ABSTRACT

Title of dissertation: ESSAYS ON INTERNATIONAL MACROECONOMICS AND TRADE

Jorge M. Salas Maldonado
Doctor of Philosophy, 2015

Dissertation directed by: Professor John Shea
Department of Economics

This dissertation explores quantitative implications of heterogeneity in price stickiness and contractual vulnerability within an open economy context. Particular attention is given to the effects of these frictions on sectoral variables during economic downturns.

Chapter 1 documents that U.S. producer prices of differentiated goods barely moved in recent recessions, while the declines in industrial production were large. In contrast, for nondifferentiated goods, large producer price reductions were accompanied by relatively small adjustments in production. Similar patterns have been found in international trade data on prices and quantities. I use a two-country general equilibrium model with trade in nondifferentiated and differentiated 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 main 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 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.

Chapter 2, co-written with Renzo Castellares, starts from the observation that some products are more sensitive to imperfect contracting than others. Hence, industries exhibit different degrees of contractual vulnerability. We build a simple theory in which: (i) exporters are paid after delivery of the goods, and (ii) a complementarity exists between (procyclical) contract enforcement at the importing-country level and contractual vulnerability at the industry level. In this model, an adverse aggregate shock in an importing country generates a disproportional decline in imports in more contractually vulnerable industries. Using disaggregated bilateral trade data for many countries from 1989 to 2000, and exploiting the variation in contractual dependence across manufacturing industries, we find robust empirical support for the model's predictions. The estimated sectoral effects are statistically significant and economically important. Our analysis employs different industry measures of contractual vulnerability, including a novel indicator that reflects payment defaults among firms.

ESSAYS ON INTERNATIONAL MACROECONOMICS AND TRADE

By

Jorge M. Salas Maldonado

Dissertation submitted to the Faculty of the Graduate School of the
University of Maryland, College Park in partial fulfillment
of the requirements for the degree of
Doctor of Philosophy
2015

Advisory Committee:
Professor John Shea, Chair
Professor Sebnem Kalemli-Özcan
Professor Nuno Limão
Professor Luminita Stevens
Professor Phillip L. Swagel

Dedication

To my parents and my wife

Acknowledgments

This thesis would not have been written without the support of my advisors. I am particularly indebted to John Shea for his invaluable guidance and insightful comments. His faith in my research was instrumental for the completion of this dissertation. I am also grateful to Nuno Limão for his excellent advising and helpful suggestions. I owe my gratitude as well to Luminita Stevens and Sebnem Kalemli-Özcan for providing very useful comments on my work. Special thanks go to Phillip Swagel for agreeing to serve on my dissertation committee. Additionally, I thank Pablo D'Erasmo, Rafael Dix-Carneiro, Borağan Aruoba, John Haltiwanger, and Felipe Saffie for valuable suggestions.

My wife Claudia has given me love and encouragement throughout these years. I will always be grateful for her amazing patience. I am also fortunate to have the unconditional support of my loving parents and my sister. The rest of my family in Peru, my wife's parents and siblings, and my closest friends were additional sources of confidence and affection. I owe my deepest thanks to all of them.

My friend and colleague from the University of Virginia, Renzo Castellares, deserves a special mention for his work as a coauthor of the second chapter of this dissertation.

At Maryland, I was lucky to have excellent classmates and friends. I am especially grateful to Kristian López-Vargas, Miguel Sarzosa, Jerónimo Carballo, Pablo Cuba, John León, David Moreno, Lixin Tang, and Yiqun Chen.

During my doctoral studies I benefited from financial support from the Department of Economics and the Graduate School at the University of Maryland. I also acknowledge the support from the Central Bank of Peru in the Winter of 2013.

Finally, I thank God, Mother Mary, and my dear grandma Emilia for constantly making me feel their presence in my life.

Table of Contents

List of Tables vii

List of Figures ix

List of Abbreviations x

1 A Tale of Two Sectors: Product Differentiation and Heterogeneity in Price Stickiness in a General Equilibrium Model 1
1.1 Introduction 1
1.2 Related literature 5
1.3 Empirical evidence: sectoral prices and quantities 7
1.3.1 Domestic data: new facts 8
1.3.2 International trade data: a brief survey 12
1.3.3 Effects of monetary policy shocks: VAR evidence 13
1.4 Model 19
1.5 Quantitative results 32
1.5.1 Calibration 32
1.5.2 Responses to a monetary shock 38
1.5.3 The key mechanisms: nominal rigidities and variable markups 43
1.5.4 Model-based vs. empirical impulse response functions 48
1.6 Conclusions 51

2 Contractual Imperfections and the Impact of Crises on Trade 54
2.1 Introduction 54
2.2 A simple framework of trade and contractual imperfections 59
2.2.1 Setup 59
2.2.2 Main theoretical predictions 63
2.3 Empirical strategy 64
2.3.1 Methodology 64
2.3.2 Data 67
2.4 Results 77
2.4.1 Baseline results 78

2.4.2 Sensitivity analysis . . . 82
2.4.2.1 Contract enforcement (rule of law) at the country level 82
2.4.2.2 Extending the sample period . . . 84
2.4.2.3 Controlling for financial vulnerability . . . 86
2.4.2.4 Controlling for cyclicality . . . 89
2.4.2.5 Controlling for product differentiation . . . 91
2.4.2.6 Controlling for financial vulnerability, cyclicality, and product differentiation . . . 93
2.4.2.7 Time-variant multilateral trade resistance . . . 95
2.5 Conclusions . . . 97

A Appendix for Chapter 1 100
A.1 Methodological details . . . 100
A.2 Model: market clearing conditions . . . 108
A.3 Additional experiments . . . 109

B Appendix for Chapter 2 112
B.1 Data sources . . . 112
B.2 Data from the National Summary of Domestic Trade Receivables . . . 113
B.3 Sensitivity analysis: additional results . . . 115

List of Tables

1.1 Cumulated changes in industrial production indices and producer price indices during recessions and selected sub-periods, by sector . . 10
1.2 Benchmark calibration 33
1.3 Cumulated responses of sectoral and relative prices to contractionary nominal spending shock, under alternative model specifications 44
1.4 Summary measures of theoretical and empirical impulse responses to contractionary nominal spending shock 50

2.1 List of countries 69
2.2 Summary statistics: indicators of contractual and financial vulnerability, at industry level 76
2.3 Pairwise correlation coefficients: indicators of contractual and financial vulnerability 76
2.4 Summary statistics: trade and country-level data 77
2.5 Effects of crises and contractual vulnerability on trade across countries and industries 79
2.6 Amplification effects of contractual vulnerability on industry imports 81
2.7 Effects of crises and contract vulnerability on trade: contract enforcement at importing-country level 83
2.8 Effects of crises and contract vulnerability on trade: extended sample period 85
2.9 Effects of crises and contract vulnerability on trade: controlling for financial vulnerability 88
2.10 Effects of crises and contract vulnerability on trade: controlling for cyclicality 92
2.11 Effects of crises and contract vulnerability on trade: controlling for product differentiation 94
2.12 Effects of crises and contract vulnerability on trade: controlling for financial vulnerability, cyclicality, and differentiation 96
2.13 Effects of crises and contract vulnerability on trade: time-variant multilateral resistance 98

A.1 Classification of manufacturing industries by product differentiation at the 4-digit NAICS level . . . 101
A.2 Summary measures of theoretical and empirical impulse responses to contractionary nominal spending shock . . . 110

B.1 Data sources . . . 112
B.2 Uncollectible index: the ten highest and lowest ranked industries . . . 114
B.3 Effects of crises and contract vulnerability on trade: controlling for financial vulnerability . . . 115
B.4 Effects of crises and contract vulnerability on trade: controlling for cyclicality . . . 116

List of Figures

1.1 Industrial production index by sector: Percentage (log) changes in smoothed indices . . . 10
1.2 Producer price index by sector: Percentage (log) changes in smoothed indices . . . 12
1.3 Empirical impulse responses to contractionary monetary policy shock: VAR with domestic data . . . 15
1.4 Empirical impulse responses to contractionary monetary policy shock: VAR with import data . . . 18
1.5 Sensitivity of elasticities of markups to distribution margins . . . 38
1.6 Theoretical impulse responses to contractionary nominal spending shock: aggregate variables . . . 39
1.7 Theoretical impulse responses to contractionary nominal spending shock: sectoral domestic variables . . . 41
1.8 Theoretical impulse responses to contractionary nominal spending shock: sectoral trade variables . . . 42
1.9 Theoretical impulse responses to contractionary nominal spending shock: relative prices and quantities under alternative models . . . 47
1.10 Theoretical and empirical impulse responses to contractionary nominal spending shock: relative prices and quantities . . . 49

2.1 Share of observations with crises, by year . . . 71
2.2 Distribution of the uncollectible index . . . 75
2.3 Distribution of complexity index and concentration index, by cyclicality of industries . . . 90

A.1 Aggregate sectoral distribution margins, 1997–2012 . . . 107

List of Abbreviations

BLS	Bureau of Labor Statistics
CES	Constant Elasticity of Substitution
CRF	Credit Research Foundation
FTA	Free Trade Agreement
GDP	Gross Domestic Product
HS	Harmonized System
IFS	International Financial Statistics
I-O	Input-Output
IRF	Impulse-Response Function
LCP	Local Currency Pricing
NAICS	North American Industry Classification System
NBER	National Bureau of Economic Research
NESOI	Not Elsewhere Specified Or Included
NIPAs	National Income and Product Accounts
NSDTR	National Summary of Domestic Trade Receivables
OLS	Ordinary Least Squares
PCE	Personal Consumption Expenditure
PCP	Producer Currency Pricing
PPI	Producer Price Index
SIC	Standard Industrial Classification
SITC	Standard International Trade Classification
USITC	United States International Trade Commission
VAR	Vector Autoregression
WDI	World Development Indicators

Chapter 1: A Tale of Two Sectors: Product Differentiation and Heterogeneity in Price Stickiness in a General Equilibrium Model

1.1 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 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

[1] See 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.

(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.[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).[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,

[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.

[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).

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, producers reset prices 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.

1.2 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

[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

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 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.[5]

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

in an open economy context, Carvalho and Nechio (2011) who study the real exchange rate.

[5] 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.

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.[6] 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.

1.3 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.[7]

[6] 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.

[7] 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).

1.3.1 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 (SITC) level, into the NAICS classification of manufacturing industries. Details on the concordance method are provided in Appendix A.1. In the Rauch (1999) classification, I treat goods traded on organized exchanges or with reference prices as nondifferentiated. For illustrative purposes, Appendix Table A.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.

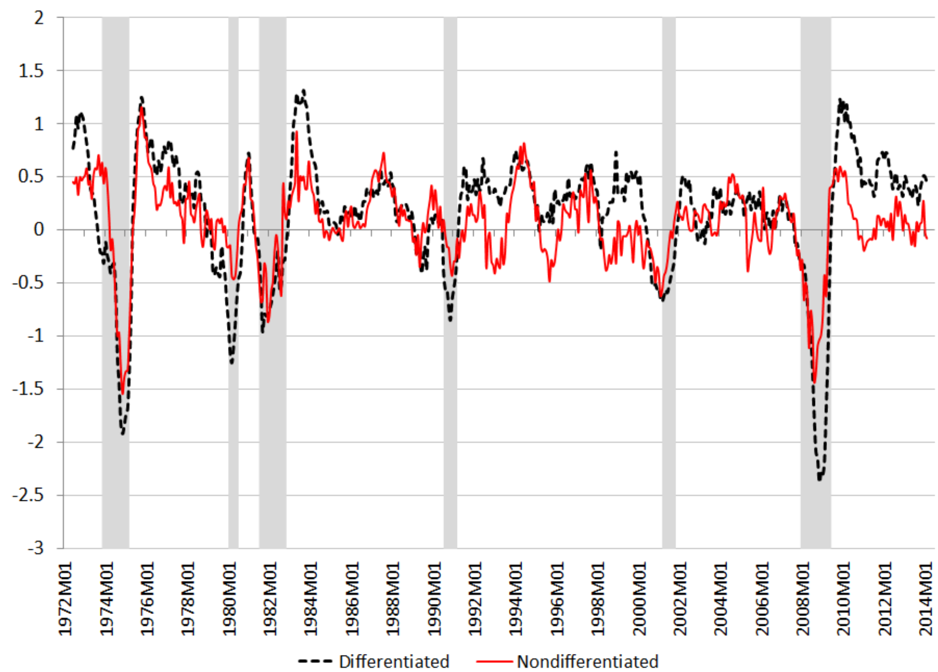
Aggregate production of nondifferentiated and differentiated goods is measured as the weighted average of (seasonally-adjusted) industrial production indices corresponding to 4-digit NAICS manufacturing industries classified as nondifferentiated and differentiated, respectively. The weights are calculated 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 1.1 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 1.1 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% (differentiated goods) and −13% (nondifferentiated goods).

