results are unchanged if I use the liberal one. I specifically employ disaggregated data on U.S. trade flows from 2006 assembled by Robert Feenstra, available on http://www.internationaldata.org/. The original source of Feenstra’s data is the U.S. Census Bureau. In particular, I use the export and import values at the HS10 level to label a given NAICS code as nondifferentiated, differentiated or unclassified in the following way: if a certain Rauch category captures at least 70 percent of trade of a NAICS aggregate, I assign that Rauch category to the industry in question; otherwise, I leave the NAICS industry unclassified. This mapping method is developed for the 3-, 4-, 5- and 6-digit NAIC Systems. I checked that the results are not affected by using data on U.S. trade flows from other years (in particular, 2001 and 1996).

Construction of measures of sectoral trade prices and volumes. As explained in section 2.3, I construct measures of sectoral trade volumes and prices (unit values)
using disaggregated data on U.S. exports and imports from the USITC at the 4-digit NAICS level. Since the USITC reports trade volume data for all of the different units in which transactions are recorded (e.g., kilograms, liters, dozens, etc.), a unique NAICS code frequently features non-zero monthly observations for several different units. For simplicity, to construct the sectoral measures of volumes and prices I work with the unit that captures the largest fraction of total trade value within each NAICS code—in most cases, this means the unit that captures over 50 percent of trade in each industry. Naturally, the predominant unit is not necessarily the same across all industries, so before computing the sectoral averages I transform the trade volume series at the industry level into indices normalized to a common base period. To be consistent, I also transform the industry-level series for unit values (originally reported in dollars) into indices. (Although the BLS publishes data on export and import price indices at the 4-digit NAICS level, these series are only available since 2005, so their number of observations was insufficient to be used in the econometric analysis.)

Calibration of sectoral expenditure shares. To classify the NIPAs for nondurable PCE categories as nondifferentiated and differentiated, I map the Rauch categories into the commodities in the input-output (I-O) accounts included in each PCE category. I proceed in three steps. First, I use my own classification of NAICS industries as nondifferentiated and differentiated, described earlier in this Appendix.

Second, I use the concordance from NAICS codes (at the 3-, 4-, 5-, and 6-digit level) to the I-O codes, which is available from the BEA. Although this concordance often yields a many-to-one mapping, most of the times an I-O code is matched with different NAICS codes which are characterized by a unique Rauch category. Conversely, when different Rauch categories map into a given I-O code (or when the concordance method does not match an I-O code with any NAICS code), I either classify the I-O codes manually or leave them unclassified. As a result of this procedure, I classify 91 (out of 98) I-O codes as either nondifferentiated or differentiated.

Finally, I use the bridge tables that map the PCE categories into the final use categories in the I-O accounts, which are available from the BEA. These bridge tables feature, among
other information, the I-O commodities included in each PCE category and the value of the transactions in purchasers’ prices of each I-O commodity. The latest annual tables at the most disaggregated level (or “detail level”, in BEA’s terminology) correspond to 2007. I use data at the detail level from that year to map the Rauch categories into the PCE categories. Specifically, if at least 70 percent of the total purchasers’ value for a PCE category correspond to, say, nondifferentiated I-O commodities, then that PCE category is labeled as nondifferentiated. If this 70 percent criterion is not satisfied, the PCE category is left unclassified.

By combining the Rauch classification of PCE categories with the NIPA annual data for PCE by type of product (NIPA table 2.4.6U), I calculate the average expenditure shares in nondifferentiated and differentiated consumption goods for the period 1999–2012. This calculation ignores the PCE categories which are classified neither as nondifferentiated nor as differentiated to ensure that the sectoral expenditure shares in the model add up to one. (In the raw data, the unclassified PCE categories represent roughly 10 percent of total nondurable consumption.)

**Calibration of sectoral import shares.** I calculate the sectoral import shares $1 - \psi_s$ as the sum product of import shares of sector-$s$ goods and the relative weights of these sector-$s$ goods in their respective sectoral consumption basket. A detailed explanation of the procedure is in order (see accompanying Table A1). I use the import share estimates for nondurable major PCE categories reported by McCulley (2011) for 2009 (column (2) in Table A1). These major PCE categories are: ‘Food and beverages purchased off-premises consumption’, ‘Clothing and footwear’, ‘Gasoline and other energy goods’, and ‘Other Nondurables’.