To analyze the behavior of prices in both sectors, I show the percentage changes of the (smoothed) sectoral PPIs for the period 1976:1–2014:6. As observed in Figure

Figure 1.1
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 IP indices, smoothed using 9-month centered moving averages, multiplied by 100. The shaded areas indicate NBER recession dates.

Table 1.1
Cumulated changes in industrial production indices and producer price indices during recessions and selected sub-periods, by sector (%)

	1973–75	1980	1981–82	1990–91	2001	2007–09	Aug.08–Jun.09
Production							
Nondiff.	-11.5	-5.2	-5.5	-3.2	-1.2	-13.4	-8.2
Diff.	-20.1	-10.5	-11.9	-7.7	-4.7	-27.2	-22.4
Prices							
Nondiff.	n/a	3.1	-5.0	-3.8	-2.7	-1.9	-15.0
Diff.	n/a	6.9	4.0	1.8	0.7	6.9	0.8

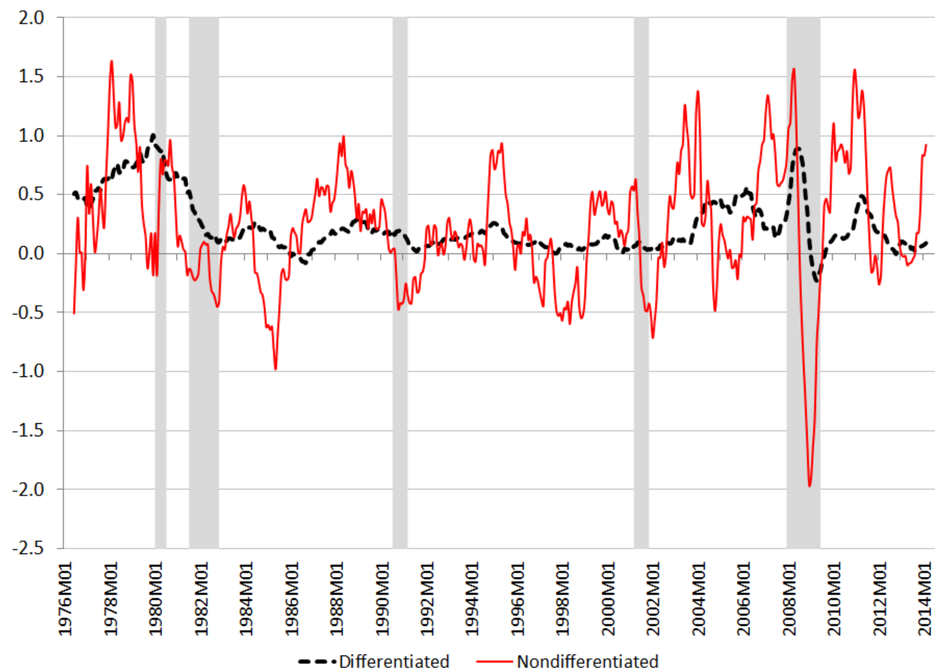
Notes: The table reports the percentage (log) changes cumulated during NBER recession dates and the acute phase of the 2007–2009 recession (August 2008 to June 2009). The underlying monthly data for production corresponds to the sectoral industrial production indices for nondifferentiated and differentiated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

1.2, 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 exhibit higher price flexibility than other categories of more differentiated manufacturers (Nakamura and Steinsson, 2008). The last two rows of Table 1.1 show that 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%). On the other hand, differentiated goods prices either increased or remained fairly stable in all of the recessions included in the sample period.[8]

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 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

[8] I 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.

Figure 1.2
Producer price index (PPI) by sector: Percentage (log) changes in smoothed indices, 1976:1–2014:6



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.

fact, a flatter supply curve for differentiated goods than for nondifferentiated goods appears to be a key element to explain the empirical findings.

1.3.2 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 (HS) level for a group of countries including the U.S.[9]

1.3.3 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

[9] 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.

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.

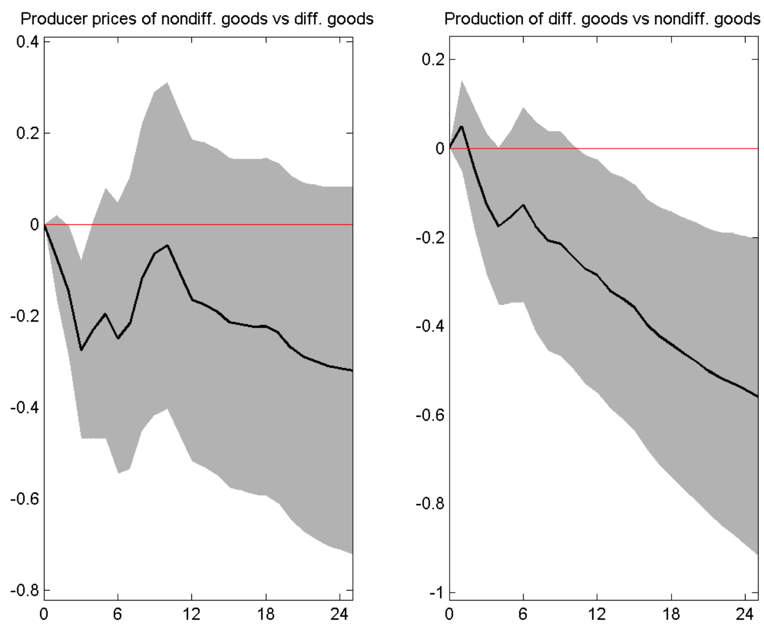
I estimate VARs on monthly U.S. data to generate impulse response functions (IRFs) to contractionary monetary policy shocks. The identification method is based on the standard recursiveness assumption developed in Christiano et al. (1999)—the central bank observes the current and lagged values of all 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).[10]

[10] 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

As shown in Figure 1.3, 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% 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%.[11]

Figure 1.3
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.

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.

[11]I also estimated two additional sectoral VARs, each including sector-specific measures of PPI and industrial production, and found no evidence of price puzzles, as desired.

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.

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.[12] 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 1.3.2. 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.

[12]Outliers are excluded by eliminating changes of a magnitude greater than 2 log points in the disaggregated trade volume series.

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.[13] 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 order is 9.[14]

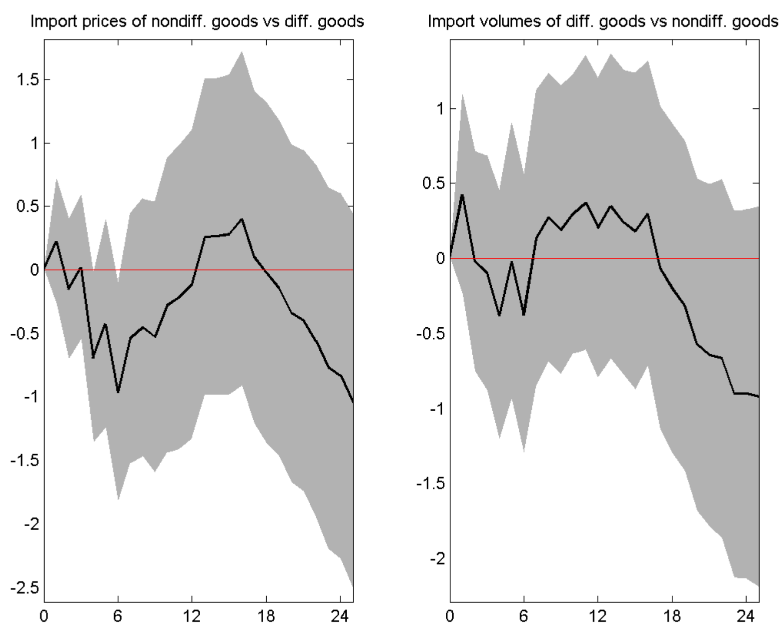
In Figure 1.4, the left panel shows that the import prices of nondifferentiated

[13] 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).

[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.

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% 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%.

Figure 1.4
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.

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.

1.4 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 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:

$$\mathsf{E}_0 \sum_{t=0}^{\infty} \beta^t \left[\frac{\overline{C}_{H,t}^{1-\gamma}}{1-\gamma} - \frac{L_{H,t}^{1+\phi}}{1+\phi} \right],$$

where $\overline{C}_{H,t}$ is an overall consumption index, $L_{H,t}$ denotes labor supply, E_0 is the time-0 expectations operator, β is the discount factor, γ is the inverse of the intertemporal elasticity of substitution, and ϕ 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 $C^z_{H,t}$, respectively:

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

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 $C^z_{H,t}$ is given by an aggregator of varieties $i \in [0,1]$ with elasticity of substitution $\xi > 1$: $C^z_{H,t} = \left(\int_0^1 C^z_{H,t}(i)^{(\xi-1)/\xi} di \right)^{\xi/(\xi-1)}$. 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:

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

$$C_{kHs,t} = \left(\int_0^1 C_{kHs,t}(i)^{(\sigma_s-1)/\sigma_s} di\right)^{\sigma_s/(\sigma_s-1)},$$

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.[15] Note that the notation for $C_{HHs,t}$ and $C_{FHs,t}$ is such that the first subscript denotes the country where the goods are produced, the second subscript denotes the country where the goods are sold, and the third subscript denotes the sector.

Consumer optimization yields the following set of demands:

$$C_{H,t} = \chi \left(\frac{P_{H,t}}{\overline{P}_{H,t}}\right)^{-\omega} \overline{C}_{H,t}, \qquad C^z_{H,t} = (1-\chi) \left(\frac{P^z_{H,t}}{\overline{P}_{H,t}}\right)^{-\omega} \overline{C}_{H,t},$$

$$C_{Hs,t} = \mu_s \left(\frac{P_{Hs,t}}{P_{H,t}}\right)^{-\eta} C_{H,t}, \qquad C_{HHs,t} = \psi_s \left(\frac{P_{HHs,t}}{P_{Hs,t}}\right)^{-\sigma_s} C_{Hs,t},$$

$$C_{FHs,t} = (1-\psi_s) \left(\frac{P_{FHs,t}}{P_{Hs,t}}\right)^{-\sigma_s} C_{Hs,t}, \qquad C_{kHs,t}(i) = \left(\frac{P_{kHs,t}(i)}{P_{kHs,t}}\right)^{-\sigma_s} C_{kHs,t},$$

$$C^z_{H,t}(i) = \left(\frac{P^z_{H,t}(i)}{P^z_{H,t}}\right)^{-\xi} C^z_{H,t},$$

where $\overline{P}_{H,t}$ is the overall consumption price index, $P_{H,t}$ is the price index of tradable

[15] 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.

goods, $P^z_{H,t}$ is the price index of nontradable goods, $P_{Hs,t}$ is the price index of tradable goods from sector s, $P_{kHs,t}$ 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 $P^z_{H,t}(i)$ is the price of variety i of nontradable goods. The CES price indices are given by:

$$\overline{P}_{H,t} = \left[\chi \left(P_{H,t}\right)^{1-\omega} + (1-\chi)\left(P^z_{H,t}\right)^{1-\omega}\right]^{1/(1-\omega)}, \qquad P^z_{H,t} = \left(\int_0^1 P^z_{H,t}(i)^{1-\xi} di\right)^{1/(1-\xi)},$$

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

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

The household's flow budget constraint is:

$$\overline{P}_{H,t}\overline{C}_{H,t} + \mathsf{E}_t \Theta_{t,t+1} B_{H,t+1} \leq W_{H,t} L_{H,t} + B_{H,t} + \Pi_{H,t},$$

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:

$$\overline{C}^{\gamma}_{H,t} L^{\phi}_{H,t} = \frac{W_{H,t}}{\overline{P}_{H,t}},$$

and an international risk-sharing condition:

$$\frac{E_{HF,t}\overline{P}_{F,t}}{\overline{P}_{H,t}} = \left(\frac{\overline{C}_{H,t}}{\overline{C}_{F,t}}\right)^{\gamma},$$

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 $\overline{C}_{F,t}$ and $\overline{P}_{F,t}$ 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.[16] Firms in the distribution sector are perfectly competitive. They combine varieties of distribution services $D^z_{H,t}(i)$ supplied by firms in the nontradable sector to produce a CES composite defined by $D^z_{H,t} = \left(\int_0^1 D^z_{H,t}(i)^{(\xi-1)/\xi} di\right)^{\xi/(\xi-1)}$. For simplicity, no distinction is made between nontradable consumption goods and distribution services. Hence, the price of the composite $D^z_{H,t}$ is given by the price index of nontradable goods $P^z_{H,t}$, and the optimal demand from the distribution sector for a variety of nontradables is given by $D^z_{H,t}(i) = \left(\frac{P^z_{H,t}(i)}{P^z_{H,t}}\right)^{-\xi} D^z_{H,t}$.

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

[16]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.

domestically is given by:

$$P_{HHs,t}(i) = P^P_{HHs,t}(i) + \kappa_s P^z_{H,t}, \tag{1.1}$$

where $P^P_{HHs,t}(i)$ 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 $P^z_{H,t}$ 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:

$$P_{HFs,t}(i) = \frac{P^P_{HFs,t}(i)}{E_{HF,t}} + \kappa_s P^z_{F,t}, \tag{1.2}$$

where $P^P_{HFs,t}(i)$ denotes the producer price set by a Home firm i in units of the Home currency, and $P^z_{F,t}$ 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:

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

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

[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, κ_s, is assumed to be constant across countries.

[18] As shown below, due to the local distribution costs, $P^P_{HHs,t}(i)$ is different from $P^P_{HFs,t}(i)$. In other words, a given producer discriminates between markets.

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 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:

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

Note that this maximization problem is subject to the relationship between producer and consumer prices shown in equations (1.1) and (1.2). By symmetry, the *optimal sectoral producer price under flexible prices*, denoted by $\widetilde{P}^P_{Hks,t}$, is common to all firms i in a given sector s, and can be written as an optimal markup over the marginal cost:

$$\widetilde{P}^P_{Hks,t} = \begin{cases} \zeta_{HHs,t} \frac{W_{H,t}}{A_{Hs,t}} & \text{, if } k = H \\ \zeta_{HFs,t} \frac{W_{H,t}}{A_{Hs,t}} & \text{, if } k = F \end{cases},$$

where $\zeta_{HHs,t}$ and $\zeta_{HFs,t}$ are the sector- and destination country-specific optimal markups, given by:

$$\zeta_{HHs,t} \equiv \frac{\sigma_s}{\sigma_s - 1}\left[1 + \frac{1}{\sigma_s}\frac{A_{Hs,t}}{W_{H,t}}\kappa_s P^z_{H,t}\right], \qquad \zeta_{HFs,t} \equiv \frac{\sigma_s}{\sigma_s - 1}\left[1 + \frac{1}{\sigma_s}\frac{A_{Hs,t}}{W_{H,t}}\kappa_s E_{HF,t} P^z_{F,t}\right].$$

The familiar result that a higher elasticity of substitution σ_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 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:

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

where $x_{HFs,t}(i) \equiv \frac{\kappa_s P^z_{F,t}}{\left(P^P_{HFs,t}(i)/E_{HF,t}\right)+\kappa_s P^z_{F,t}}$ is the sectoral share of the distribution cost

[19]Analogous maximization problems for the Foreign firm yield the following expressions for the optimal producer prices in units of the producer's currency: $\widetilde{P}^P_{FFs,t} = \zeta_{FFs,t}\frac{W_{F,t}}{A_{Fs,t}}$, and $\frac{\widetilde{P}^P_{FHs,t}}{E_{HF,t}} = \zeta_{FHs,t}\frac{W_{F,t}}{A_{Fs,t}}$, where the optimal markups $\zeta_{FFs,t}$ and $\zeta_{FHs,t}$ are given by: $\zeta_{FFs,t} \equiv \frac{\sigma_s}{\sigma_s-1}\left[1 + \frac{1}{\sigma_s}\frac{A_{Fs,t}}{W_{F,t}}\kappa_s P^z_{F,t}\right]$, and $\zeta_{FHs,t} \equiv \frac{\sigma_s}{\sigma_s-1}\left[1 + \frac{1}{\sigma_s}\frac{A_{Fs,t}}{W_{F,t}}\kappa_s \frac{P^z_{H,t}}{E_{HF,t}}\right]$.

in the retail price $P_{HFs,t}(i)$—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., $\zeta_{HFs,t} = \frac{\epsilon_{HFs}(i)}{\epsilon_{HFs}(i)-1}$), 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 $\zeta_{HFs,t}$ with respect to the relative price $\frac{P^P_{HFs,t}(i)}{E_{HF,t}P^z_{F,t}}$ is:

$$\Gamma_{HFs,t}(i) \equiv -\frac{\partial \log \zeta_{HFs,t}}{\partial \log \left(\frac{P^P_{HFs,t}(i)}{E_{HF,t}P^z_{F,t}}\right)} = \left[(\sigma_s - 1)\frac{1 - x_{HFs,t}(i)}{x_{HFs,t}(i)} - 1\right]^{-1}. \qquad (1.3)$$

The elasticity of the markup $\Gamma_{HFs,t}(i)$ is greater (in absolute value) in a sector s characterized by relatively high distribution margin $x_{HFs,t}(i)$ and low elasticity of substitution σ_s. More specifically, equation (1.3) implies that $\Gamma_{HFs,t}(i) = 0$ if $x_{HFs,t}(i) = 0$. Below I show that for realistic values of $x_{HFs,t}(i)$ and σ_s, the case in which $x_{HFs,t}(i) > 0$ implies $\Gamma_{HFs,t}(i) > 0$. 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 ($x_{HFs,t}$), the price elasticity of demand (ϵ_{HFs}), and the elasticity of the markup ($\Gamma_{HFs,t}$).[20]

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

[20]The producer's elasticity of the markup for *domestic* sales is given by $\Gamma_{HHs,t} = \left[(\sigma_s - 1)\frac{1-x_{HHs,t}}{x_{HHs,t}} - 1\right]^{-1}$, where $x_{HHs,t} \equiv \frac{\kappa_s P^z_{H,t}}{P^P_{HHs,t}(i)+\kappa_s P^z_{H,t}}$ is the sectoral distribution margin for final consumption of domestically produced goods sold in the Home country.

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:

$$\widetilde{p}^{P}_{Hks,t} = (1-\beta\theta_s)\sum_{j=0}^{\infty}(\beta\theta_s)^j\,\mathsf{E}_t\widetilde{p}^{P}_{Hks,t+j}, \tag{1.4}$$

where lower-case letters denote log-deviations from the steady state, so $\widetilde{p}^{P}_{Hks,t}$ 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 $\widetilde{p}^{P}_{Hks,t}$:

$$\widetilde{p}^{P}_{Hks,t+j} = \begin{cases} \frac{1}{1+\Gamma_{HHs}}\left(w_{H,t+j}-a_{Hs,t+j}\right)+\frac{\Gamma_{HHs}}{1+\Gamma_{HHs}}p^{z}_{H,t+j} & \text{, if } k=H \\ \frac{1}{1+\Gamma_{HFs}}\left(w_{H,t+j}-a_{Hs,t+j}\right)+\frac{\Gamma_{HFs}}{1+\Gamma_{HFs}}\left(p^{z}_{F,t+j}+e_{HF,t+j}\right) & \text{, if } k=F \end{cases}. \tag{1.5}$$

For either $k=H$ or $k=F$, the first terms in (1.5) indicate that with variable markups ($\Gamma_{Hks}>0$), movements in the marginal cost are incompletely passed through into $\widetilde{p}^{P}_{Hks,t}$. The pass-through rate is given by $\frac{1}{1+\Gamma_{Hks}}$, which declines with the *steady-state* sectoral elasticity of the markup Γ_{Hks}. Equations (1.4) and (1.5) 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 (1.5) reflect the dependence of optimal markups on local distribution costs, which are determined by the price of nontradables in the destination country $p^{z}_{k,t}$. For the price $\widetilde{p}^{P}_{HFs,t}$ 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 passed through into the optimal producer prices with a coefficient $\frac{\Gamma_{Hks}}{1+\Gamma_{Hks}}$, which increases with the sensitivity of the markup Γ_{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 ($\widetilde{p}^{P}_{FHs,t+j}$) is denominated in units of the Home currency due to the LCP assumption for Home imports. This price is given by:

$$\widetilde{p}^{P}_{FHs,t+j} = \frac{1}{1+\Gamma_{FHs}} \left(w_{F,t+j} + e_{HF,t+j} - a_{Fs,t+j}\right) + \frac{\Gamma_{FHs}}{1+\Gamma_{FHs}} p^{z}_{H,t+j}, \tag{1.6}$$

where the presence of the exchange rate $e_{HF,t+j}$ in the first term of equation (1.6) 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 $\widetilde{p}^{P}_{FHs,t+j}$ is incomplete, and the degree of incompleteness is larger for a sector with higher markup elasticity Γ_{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, $Y^z_{H,t}(i) = L^z_{H,t}(i)$, and face Calvo price rigidities with probability of non-adjustment denoted by θ^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:

$$\widehat{p}^z_{H,t} = (1 - \beta\theta^z) \sum_{j=0}^{\infty} (\beta\theta^z)^j \, \mathsf{E}_t \widetilde{p}^z_{H,t+j},$$

where the optimal price in the absence of nominal rigidity is $\widetilde{p}^z_{H,t+j} = w_{H,t+j}$.

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: $\overline{P}_{H,t}\overline{C}_{H,t} = M_{H,t}$, 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 A.2. 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 $P^P_{HFs,t}(i)$. Under the assumptions of symmetric firms and time-dependent pricing, $EPI_{Hs,t}$ is then given by:

$$\begin{aligned} EPI_{Hs,t} &\equiv \left(\int_0^1 P^P_{HFs,t}(i)^{1-\sigma_s} di\right)^{1/(1-\sigma_s)}. \\ &= \left[\theta_s \left(EPI_{Hs,t-1}\right)^{1-\sigma_s} + (1-\theta_s)\left(\widehat{P}^P_{HFs,t}\right)^{1-\sigma_s}\right]^{1/(1-\sigma_s)}, \end{aligned} \tag{1.7}$$

where $\widehat{P}^P_{HFs,t}$ 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:

$$\begin{aligned} IPI_{Hs,t} &\equiv \left(\int_0^1 P^P_{FHs,t}(i)^{1-\sigma_s} di\right)^{1/(1-\sigma_s)} \\ &= \left[\theta_s \left(IPI_{Hs,t-1}\right)^{1-\sigma_s} + (1-\theta_s)\left(\widehat{P}^P_{FHs,t}\right)^{1-\sigma_s}\right]^{1/(1-\sigma_s)}, \end{aligned} \tag{1.8}$$

where $\widehat{P}^P_{FHs,t}$ denotes the level of the optimal reset export price chosen by a Foreign firm. As observed in equations (1.7) and (1.8), 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:

$$PPI_{Hs,t} = (DPI_{Hs,t})^{\psi_s} (EPI_{Hs,t})^{1-\psi_s}, \tag{1.9}$$

where ψ_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:

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

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

1.5 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 1.3.3.

1.5.1 Calibration

The model is calibrated to monthly data. Table 1.2 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%, 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 ω is set to 0.74, as estimated in Mendoza (1991) based on data from industrialized countries.

Table 1.2
Benchmark calibration

Parameter	Value	Description/Source/Target
β	0.997	Discount factor.
γ	1	CRRA.
ϕ	0.40	Inverse Frisch elasticity.
χ	0.53	Share of tradable goods in consumption. Source: Corsetti et al. (2008).
ω	0.74	Elast. of subst. between tradables and nontradables. Source: Mendoza (1991).
η	1.50	Elast. of subst. between diff. and nondiff. goods.
μ_D	0.54	Share of diff. goods in tradable consumption. Source: NIPAs.
ψ_D	0.81	Domestic share in diff. goods consumption. Source: NIPAs and McCulley (2011).
ψ_N	0.92	Domestic share in nondiff. goods consumption. Source: NIPAs and McCulley (2011).
σ_D	2.3	Elast. of subst. across diff. varieties. Source: Broda and Weinstein (2006).
σ_N	3.2	Elast. of subst. across nondiff. varieties. Source: Broda and Weinstein (2006).
κ_D	5.67	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).
κ_N	1.32	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).
θ_D	0.93	Frequency of price non-adjustment, diff. goods. Source: Gopinath and Rigobon (2008) and Nakamura and Steinsson (2008).
θ_N	0.60	Frequency of price non-adjustment, nondiff. goods. Source: Gopinath and Rigobon (2008) and Nakamura and Steinsson (2008).
θ^z	0.94	Frequency of price non-adjustment, nontradable goods. Source: Nakamura and Steinsson (2008).
ρ_{MH}, ρ_{MF}	0.97	Persistence monetary shock. Source: M2 data.
σ_{MH}, σ_{MF}	0.003	Std. dev. monetary shock. Source: M2 data.
$\rho_{MH,MF}$	0.5	Cross-country correlation between monetary shocks.