[Table A1]

I compute the differentiated and nondifferentiated content of each major PCE product using the Rauch classification of PCE categories at the *detail* level (described earlier in this Appendix for the calculation of the sectoral expenditure shares) and the value of transactions in purchasers’ prices within each PCE category at the detail level (columns
For some major PCE categories the values in columns (3) and (6) do not add up to 100. This is because, as explained above, some detail-level PCE categories are classified neither as nondifferentiated nor as differentiated. Due to the lack of disaggregated information, I assume the same import shares for both the nondifferentiated and the differentiated content within each type of major product.

For each major PCE category, the relative weight of sector-s goods in the sectoral consumption basket (columns (5) and (8)) is calculated as the ratio of the relative weight of sector-s goods in PCE (columns (4) and (7)), to the relative weight of the overall sector-s consumption basket in PCE (sums of columns (4) and (7)). Note that both column (5) and column (8) add up to 100. As part of these calculations I use the shares of PCE by major type of product reported by McCulley (2011) (column (1)).

Calibration of sectoral distribution margins. To compute the sectoral distribution margins, I use the 2007 bridge tables at the detail level that map the PCE categories into the final use categories in the I-O accounts (source: BEA). These tables show the I-O commodities included in each PCE category, the value of the transactions purchasers’ prices, and the associated transportation costs and trade margins. As explained above for the calculation of the sectoral expenditure shares, I develop a classification of I-O commodities as differentiated and nondifferentiated goods. Relying on this classification and on the 2007 PCE bridge tables, I compute the baseline sectoral distribution margins in the way detailed in the text.

The BEA also reports annual PCE bridge tables at a more aggregate level (“summary level”, in BEA’s terminology), which are available for the period 1997–2012. Using these data, I compute annual sectoral distribution margins for robustness purposes. The results are summarized in Figure A1. For this exercise, the summary-level PCE categories corresponding to the following nondurable major PCE products are coded as nondifferentiated: ‘Food and beverages purchased for off-premises consumption’, ‘Gasoline and other energy goods’, and ‘Tobacco’; whereas the PCE categories that belong to the following

---

28The NIPA tables for PCE by type of product published by the BEA imply the following mapping from the major PCE products to the NIPA lines for detail-level PCE categories: ‘Food and beverages...’: lines 72-101; ‘Clothing and footwear’: lines 103-110; ‘Gasoline and other energy goods’: lines 112-117; and ‘Other Nondurables’: lines 120-142.
major products are coded as differentiated: ‘Clothing and footwear’, ‘Pharmaceutical and other medical products’, ‘Household supplies’, ‘Personal care products’, and ‘Magazines, newspapers, and stationery’.\footnote{The PCE categories within the remaining major product ‘Recreational Items’ are left unclassified because the nondifferentiated content of this major product is virtually the same as its differentiated content, as revealed by my Rauch-type classification of PCE categories at the detail level.}

**B Model: market clearing conditions**

Total labor in the nontradable sector is given by $L_{H,t}^z \equiv \int_0^1 L_{H,t}^z(i)di$. Using this definition along with the technology function for nontradables, the goods market clearing condition in that sector can be written as:

$$L_{H,t}^z = \left( C_{H,t}^z + D_{H,t}^z \right) \int_0^1 \left( \frac{P_{H,t}^z(i)}{P_{H,t}^z} \right)^{-\xi} di,$$

where the integral term is a measure of sectoral price dispersion (which is equal to zero to a first-order approximation in a neighborhood of the zero-inflation steady state), and where market clearing in the distribution sector implies:

$$D_{H,t}^z = \sum_{s=N,D} \kappa_s \left[ C_{HHs,t} \int_0^1 \left( \frac{P_{HHs,t}(i)}{P_{HHs,t}} \right)^{-\sigma_s} di + C_{HFHs,t} \int_0^1 \left( \frac{P_{HFHs,t}(i)}{P_{HFHs,t}} \right)^{-\sigma_s} di \right].$$

Likewise, let total labor in tradable sector $s$ be given by $L_{H,s,t} \equiv \int_0^1 L_{H,s,t}(i)di$. Using this expression and the production function for sector $s$, the corresponding goods market clearing condition is:

$$A_{H,s,t}L_{H,s,t} = C_{HHs,t} \int_0^1 \left( \frac{P_{HHs,t}(i)}{P_{HHs,t}} \right)^{-\sigma_s} di + C_{HFHs,t} \int_0^1 \left( \frac{P_{HFHs,t}(i)}{P_{HFHs,t}} \right)^{-\sigma_s} di.$$

Finally, market clearing in the aggregate labor market requires:

$$L_{H,t} = \sum_{s=N,D} L_{H,s,t} + L_{H,t}^z.$$
An analogous set of goods and labor market clearing conditions hold for the Foreign country.