Note: See section 1.5.1 for further details.

For both countries, the persistence of the money supply shock is set to $\rho_M = 0.97$, and the standard deviation to $\sigma_M = 0.33\%$. 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 σ_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.[21]

[21]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 calculate θ_N.

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.[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.1. 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 (μ_N and μ_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

[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.

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 ψ_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 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 ψ_N and ψ_D of 0.92 and 0.81, respectively.[23]

To compute the sectoral distribution margins ($x_s = \frac{\kappa_s P^z}{P_s}$) 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 σ_N and σ_D specified above imply sectoral inputs of distribution services

[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.1, 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 $x_D = 0.49$.

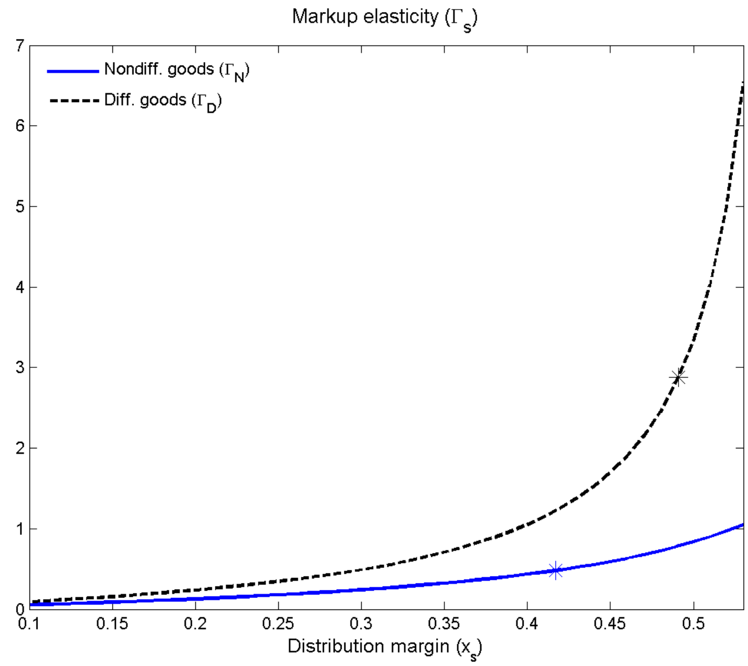
of $\kappa_N = 1.32$ and $\kappa_D = 5.67$.

Figure 1.5 shows the sector-s steady-state elasticity of the markup, denoted by Γ_s, as a function of steady-state values for the distribution margins x_s. Following equation (1.3), Γ_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 σ_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: $\Gamma_N < \Gamma_D$. Since the evidence suggests that distribution margins above 60% are rare (Burstein et al., 2003; Goldberg and Campa, 2010), the figure indicates that for a plausible calibration, the model may only predict potentially high values for Γ_D but not for Γ_N.

As indicated by asterisks in the graph, the baseline values for Γ_N and Γ_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 $1 - \frac{\Gamma_s}{1+\Gamma_s}$, eliminate the effect of nominal rigidities.

The calibration of the model is completed as follows. The frequency of price non-adjustment for nontradable goods θ^z is set equal to 0.94, based on the median

Figure 1.5
Sensitivity of elasticities of markups to distribution margins



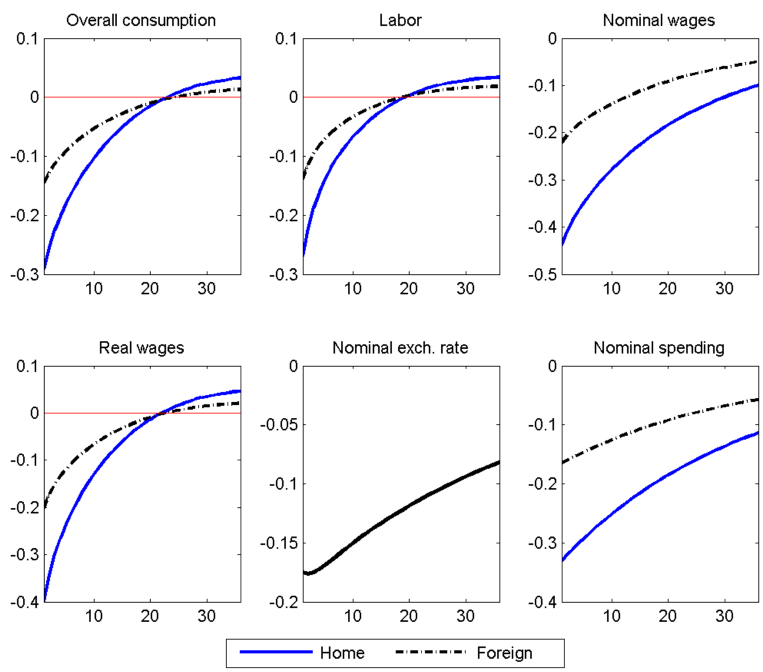
Notes: The depicted values assume $\sigma_N = 3.2$ and $\sigma_D = 2.3$. The asterisk on each curve indicates the parameter value implied by the benchmark calibration targets for the distribution margins x_N and x_D.

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 $\eta = 1.5$.

1.5.2 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 1.6, which depicts impulse responses of aggregate variables.

Figure 1.6
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.

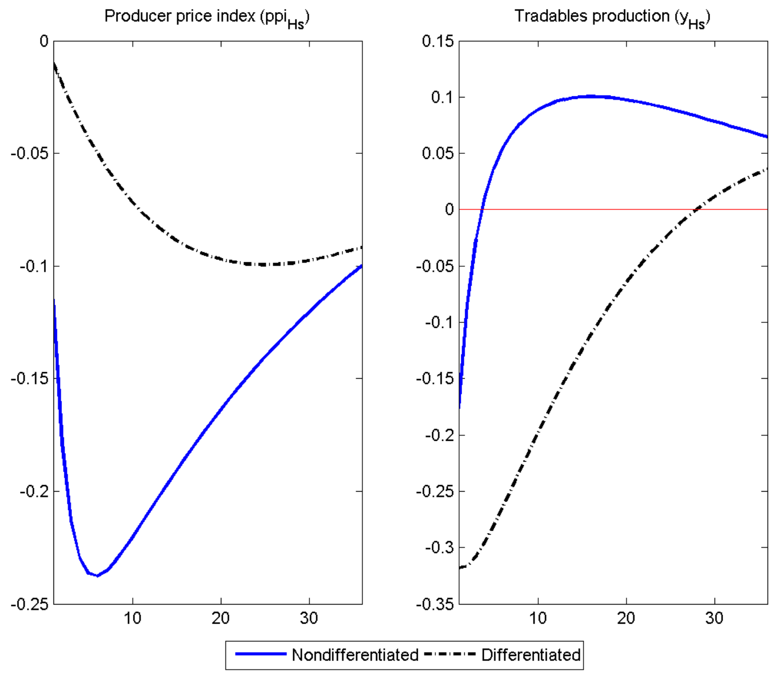
[25]Mussa (1986) is a classic reference for evidence that fluctuations in nominal and real exchange rates are driven by monetary shocks. For more recent evidence, see e.g. Bouakez and Normandin (2010). Using a VAR analysis, Eichenbaum and Evans (1995) find empirically that contractionary shocks to U.S. monetary policy lead to persistent and significant appreciations in U.S. nominal and real exchange rates.

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 1.7 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 (1.9)) 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.

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 more protracted than that of nondifferentiated goods (right panel in Figure 1.7). 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.

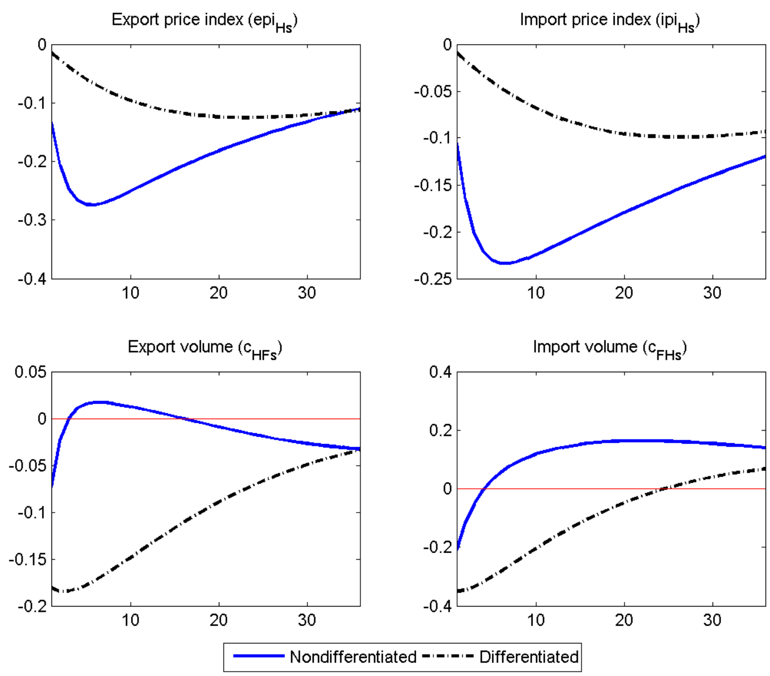
Figure 1.7
Theoretical impulse responses to contractionary nominal spending shock: sectoral domestic variables, Home country



Notes: The horizontal axis indicates the months after the shock. The variables are measured in percent.

The responses of sectoral trade variables for the Home country to a contractionary monetary shock are depicted in Figure 1.8. 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 1.8
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.

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.

1.5.3 The key mechanisms: nominal rigidities and variable markups

To examine the quantitative role of the key mechanisms of the model, Table 1.3 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 use 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 1.3 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 '*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,

Table 1.3

Cumulated responses of sectoral and relative prices to contractionary nominal spending shock (Home country), under alternative model specifications (%)

	Baseline model			Constant markups			Same Calvo parameter		
	Nondiff.	Diff.	Gap $N-D$	Nondiff.	Diff.	Gap $N-D$	Nondiff.	Diff.	Gap $N-D$
Optimal prices under flexible prices									
Domestic ($\widetilde{p}^{P}_{HHs}$)	-6.4	-4.6	-1.8	-7.9	-7.9	0.0	-6.8	-4.9	-2.0
Export ($\widetilde{p}^{P}_{HFs}$)	-7.3	-6.7	-0.7	-7.9	-7.9	0.0	-7.6	-6.6	-1.0
Import ($\widetilde{p}^{P}_{FHs}$)	-6.8	-4.7	-2.1	-8.8	-8.8	0.0	-6.8	-4.9	-2.0
Optimal reset prices									
Domestic ($\widehat{p}^{P}_{HHs,t}$)	-6.1	-3.8	-2.3	-7.5	-5.3	-2.1	-4.8	-4.0	-0.8
Export ($\widehat{p}^{P}_{HFs,t}$)	-7.0	-5.2	-1.8	-7.5	-5.3	-2.1	-5.3	-5.2	-0.2
Import ($\widehat{p}^{P}_{FHs,t}$)	-6.6	-4.1	-2.5	-8.4	-6.4	-2.1	-5.2	-4.1	-1.1
Industry-level price indices									
Domestic (dpi_{Hs})	-6.0	-2.7	-3.3	-7.3	-4.0	-3.4	-3.5	-2.8	-0.7
Producer (ppi_{Hs})	-6.0	-2.9	-3.2	-7.3	-4.0	-3.4	-3.6	-3.0	-0.6
Export (epi_{Hs})	-6.8	-3.7	-3.1	-7.3	-4.0	-3.4	-3.9	-3.7	-0.2
Import (ipi_{Hs})	-6.4	-2.8	-3.6	-8.2	-4.6	-3.6	-3.7	-2.9	-0.8

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.)

variable markups generate price stickiness above and beyond the nominal rigidity channel. Table 1.3 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 differentiated goods decreases around 3% both in the baseline model and in the model with constant markups, but it only declines by 0.6% 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.[26]

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.

[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.