C Additional experiments

The analysis on the relative importance of nominal rigidities and variable markups is mainly based on the comparison of impulse responses between the baseline model (benchmark calibration) and two alternative models. This appendix shows that the key conclusions are robust to the consideration of further alternative models, which feature different parameterizations. I first consider two scenarios that assume the same (non-zero) distribution margin in both sectors, implying that the variability of markups is not as different across sectors as in the baseline model. I also compute another model with a unique frequency of price adjustment in both sectors.

The first of these additional models sets the target for the distribution margin to $x_N = x_D = 0.49$, the value for the differentiated goods sector employed in the benchmark calibration. The second model uses $x_N = x_D = 0.42$, the value for the distribution margin in the nondifferentiated goods sector employed in the benchmark calibration. The third model assumes $\theta_N = \theta_D = 0.6$, the value for the Calvo parameter in the nondifferentiated goods sector employed in the benchmark calibration. Summary measures for the impulse responses based on these models are displayed in columns 5, 6, and 7 of Table A2, respectively. To facilitate comparison with previous findings, the first four columns of this Table replicate Table 5.

[Table A2]

In columns 5 and 6, we observe that regardless of the specific values assumed for the distribution margins, the responses of our variables of interest (i.e., relative prices and quantities) to a monetary shock are similar. Moreover, the magnitudes of the responses

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30 The first model implies the following values for the sectoral inputs of distribution services and the sectoral markup elasticities: $\kappa_N = 2.14$, $\kappa_D = 5.67$, $\Gamma_N = 0.78$, and $\Gamma_D = 2.88$. The second model implies: $\kappa_N = 1.32$, $\kappa_D = 2.41$, $\Gamma_N = 0.48$, and $\Gamma_D = 1.22$. 

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under these two models are only marginally smaller (in absolute value) relative to the baseline model (column 2). These results are in line with the conclusion that the variable markup channel plays a limited quantitative role.

The responses observed under a homogeneous degree of nominal rigidity across sectors show that with a low Calvo parameter (column 7), the model predicts larger declines in relative prices and quantities than with a high Calvo parameter (column 4). That said, the responses in column 7 are appreciably subdued relative to the baseline model. This finding confirms that the differences in the sectoral frequency of price changes constitute the central mechanism behind the main quantitative results.
### Table A1
Intermediate results for calculation of sectoral import shares (%)

<table>
<thead>
<tr>
<th>Major type of PCE product (nondurables only)</th>
<th>Share of PCE (1)</th>
<th>Import share (2)</th>
<th>Nondifferentiated</th>
<th>Sectoral composition</th>
<th>Differentiated</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td></td>
<td></td>
<td>Share of total purchasers' value (3)</td>
<td>Share of PCE (4)</td>
<td>Share of total nondiff. cons. (5)</td>
</tr>
<tr>
<td>Food and beverages purchased for off-premises consumption</td>
<td>7.8</td>
<td>6.1</td>
<td>57.3</td>
<td>4.5</td>
<td>48.2</td>
</tr>
<tr>
<td>Clothing and footwear</td>
<td>3.2</td>
<td>31.9</td>
<td>0.3</td>
<td>0.0</td>
<td>0.1</td>
</tr>
<tr>
<td>Gasoline and other energy goods</td>
<td>3.0</td>
<td>4.1</td>
<td>99.9</td>
<td>3.0</td>
<td>32.3</td>
</tr>
<tr>
<td>Other nondurable goods</td>
<td>8.0</td>
<td>18.6</td>
<td>22.5</td>
<td>1.8</td>
<td>19.4</td>
</tr>
</tbody>
</table>

**MEMO:**

Share of nondurable PCE by sector:
- Nondiff. (sum of column (4)) = 9.3
- Diff. (sum of column (7)) = 11.8

**Notes:** All numbers are percentages. The values in columns (1) and (2) are taken from Table 1 (“shares of PCE by major type of product”) and Table 4 (“import shares of PCE by major type of product”) of McCulley (2011). The numbers in columns (3) and (6) are the shares of total purchasers’ value of detail-level PCE categories within each major type of product that I classify either as nondifferentiated or differentiated (source: BEA’s bridge tables that map PCE categories into final use categories in the 2007 input-output accounts at the detail level). Column (4) is the product of columns (1) and (3). Column (7) is the product of columns (1) and (6). Column (5) results from dividing the values in column (4) by the overall nondifferentiated goods content of PCE, which is 9.3. Column (8) results from dividing the values in column (7) by the overall differentiated goods content of PCE, which is 11.8. The import content of nondifferentiated goods \((1 - \psi_N)\) is the sum product of columns (2) and (5), and the import content of differentiated goods \((1 - \psi_D)\) is the sum product of columns (2) and (8). See further details in the text.
Figure A1
Aggregate sectoral distribution margins, 1997–2012