To illustrate this, observe that in the top part of Table 1.3, the cumulated declines in the relative optimal prices (under flexible prices) are roughly between 1% and 2% 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.

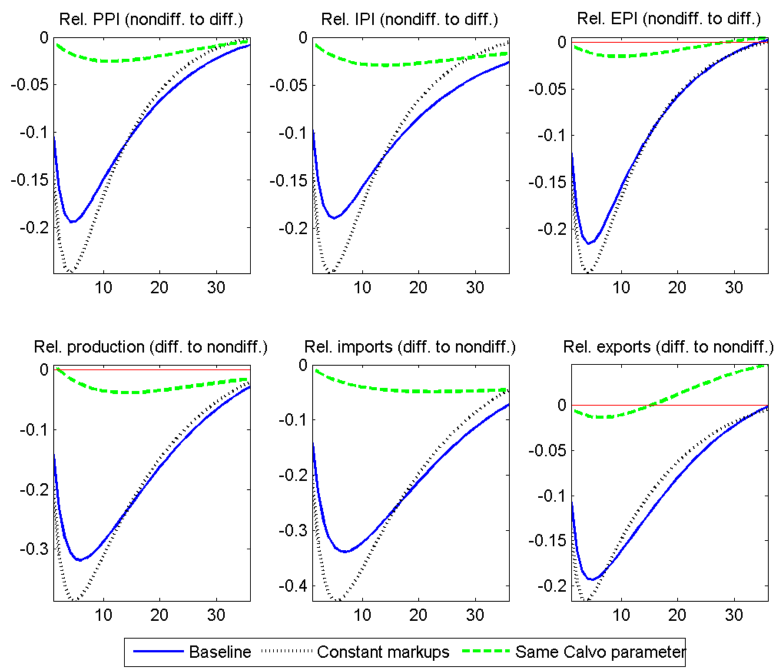
But as the results for the model with a single Calvo parameter indicate, the variable markup channel generates smaller movements in the relative prices between sectors if we focus instead on the optimal reset prices (middle part of Table 1.3) and the industry-level price indices (bottom part). This is 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).[27] 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 1.9 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 1.3, 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.

[27]The compounding effects of nominal rigidities arise because, as observed in equations (1.4), (1.7), and (1.8), the optimal reset prices and the industry-level price indices are functions of the degree of nominal rigidity measured by the Calvo parameter θ_s.

Figure 1.9 also shows that all of the models predict declines in the relative quantities produced and relative trade volumes of differentiated goods to nondifferentiated goods.

Figure 1.9
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.)

It is apparent from Figure 1.9 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.

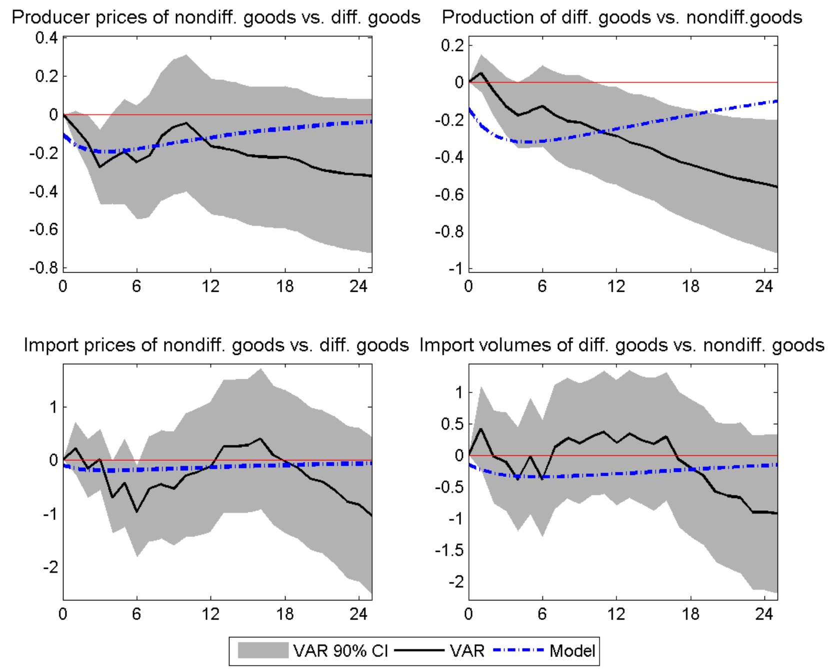
1.5.4 Model-based vs. empirical impulse response functions

Finally, I evaluate the model in terms of its ability to fit the VAR evidence. Figure 1.10 compares the empirical impulse responses of relative prices and quantities to a tightening monetary shock, as reported in section 1.3.3, 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 (such as habit formation, capital accumulation and adjustment costs).

To study the quantitative effects of a monetary shock in more detail, I compute three summary measures for the empirical and theoretical impulse responses: the

Figure 1.10
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 1.3 and 1.4. 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).

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. Table 1.4 displays the results. 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 A.3).

Table 1.4
Summary measures of theoretical and empirical impulse responses to contractionary nominal spending shock: relative prices and quantities, Home country (%)

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

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 1.3 and 1.4. 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: $\theta_N = \theta_D = 0.93$.

From Table 1.4, 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.

1.6 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.

Chapter 2: Contractual Imperfections and the Impact of Crises on Trade

Note: This chapter is coauthored with Renzo Castellares.

2.1 Introduction

Recent papers document the negative impact of crises on international trade. For example, Abiad et al. (2014) find empirically that financial crises are associated with significant declines in exports to the crisis country. In this paper, we argue that contractual imperfections are important to understand the causality between crises and trade disruptions. Our main finding is that exports to destinations in crisis are disproportionately affected in industries that are more contractually vulnerable. In this way, we provide empirical evidence on a new mechanism that has been thus far ignored in the literature on crises and trade.

We first propose a simple model of trade to explain the relevance of industries' contractual dependence during crises.[1] Our theory builds on the intuition that when international transactions are arranged in post-shipment terms (i.e., exporters are paid by importers after delivery of the goods), the risk of default of importers

[1]Throughout the paper we use the expressions "contractual vulnerability" and "contractual dependence" interchangeably.

matters (Schmidt-Eisenlohr, 2013). Importers are presumably less likely to honor their contracts when the state of their country's economy is weak—as would be the case if the economy were hit by a recession or if it entered into a financial crisis. But the probability of repayment under post-shipment terms can also be affected by industry-specific characteristics. In particular, when goods are more complex and/or customized, it is harder to verify their quality in court and the market value of the goods inside the original importer-exporter relationship is higher than outside this relationship.[2] Therefore exporters in some industries are more contractually vulnerable than in others. We then show that when an importing country suffers an adverse aggregate shock, a complementarity between contract enforcement at the country level and contract dependence at the industry level gives rise to a larger decline in imports in more contractually vulnerable industries. This is our key theoretical insight.

Using disaggregated bilateral trade data, we quantify the importance of contractual dependence at the industry level during crises. Our empirical approach exploits the variation in the occurrence of crises across 118 importing countries from 1989 to 2000, and the variation in contractual vulnerability across (up to) 351 SIC manufacturing industries. We simultaneously use three measures of crises in our regressions: recessions alone, financial crises alone, and recessions with financial crises. We confirm the negative average effects of crises on trade flows found in previous papers, but we also show that trade declines disproportionately in more

[2]Some important references in the literature on incomplete contracts include Williamson (1979), Williamson (1985), Grossman and Hart (1986), and Hart and Moore (1999). See Berkowitz et al. (2006) for an early study of the relationship between product complexity, contracting institutions, and trade.

contract-dependent industries. In most of our estimates, these sectoral effects are statistically significant and economically important. This finding is the central contribution of the paper.

We find that when crises involve recessions and financial disruptions at the same time, the amplification effects associated with contractual dependence at the industry level are stronger. According to one of our estimates, a recession with financial crisis is associated with a 8.6% larger drop in imports in an industry that is highly contract dependent relative to an industry that exhibits little dependence. To put this result in perspective, we find that the *average* impact of a recession with financial crisis on sectoral imports is close to −18%, while the analogous estimate in the case of a recession alone is nearly −6%.[3]

Our main empirical results are mostly robust to the following exercises: (i) extending the sample period back to 1980, (ii) controlling for industry measures of financial vulnerability, (iii) controlling for industry measures of cyclicality (or durability), (iv) controlling for industry measures of product differentiation, and (v) considering alternative specifications of the estimating equation. We also find evidence that in countries with lower institutional quality (proxied by the rule of law) the amplification effect of contractual vulnerability on sectoral imports is greater.

We use three industry measures of contractual vulnerability. Two of them are standard in the literature. The first one is the Nunn (2007) index of contract-intensity of goods, measured by the value share of inputs that Nunn identifies as

[3] Our definition of recessions is based on the methodology of Braun and Larrain (2005) and our definition of financial crises relies on Laeven and Valencia (2013).

relationship-specific. Levchenko (2007) provides us with a second indicator, which he constructs as an index of input-use concentration. Levchenko explicitly points out that his index represents a measure of product complexity; in our paper, as in Krishna and Levchenko (2012) and Hoefele et al. (2013), we make a similar assumption in terms of the Nunn index. We introduce an additional novel measure of contractual vulnerability, which we call the "uncollectible index". By quantifying the share of total account receivables uncollected compared to what was available to collect in a given period, the uncollectible index directly reflects payment defaults in business-to-business transactions. We obtain the data to construct this indicator from the National Summary of Domestic Trade Receivables, a proprietary quarterly survey of large U.S. firms. Our results are robust to the use of the Nunn, the Levchenko, and the uncollectible indices.

As summarized in Antràs (2015), several difficulties underlie the contractual imperfections associated with international transactions. First, it is sometimes difficult to determine which country's laws apply to a particular contract, especially since many contracts do not include a *choice-of-law* clause. Second, there is potential bias of courts in favor of their national citizens. Third, it is practically impossible in many cases to enforce decisions stipulated in a court's verdict. Recent coordinated attempts to reduce the contractual risk involving international transactions—notably, the Contracts for the International Sale of Goods initiative, and resorting to international arbitrators such as the International Chamber of Commerce—have fallen short of their objectives and constitute partial solutions at best. Quoting Rodrik (2000), Antràs (2015) concludes that ultimately international contracts remain incomplete.

This paper is related to the literature on the impact of financial crises and recessions on trade (Levchenko et al., 2010; Eaton et al., 2011; Berman et al., 2012; Bricongne et al., 2012; Chor and Manova, 2012; Bems et al., 2013; Abiad et al., 2014). A large part of this literature analyzes the so-called Great Trade Collapse of 2008–09. These papers have documented the role of several mechanisms, such as composition effects, protectionism, supply chains, credit constraints, and exchange rate dynamics. Our work contributes to this literature by emphasizing a new mechanism—contractual imperfections—that helps explain the important effects of crises on trade, and the heterogeneous impact across industries.

Our theoretical mechanism heavily relies on the role of default risk in trade. Other recent papers also study the implications of importers' repayment probability, but they mainly focus on a different problem, namely how this risk affects the choice of financing terms that support international trade (Hoefele et al., 2013; Schmidt-Eisenlohr, 2013; Ahn, 2014; Antràs and Foley, forthcoming).

Finally, this paper is connected to the literature on contracting institutions and trade (see Antràs, 2015 and Nunn and Trefler, 2014 for comprehensive reviews). A large bulk of this research has studied the relationship between domestic institutions and comparative advantage. Levchenko (2007) and Nunn (2007) constitute seminal contributions to that literature. Our use of the contractual-vulnerability indices introduced in those two papers to analyze the effects of crises on trade is new relative to previous work.

2.2 A simple framework of trade and contractual imperfections

To fix ideas, we propose a static, partial equilibrium model of trade. The model incorporates contractual frictions in a reduced-form way, which reflect contracting imperfections affecting the outputs produced by different industries. We then use the model to derive our main testable implications.[4]

2.2.1 Setup

Basic assumptions. Our framework is in line with the traditional monopolistic competition models of trade. In each country, a continuum of firms produce differentiated goods in multiple industries (sectors), indexed by s, using labor (supplied inelastically). A numeraire sector produces a freely-traded homogeneous good under constant returns to scale. Relative wages are pinned down by productivity in this numeraire sector. Preferences are identical across countries and are described by a Cobb-Douglas utility function. For country i, the utility function is $U_i = \prod_s C_{is}^{\mu_s}$, defined over CES consumption indices $C_{is} = \left(\int_{\Omega_{is}} x_{is}(\omega)^{(\sigma-1)/\sigma} d\omega\right)^{\sigma/(\sigma-1)}$, where ω is a variety, Ω_{is} is the set of available varieties, $\sigma > 1$ is the elasticity of substitution, and μ_s is the sectoral expenditure share.

Production technology in the differentiated sectors exhibits increasing returns to scale. A firm from sector s in country e that sell x_{eis} units of a good to an importer in country i faces the cost function $w_e \tau_{ei} x_{eis} + f_{ei}$, where w_e is the wage

[4] For simplicity the model is written in terms of final goods, but its key implications could be generalized for transactions involving intermediate inputs as well.

rate, $\tau_{ei} > 1$ is an iceberg trade cost, and f_{ei} is a fixed cost in units of the numeraire. Since local sales are immaterial for our purposes, we can simply assume that $f_{ii} = 0$.

Post-shipment payment. We assume that exporters are risk neutral and use open account contracts, meaning that they are paid by importers after delivery of the goods. (Importers can be thought as wholesalers who sell to domestic consumers.) Using trade data at the transaction level, Antràs and Foley (forthcoming) (U.S.) and Ahn (2014) (Chile and Colombia) show that in terms of payment methods, open account contracts comprise the majority of international transactions, both by number and by value. Asmundson et al. (2011) report a similar finding for worldwide trade based on survey data.[5]

Contractual frictions. Importers in country i are assumed to honor their contractual obligations (i.e., pay in full and on time to exporters) with probability λ_i. We assume that this probability increases with aggregate real expenditure in the importing country, Y_i. That is, λ_i is procyclical: $\lambda_i = \lambda(Y_i)$, with $\lambda'(Y_i) \equiv \frac{\partial \lambda_i}{\partial Y_i} > 0$.

A simple way to interpret this assumption is that in the wake of an adverse aggregate demand or financial shock in country i, some importing firms become insolvent or illiquid and are unable to pay in full and/or on time. In support of this argument, Mora and Powers (2011) document the increased perception of counterparty risk among international traders during the 2008 crisis, evidenced by the fact that exporters raised their demand for low-risk financing. Similarly, Auboin and Engemann (2014) use a comprehensive database of export credit insurance

[5] Payment under open account terms typically occurs between one and three months after the goods arrive to the importer's location.

covering 91 countries and find that the risk of international trade, as proxied by claims paid on insured open account contracts, steadily increased during the acute phase of the 2008 crisis. Additionally, Jacobson et al. (2013) use data on Swedish businesses and document that the output gap is a good predictor of firm insolvency.

In the context of models of trade financing terms, Schmidt-Eisenlohr (2013) and Antràs and Foley (forthcoming) propose a related setup in which λ_i represents instead a structural index of the quality of contracting institutions in country i. In one of our empirical exercises below, we take that modeling approach into account by dividing our sample of importing countries into two groups: countries with weak and strong rule of law.

We also assume that contract enforcement has an additional industry-specific dimension, captured by the index $z_s \in [0, 1]$. A higher value of z_s implies that the good s is more *contract dependent*, in the sense that it is more complex and hence more sensitive to imperfect contracting. Intuitively, complex goods require a high share of relationship-specific inputs and often involve customization. Moreover, the quality of a complex good can be difficult to verify in court. Importers of this type of goods are thus more likely to renege on the contract due to disagreement on the quality of the delivered products.[6] Further, due to their customized nature, complex goods may be hard to resell, so their market value outside the original importer-exporter relationship arguably declines following an importer's default.

As in Hoefele et al. (2013), we assume a complementarity between contract

[6]According to Burstein and Gopinath (2014), a typical trade credit insurance contract covers against defaults due to insolvency, but not due to disagreement. In their analysis of trade credit defaults among French firms, Burstein and Gopinath (2014) document that the most prevalent reason for defaulting on trade credit is disagreement, followed by illiquidity and insolvency.

enforcement at the importing-country level and contract dependence at the industry level. In particular, we assume that the probability of enforcement in country i and sector s is given by $\overline{\lambda_{is}} = \lambda_i^{z_s}$. For a given λ_i, higher values of z_s associated with more complex goods imply a lower effective probability of contract enforcement in country i. For the least contract dependent product, $z_s = 0$, the importer in country i honors the contract with probability $\overline{\lambda_{is}} = 1$.[7]

The exporter's problem. An exporter in country e and sector s maximizes her expected profits from selling to country i, which are given by:

$$\pi_{eis} = \overline{\lambda_{is}} p_{eis} x_{eis} - w_e \tau_{ei} x_{eis} - f_{ei} \tag{2.1}$$

Exporters choose prices recognizing the risk of default. Following Antràs and Foley (forthcoming), equation (2.1) assumes that importers have no wealth and are protected by limited liability, so that they cannot pay beyond the market value of the purchased goods.[8]

The exporter decides on the optimal price p_{eis}, taking as given the demand for her varieties, $x_{eis} = \left(\frac{p_{eis}}{P_{is}}\right)^{-\sigma} \frac{\mu_s P_i Y_i}{P_{is}}$, where Y_i, P_i and $P_{is} = \left(\int_{\Omega_{is}} p_{is}(\omega)^{1-\sigma} d\omega\right)^{1/(1-\sigma)}$ are specific to the importer's country, and represent aggregate real expenditure (or, with balanced trade, real GDP), the overall consumer price index, and the price index in sector s, respectively. We treat Y_i, P_i and P_{is} as exogenous and solve for

[7] Hoefele et al. (2013) and Demir and Javorcik (2014) find empirically that for a given quality of institutions in the importing country, more complex goods are less likely to be exported on open account terms, and more likely to be exported on cash in advance or bank-intermediated terms. Yet, using detailed exports data, Ahn (2014) (Chile and Colombia) and Demir and Javorcik (2014) (Turkey) report that the share of complex goods traded on open account terms is very high—around 70% to 80%.

[8] For a related model that incorporates incentive-compatibility and participation constraints to enforce international payments, see chapter 3 in Antràs (2015).

the optimal sectoral export price and quantity decisions in partial equilibrium.

2.2.2 Main theoretical predictions

In equilibrium, the export value in sector s is given by:

$$p_{eis}x_{eis} = \left[\frac{\sigma}{\sigma-1}\frac{1}{\overline{\lambda_{is}}}w_e\tau_{ei}\right]^{1-\sigma}\frac{\mu_s P_i Y_i}{P_{is}^{1-\sigma}} \tag{2.2}$$

Equation (2.2) shows that the export value is a function of standard variables (constant markup over marginal cost, relative price, and sectoral expenditure), but is also an increasing function of the probability of contract enforcement $\overline{\lambda_{is}}$. Intuitively, the riskiness of the transaction acts as wedge on the price, and this wedge increases when the exporter is more likely to face a default. Therefore, the lower the $\overline{\lambda_{is}}$, the higher is the optimal price p_{eis} and the lower is the quantity exported to country i, x_{eis}. The model thus predicts that for a given industry s, a "crisis" in country i (represented by a decline in Y_i) reduces the export value to that country, $p_{eis}x_{eis}$, because of the assumed procyclical movement of λ_i (and hence of $\overline{\lambda_{is}}$; first term in equation (2.2)). This mechanism works on top of a direct demand effect (second term in (2.2)).

Furthermore, the impact of a crisis in the importing country on $p_{eis}x_{eis}$ is amplified in more contract-dependent industries. Formally, consider the effect of the industry measure of contractual vulnerability z_s on the export value response to a decline in Y_i. The elasticity of the sectoral export value with respect to Y_i is:

$$\varepsilon_{px,s} \equiv -\frac{\partial p_{eis}x_{eis}}{\partial Y_i}\frac{Y_i}{p_{eis}x_{eis}} = (1-\sigma)\, z_s\frac{\lambda'(Y_i)Y_i}{\lambda_i} - 1 \tag{2.3}$$

Since $1-\sigma < 0$, equation (2.3) shows that $\varepsilon_{px,s} < 0$. The first term on the right-hand side of (2.3) again indicates that, all else equal, sectoral exports fall as macroeconomic conditions in the destination country i deteriorate and the country-specific probability of contract enforcement λ_i decreases. But crucially, a higher value of z_s magnifies the decline in exports in industry s to country i. This prediction constitutes our main testable implication. Meanwhile, the second term on the right-hand side of (2.3) implies a unit demand elasticity, common to all industries, which naturally follows from our CES demand assumption.

In the absence of firm or consumer heterogeneity, the predictions of the model are directly applicable to country-industry trade flows.

2.3 Empirical strategy

In this section we explain our empirical methodology and describe the data to be used in the regression analysis. The sources of our data are summarized in Appendix Table B.1.

2.3.1 Methodology

We estimate the following baseline equation to test the hypothesis that the negative trade effects of a crisis in the destination country are amplified in industries with higher contractual vulnerability:

$$\begin{aligned} lnX_{eist} = \sum_{k=1}^{3} \alpha_k \, Crisis_{it}^k + \sum_{k=1}^{3} \beta_k \, Crisis_{it}^k \times z_s + \\ \eta \Upsilon_{it} + \delta \Theta_{et} + \varphi \Psi_{eit} + \gamma_{eis} + \gamma_t + \varepsilon_{eist}, \end{aligned} \tag{2.4}$$

where lnX_{eist} represents the log of exports of country e to country i in industry s at time t. $Crisis_{it}^k$ is an indicator variable that takes the value of 1 if the importing country i suffers a crisis at time t, and 0 otherwise. The superscript $k \in \{1, 2, 3\}$ denotes a specific measure of crisis, as defined below. In line with the model, we expect the coefficients associated with crises to be negative ($\alpha_k < 0$). These coefficients capture the average effect of crises on industry imports of the crisis country.

We incorporate three measures of crises in the regressions. In particular, $Crisis^k$ is defined as:

$$Crisis^k = \begin{cases} \text{Recession alone} & \text{, if } k = 1 \\ \text{Financial Crisis alone} & \text{, if } k = 2 \\ \text{Recession and Financial Crisis} & \text{, if } k = 3 \end{cases},$$

where 'Recession alone' captures an economic downturn without a financial crisis, 'Financial Crisis alone' characterizes a financial disruption that is not accompanied by a recession, and 'Recession and Financial Crisis' captures the simultaneous occurrence of both of these events.

To identify the amplification effect of crises in industries with higher contractual vulnerability, we include interaction terms of $Crisis_{it}^k$ with z_s, a demeaned index that represents the degree of contractual vulnerability of industry s. Our model's key prediction is that the coefficients associated with these interaction terms are negative ($\beta_k < 0$). That is, imports of the crisis country decline disproportionately in more contract-dependent industries. We identify β_k by relying on the variation of contractual vulnerability across industries, and the occurrence (or not) of crises in importing countries across years.

Additionally, equation (2.4) includes a first set of control variables, Υ_{it}, that contains the log of real GDP (as a proxy for market size and demand) and the degree of financial development of the importing country. A second set of controls, Θ_{et}, includes the log of real GDP and the degree of financial development of the exporting country. The final set of controls, Ψ_{eit}, includes the log of the bilateral real exchange rate and a dummy variable that captures whether the trading partners have a free trade agreement at time t.

We add proxies for financial development in the estimating equation for three reasons. First, financing conditions at the country level affect decisions on trade finance terms (e.g., using open account or cash in advance), as documented in Antràs and Foley (forthcoming) and Hoefele et al. (2013). Second, trade is intensive in working capital, and as such it depends on financial conditions (Manova, 2013). Third, financial development might reflect to some extent the general contractual environment.[9] The regression also includes an interaction term of financial development of both the exporting and importing country. We expect a negative coefficient on this interaction term under the consideration that the more financially developed

[9]Measures of contractual enforcement at the country level are typically unavailable for a wide range of countries *and* for a long span of years. Some indicators included in the International Country Risk Guide constitute an exception, but these data are not publicly available.

the exporting country is, the lower may be the relevance of the importing country as a source of financing for trade, and vice versa.

Equation (2.4) also includes fixed effects at the exporter-importer-industry level, γ_{eis}, and at the year level, γ_t. The inclusion of γ_{eis} in equation (2.4) accounts for time-invariant bilateral characteristics such as distance, common language, contiguity or colonial links, and any specific relationship between a pair of trading partners at the industry level.[10] Additionally, γ_{eis} also accounts for the time-invariant component of multilateral trade resistance effects (Anderson and van Wincoop, 2003). Finally, γ_t captures factors that affect all countries in the same period, such as global recessions or changes in commodity prices. We compute clustered standard errors at the importing country-year level. Below we perform several robustness checks.