Notes: The graph relies on the BEA’s bridge tables at the summary level that map PCE categories into final use categories in the I-O accounts. Based on my own Rauch classification of I-O commodities, the PCE categories that belong to the following major products are treated as nondifferentiated: ‘Food and beverages purchased for off-premises consumption’, ‘Gasoline and other energy goods’, and ‘Tobacco’; while the PCE categories that belong to the following major products are treated as differentiated: ‘Clothing and footwear’, ‘Pharmaceutical and other medical products’, ‘Household supplies’, ‘Personal care products’, and ‘Magazines, newspapers, and stationery’. The aggregate sectoral distribution margin is calculated as the sum of the wholesale and retail distribution margins and transportation costs, divided by the sum of all output valued at purchaser’s prices, across all final use categories in each sector’s I-O accounts. See further details in the text.
Table A2
Summary measures of theoretical and empirical impulse responses to contractionary nominal spending shock: relative prices and quantities, Home country (%)

<table>
<thead>
<tr>
<th></th>
<th></th>
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</thead>
<tbody>
<tr>
<td></td>
<td>(1)</td>
<td>(2)</td>
<td>(3)</td>
<td>(4)</td>
<td>(5)</td>
<td>(6)</td>
</tr>
<tr>
<td><strong>Maximum response with expected sign (1 year after shock)</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Rel. PPI</td>
<td>-0.27</td>
<td>-0.19</td>
<td>-0.25</td>
<td>-0.03</td>
<td>-0.17</td>
<td>-0.19</td>
</tr>
<tr>
<td>Rel. Production</td>
<td>-0.29</td>
<td>-0.32</td>
<td>-0.39</td>
<td>-0.04</td>
<td>-0.28</td>
<td>-0.31</td>
</tr>
<tr>
<td>Rel. Import price</td>
<td>-0.97</td>
<td>-0.19</td>
<td>-0.25</td>
<td>-0.03</td>
<td>-0.16</td>
<td>-0.18</td>
</tr>
<tr>
<td>Rel. Import volume</td>
<td>-0.38</td>
<td>-0.34</td>
<td>-0.43</td>
<td>-0.04</td>
<td>-0.30</td>
<td>-0.32</td>
</tr>
<tr>
<td><strong>Average response 1 year after shock</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Rel. PPI</td>
<td>-0.14</td>
<td>-0.16</td>
<td>-0.19</td>
<td>-0.02</td>
<td>-0.14</td>
<td>-0.15</td>
</tr>
<tr>
<td>Rel. Production</td>
<td>-0.15</td>
<td>-0.28</td>
<td>-0.32</td>
<td>-0.02</td>
<td>-0.25</td>
<td>-0.27</td>
</tr>
<tr>
<td>Rel. Import price</td>
<td>-0.32</td>
<td>-0.16</td>
<td>-0.19</td>
<td>-0.02</td>
<td>-0.14</td>
<td>-0.15</td>
</tr>
<tr>
<td>Rel. Import volume</td>
<td>0.08</td>
<td>-0.30</td>
<td>-0.36</td>
<td>-0.03</td>
<td>-0.26</td>
<td>-0.28</td>
</tr>
<tr>
<td><strong>Response 6 months after shock</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Rel. PPI</td>
<td>-0.14</td>
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<td>-0.14</td>
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<td>-0.28</td>
</tr>
</tbody>
</table>

Notes: See notes to Table 5. Columns 1-4 replicate the columns from Table 5. Columns 5 and 6 report responses for models that use the same target for the distribution margin in both sectors. Column 5 uses \(x_N = x_D = 0.49\), the target for the distribution margin in the differentiated goods sector employed in the benchmark calibration. Column 6 uses \(x_N = x_D = 0.42\), the target for the distribution margin in the nondifferentiated goods sector employed in the benchmark calibration. Column 7 assumes the same frequency of price adjustment for both sectors: \(\theta_N = \theta_D = 0.6\), the value for the Calvo parameter in the nondifferentiated goods sector employed in the benchmark calibration.