2.3.2 Data

Country-industry trade flows. We use annual data on bilateral trade flows obtained from the Feenstra et al. (2005) World Trade Flows database. These data are originally organized by the 4-digit Standard International Trade Classification (SITC), Revision 2. Since our key industry variables are constructed for 4-digit U.S. Standard Industrial Classification (SIC) industries, we convert the trade data to this format by replicating the concordance method from Cuñat and Melitz (2012).[11]

[10]By specific relationship we mean, for example, a situation in which the exporter may not be not selling exactly the same product to every destination, or using the same payment method to sell a product across different destinations (to the extent that payment methods remain relatively stable over time).

[11]We add up the value of disaggregated 10-digit Harmonized System (HS) U.S. annual exports for the period 1989–2000, using the dataset constructed by Feenstra et al. (2002). Since this dataset includes a concordance between HS, SITC and SIC categories, we derive concordance weights to map the SITC codes into SIC categories. A similar procedure is also employed in Chor (2010).

Our sample excludes zero trade flows, nonmanufacturing industries, and the oil sector represented by the SIC code 2911. We deflate the export flows (originally reported in current U.S. dollars) by using the world export price index from the International Financial Statistics database. The results presented below, however, are robust to using nominal trade values instead of real ones. Our final sample covers the period 1989–2000 and it includes 127 exporting countries, 118 importing countries, and (in most of our regressions) 351 SIC industries.[12] We show a list of the countries included in the sample in Table 2.1.

Recessions and financial crises. We identify crisis periods in importing countries as years when these countries experience recessions alone, financial crises alone, or both recessions and financial crises at the same time. In line with the spirit of the theoretical model, we think of these events as periods of increasing importers' risk of default.

We use real GDP (obtained from the World Development Indicators database) and the methodology of Braun and Larrain (2005) to construct indicators for recessions. A recession in a given country is defined following a peak-to-trough criterion—a trough occurs when cyclical GDP is more than one standard deviation below zero; a local peak associated with a trough is a year in which cyclical GDP is higher than in both the previous and the posterior years.[13] We checked that our results are

[12]The endpoint in our sample period is determined by data availability, as the World Trade Flows database is constructed until the year 2000. We start the analysis in 1989 because our concordance method relies on the SITC to SIC-87 mapping that is readily available in the Feenstra et al. (2002) dataset only since 1989 (see footnote 11). Our sample captures several clusters of recessions and crises during the 1990s, as detailed below. We report a sensitivity analysis using data since 1980.

[13]The cyclical component of GDP is computed using the Hodrick-Prescott filter with a lambda parameter value of 6.25 (Ravn and Uhlig, 2002). Whenever available, the cyclical component of GDP is constructed using data from 1960 to 2012.

Table 2.1
List of countries

Albania	Czech Republic	Kenya	Portugal
Angola	Denmark	Korea, Rep.	Russian Federation
Argentina	Dominican Republic	Kuwait	Rwanda
Armenia	Ecuador	Kyrgyz Republic	Saudi Arabia*
Australia	Egypt, Arab Rep.	Lao PDR	Senegal
Austria	El Salvador	Latvia	Singapore
Azerbaijan	Equatorial Guinea	Lithuania	Slovak Republic
Bahamas *	Estonia*	Macao*	Slovenia*
Bahrain*	Ethiopia	Madagascar	South Africa
Bangladesh	Fiji	Malawi	Spain
Barbados	Finland	Malaysia	Sri Lanka
Belarus	France	Mali	St. Kitts and Nevis*
Belgium	Gabon	Malta*	Sudan
Belize	Gambia	Mauritania	Suriname
Benin	Georgia	Mauritius	Sweden
Bolivia	Germany	Mexico	Switzerland
Brazil	Ghana	Moldova	Syrian Arab Republic
Bulgaria	Greece	Mongolia	Tanzania
Burkina Faso	Guatemala	Morocco	Thailand
Burundi	Honduras	Mozambique	Togo
Cambodia	Hong Kong	Nepal	Trinidad and Tobago
Cameroon	Hungary	Netherlands	Tunisia
Canada	Iceland	New Zealand	Turkey
Central African Republic	India	Niger	Uganda
Chad	Indonesia	Nigeria	Ukraine
China	Iran, Islamic Rep.	Norway	United Kingdom
Colombia	Ireland	Pakistan	United States
Congo, Rep.	Israel	Panama	Uruguay
Costa Rica	Italy	Paraguay	Vietnam
Cote d'Ivoire	Japan	Peru	Yemen, Rep.
Croatia *	Jordan	Philippines	Zambia
Cyprus	Kazakhstan	Poland	

Note: An asterisk (*) indicates countries that appear in the sample only as exporters.

not affected by using other definitions of recessions, such as years of negative GDP growth rates.

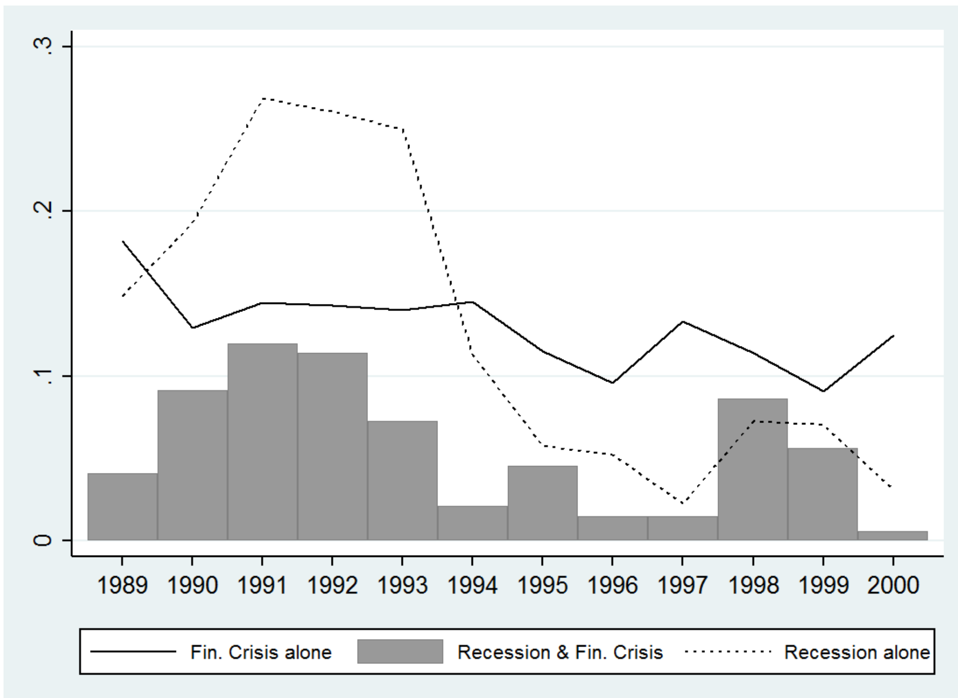
We also define an indicator for financial disruptions. Following Abiad et al. (2014), we identify financial crisis episodes as periods of banking or sovereign debt crises, based on the Laeven and Valencia (2013) database that covers 129 countries from 1970 to 2011. Laeven and Valencia's criteria to define a systemic banking crisis are: (i) significant signs of distress in the banking sector, such as liquidations, losses, and/or bank runs; and (ii) significant banking policy intervention measures in response to losses in the banking system. Laeven and Valencia also report sovereign debt crises as episodes of sovereign debt default and/or restructuring. Importantly, their data shows a marked increase in the number of crises during the 1990s, a period that we fully cover in our analysis.[14]

Of the 118 importing countries in our sample, 78 (63) suffered a recession (financial crisis) at some point between 1989 and 2000. The mean duration of a recession is close to 2 years, and the mean duration of a financial crisis is almost 4 years. Focusing on the measures of crises employed in our regressions, Figure 2.1 shows that the share of recessions alone reaches a peak of nearly 30% in the early 1990, while the maximum share of observations characterized by both recessions and financial crises is around 10%. The share of financial crises alone is fairly stable across years (13%, on average) but it tends to decline slowly over the sample period. The graph also reveals a comovement over time between the occurrence of recessions

[14]Countries of different levels of income experienced financial crises during the 1990s (e.g., Sweden, Malaysia, Mexico, Indonesia, and Kenya). Other spikes in the number of crises (which are not covered by our sample period) are found in the early 1980s and during the Great Recession, particularly in 2008.

alone and the coincidence of recessions and financial crises, although the share of recessions alone in most years is greater than that of the combination of recessions and financial crises.

Figure 2.1
Share of observations with crises, by year



Notes: Recessions are identified using the Braun and Larrain (2005) methodology. Financial crises are identified as banking or sovereign debt crises, using the Laeven and Valencia (2013) dataset.

In our data, the existence of a non-trivial share of observations characterized by 'financial crises alone' is explained in part by the fact that sovereign debt crises often last more years than the average recession. (On the contrary, banking crises are relatively short-lived and tend to be more closely correlated with economic downturns.) In a few other cases, the identification of periods simultaneously characterized by recessions and financial crises is limited by the existence of missing observations on some country GDP series.

Contractual vulnerability across industries. We need to identify industry measures of contractual vulnerability as proxies for the industry-specific components of contract enforcement described in the model.[15] We first follow the literature and use the Nunn (2007) and Levchenko (2007) indices. These are available for our desired level of sectoral disaggregation and are constructed using U.S. Input-Output Tables. As is standard in related papers, we assume that the ranking of industries remains stable across countries. This is a plausible assumption to the extent that both of these indices reflect technological factors.[16]

Nunn (2007) aims to measure the contract intensity of industries, which he defines as the fraction of an industry's intermediate inputs that are relationship-specific (i.e., that are either not traded on an organized exchange or for which no reference price exists). A higher value of the Nunn (2007) index reflects a higher degree of an industry's sensitivity to imperfect institutions.[17] Some of the most contract intensive industries include Motor Vehicles and Passenger Car Bodies, Electronic Computers, and Electromedical and Electrotherapeutic Apparatus; some of the least contract intensive industries are Poultry Slaughtering and Processing, Primary Smelting and Refining of Copper, and Rice Milling. These examples are useful to illustrate the relationship between product complexity and contract dependence described in the

[15] To our knowledge, there are no publicly available comprehensive datasets on firm defaults on international transactions for disaggregated industries.

[16] We thank Davin Chor for kindly sharing his data on the Nunn and the Levchenko indices at the 4-digit SIC level.

[17] In our analysis, the index corresponds to the z^{rs1} measure specified in Nunn (2007). We use the Nunn index that relies on the Rauch (1999) conservative classification for its construction. For more details, see Chor (2010) and its supplementary appendix.

model.

The Levchenko (2007) index measures the sensitivity of an industry to institutions such as contract enforcement and property rights. This index equals one minus the Herfindahl index of an industry's intermediate input use—an inverse measure of the concentration mix of inputs, and hence a direct measure of exposure to hold-up problems in the production process. Among the most institutionally intensive industries are Fluid Power Pumps and Motors, Small Arms Ammunition, and Surgical Appliances and Supplies; among the least institutionally intensive industries are Meat Packing Plants, Creamery Butter, and Setup Paperboard Boxes.

We also use a novel measure of uncollected credit sales, labeled as "uncollectible index", as an additional index of industry contractual vulnerability. The source of these data is the National Summary of Domestic Trade Receivables (NSDTR), a proprietary quarterly survey of large U.S. firms compiled by the Credit Research Foundation (CRF). As detailed in Appendix B.2, we construct our index as $1 - CEI$, where CEI stands for the NSDTR's Collection Effectiveness Index. The CEI is acknowledged by the CRF as the most effective measure of credit and collection performance. Our uncollectible index captures the share of total account receivables uncollected compared to what was available to collect over a quarter. The CEI is originally reported in the NSDTR as a median value for every 4-digit SIC industry that registers at least 3 respondent firms.[18]

[18]The CRF (`http://www.crfonline.org/`) is a non-profit, member-run organization. Its members include a large number of Fortune 1000 corporations. The NSDTR constitutes a unique data source of performance indicators of domestic accounts receivable, defined as claims against customers for goods sold domestically on credit, based on the answers of hundreds of Fortune 1000 U.S. firms from a broad cross section of industries. To our knowledge, the NSDTR has not been used in recent academic literature. In the early years of the survey, however, Seiden (1964) used

The uncollectible index is advantageous for our purposes as it reflects, by construction, payment defaults in business-to-business transactions at the industry level. In this sense, our index is better suited to capture the role of contractual imperfections than other standard indicators of trade credit intensity, such as the ratio of accounts receivable to sales and the ratio of accounts payable to the cost of goods sold.[19] To isolate the structural component of the uncollectible index we take industry medians across quarters (2006q1–2010q4). A ranking of industries based on the uncollectible index is displayed in Appendix Table B.2.

In using the uncollectible index for our empirical analysis, we assume that domestic receivable performance can proxy for the quality of collection of foreign receivables. We also believe that since large firms are dominant in international trade, the sample of Fortune 1000 firms surveyed by the CRF are representative of firms engaged in overseas transactions. That said, we acknowledge that the uncollectible index may not be, as desired, completely exogenous from the perspective of firms. Another limitation is that we only have available data to construct this index for 110 industries.

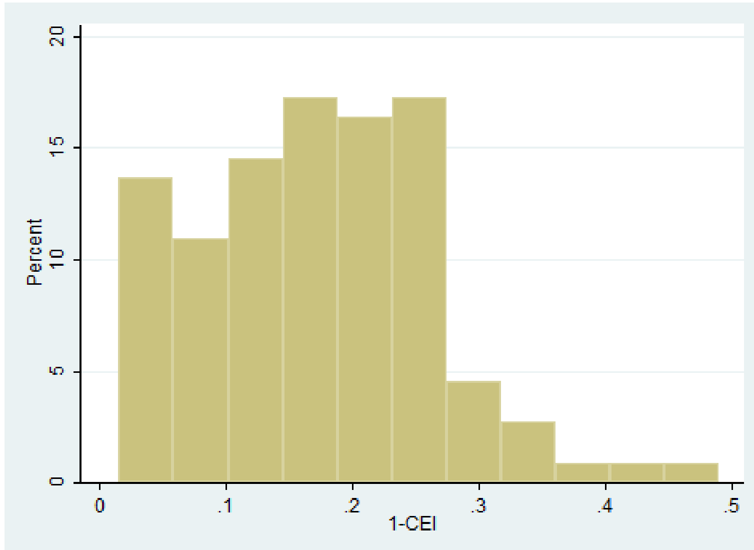
Figure 2.2 shows the distribution of the uncollectible index. Over one third of the industries fail to collect 10% of their receivables or less. On the opposite extreme, only about 10% of industries fail to collect 30% of their receivables or more.

---

it in his pioneering study of the quality of trade credit, and Nadiri (1969) employed it to calculate the delinquency rate on manufacturing accounts receivable and payable.

[19]It is worth noting that as part of our robustness checks, we control for the cash conversion cycle, an industry measure of financial dependence constructed as 365 $\times$ accounts receivable/sales $-$ 365 $\times$ accounts payable/cost of goods sold + 365 $\times$ inventories/cost of goods sold.

Figure 2.2
Distribution of the uncollectible index



Notes: The bars represent the histogram of the uncollectible index. The uncollectible index is constructed as $1 - CEI$, where CEI is the Collection Effectiveness Index reported in the Credit Research Foundation's National Summary of Domestic Trade Receivables. We calculate (4-digit SIC) industry medians over the period 2006q1-2010q4, and divide them by 100 to express them as decimals.

Table 2.2 summarizes some descriptive statistics of our contractual vulnerability indices. In this Table, the Nunn index is reported as "complexity" and the Levchenko index appears as "concentration". We maintain this notation in our regression analysis below. As shown in Table 2.3, our three indices are positively and significantly correlated at the 1% level. We find the highest correlation coefficient between the complexity and the concentration indices. (Tables 2.2 and 2.3 also report statistics and pairwise correlations for other industry indicators that will be described below.)

Country-level data. As part of our control variables, we use information on Free Trade Agreements (FTA) from de Sousa (2012). The bilateral real exchange rate is

Table 2.2
Summary statistics: indicators of contractual and financial vulnerability, at industry level

Variable	Obs.	Mean	Std.Dev.	Min	Max
Complexity	351	0.56	0.25	0.00	0.98
Concentration	351	0.86	0.11	0.21	0.97
Uncollectible	110	0.17	0.09	0.02	0.49
CashCycle	351	0.94	0.34	0.30	2.06
TangAssets	351	0.40	0.15	0.14	0.88

Notes: Industries are classified by 4-digit SIC. *Complexity* is the input relationship-specificity index from Nunn (2007). *Concentration* is the input concentration index from Levchenko (2007). The source of these indices is Chor (2010). *Uncollectible* is the account receivables' collection ineffectiveness index, based on data from the Credit Research Foundation's National Summary of Domestic Trade Receivables. *CashCycle* is a measure of the time elapsed between the moment a firm pays for its materials until the collection on its sales (in hundreds of days). *TangAssets* is a measure of tangible assets developed by Braun (2003). The last two measures are constructed using data from Compustat. See text for further details.

Table 2.3
Pairwise correlation coefficients: indicators of contractual and financial vulnerability

	Complexity	Concentration	Uncollectible	CashCycle
Concentration	0.52			
	(0.00)			
Uncollectible	0.45	0.26		
	(0.00)	(0.01)		
CashCycle	0.50	0.36	0.43	
	(0.00)	(0.00)	(0.00)	
TangAssets	-0.57	-0.30	-0.18	-0.44
	(0.00)	(0.00)	(0.07)	(0.00)

Notes: For definitions of the variables, see notes to Table 2.2. Correlations are computed across 4-digit SIC industries, with p-values reported in parentheses.

constructed using data from the Penn World Table 8.1 (Feenstra et al., forthcoming). Finally, financial development is proxied by the ratio of private credit by banks and other financial institutions to GDP (Beck et al., 2000).

Table 2.4 displays summary statistics for all of our country-level variables, including GDP and our indicators of crises. (The Table also includes statistics for other variables which will be used in our sensitivity analysis below.)

Table 2.4
Summary statistics: trade and country-level data

Variable	Obs.	Mean	Std.Dev.	Min	Max
Recession alone	5538895	0.11	0.32	0.00	1.00
Fin.Crisis alone	5538895	0.13	0.33	0.00	1.00
Recession & Fin.Crisis	5538895	0.05	0.22	0.00	1.00
Ln BRER	5538895	4.72	0.76	1.92	7.23
Ln GDP Imp.	5538895	25.77	2.00	18.80	30.08
Ln GDP Exp.	5538895	26.64	1.62	18.80	30.08
Fin. Develop.	5538895	0.65	0.47	0.00	2.28
FTA	5538895	0.23	0.42	0.00	1.00
Contract Enforcement	5538895	0.69	0.46	0.00	1.00

Notes: *Ln* denotes natural logarithm and the variable names correspond to those employed in the regression analysis. *Recession alone*, *Fin.Crisis alone*, and *Recession & Fin.Crisis* are dummy variables at the importing-country level; *BRER* is the bilateral real exchange rate; *GDP Imp.* and *GDP Exp.* are the real GDP in importing and exporting countries, respectively; *Fin. Develop.* is the ratio of private credit to GDP; *FTA* denotes free trade agreement (dummy variable); and *Contract Enforcement* is measured by the rule of law. See text and Appendix Table B.1 for further details.

2.4 Results

This section shows the results of estimating our baseline regression and different robustness exercises. The total number of data points in most of the regressions is above 5 million. When we use the uncollectible index as our contractual-vulnerability measure, the number of observations decreases because fewer industries have values for this index relative to Nunn's (2007) complexity and Levchenko's (2007) concentration indices.[20]

2.4.1 Baseline results

Table 2.5 presents the results of the OLS estimation of equation (2.4), using our three industry indices of contractual dependence. We first note that the coefficients on most of the control variables are significant and have the expected signs. Although the coefficient associated with financial development in the exporting country is not statistically significant, it is always positive, as expected.

In column 1, we show that when a country is hit by a crisis, the average impact on its industry imports is negative, even after controlling for demand. Following the notation of equation (2.4), the point estimate of α_k is statistically significant at the 1% level when a crisis is measured either by a recession alone (α_1) or a financial crisis accompanied by a recession (α_3), but not significant for a financial crisis alone (α_2). As implied by the significant coefficients, industry imports of the crisis country decline on average 5.5% ($exp(-0.057) - 1$) following a pure recession, and 18.1% ($exp(-0.200)-1$) in the aftermath of a recession with financial crisis. Our quantitative estimate of the combined effect of recessions and financial disruptions on imports is similar to the findings of Abiad et al. (2014). They document that in the year after a crisis, imports fall, on average, 19%.

[20] Our panel is unbalanced since not all countries trade in all industries and years. Moreover, not all of our control variables are available for every country over the entire sample period.

Table 2.5

Effects of crises and contractual vulnerability on trade across countries and industries. Dependent variable: Ln(bilateral sectoral imports)

	(1)	(2)	(3)	(4)
Recession alone	-0.057***	-0.054***	-0.055***	-0.056***
	(0.014)	(0.014)	(0.014)	(0.015)
Fin.Crisis alone	-0.012	-0.011	-0.011	-0.004
	(0.022)	(0.021)	(0.022)	(0.022)
Recession & Fin.Crisis	-0.200***	-0.195***	-0.196***	-0.190***
	(0.033)	(0.032)	(0.032)	(0.033)
Recession alone × Complexity		-0.132***		
		(0.026)		
Fin.Crisis alone × Complexity		-0.035		
		(0.036)		
Recession & Fin.Crisis × Complexity		-0.224***		
		(0.038)		
Recession alone × Concentration			-0.276***	
			(0.046)	
Fin.Crisis alone × Concentration			-0.089	
			(0.064)	
Recession & Fin.Crisis × Concentration			-0.401***	
			(0.069)	
Recession alone × Uncollectible				-0.296***
				(0.056)
Fin.Crisis alone × Uncollectible				-0.066
				(0.085)
Recession & Fin.Crisis × Uncollectible				-0.432***
				(0.100)
Ln BRER	-0.322***	-0.322***	-0.322***	-0.332***
	(0.035)	(0.035)	(0.035)	(0.035)
Ln GDP Imp.	0.736***	0.736***	0.736***	0.695***
	(0.095)	(0.095)	(0.095)	(0.097)
Ln GDP Exp	1.400***	1.400***	1.401***	1.430***
	(0.036)	(0.036)	(0.036)	(0.038)
FTA	0.272***	0.272***	0.272***	0.275***
	(0.028)	(0.028)	(0.028)	(0.027)
Fin. Develop. Imp.	0.415***	0.415***	0.415***	0.447***
	(0.058)	(0.058)	(0.058)	(0.058)
Fin. Develop. Exp.	0.030	0.030	0.030	0.032
	(0.030)	(0.030)	(0.030)	(0.032)
Fin. Develop. Imp*Exp	-0.175***	-0.175***	-0.175***	-0.189***
	(0.033)	(0.033)	(0.033)	(0.034)
Observations	5,538,895	5,538,895	5,538,895	2,077,588
R-squared	0.897	0.897	0.897	0.903
Importer-Exporter-Industry FE	YES	YES	YES	YES
Year FE	YES	YES	YES	YES

Notes: Standard errors are clustered by destination-year, with ***,** ,* respectively denoting significance at the 1%, 5% and 10% levels. 'Recession' and 'Fin.Crisis' are associated with importing countries. 'Complexity', 'Concentration', and 'Uncollectible' are demeaned.

Once we include the interactions of the crisis indicators with our industry measures of contractual vulnerability (columns 2-4 in Table 2.5), we observe that the estimated coefficients on these interaction terms are negative ($\beta_k < 0$). These coefficients are statistically significant, except when crises are measured as financial crises alone. In short, we confirm our key hypothesis that imports in more contract-dependent industries are disproportionately affected by crises—especially when crises involve recessions. This result is robust to the use of the complexity, concentration or uncollectible indices. For the case of the complexity index, our estimates imply that a one-standard-deviation increase from its mean magnifies the decline in imports from 5.5% to 8.6% following a recession alone, and from 18.1% to 23.1% following a recession and financial crisis.[21]

To further explore the amplification effect induced by contractual vulnerability at the industry level, we use our estimates for β_k and compare the differential impact of crises on trade across two specific industries. We define an industry in the 25^{th} percentile of each contractual-vulnerability index as a 'slightly contract-dependent' industry. Similarly, we define an industry in the 75^{th} percentile of each contractual-vulnerability index as a 'highly contract-dependent' industry. Table 2.6 summarizes the results.

Focusing on the complexity index, the impact of a recession and financial crisis on imports is 8.6 percentage points ($exp(-0.090) - 1$) larger in the highly contract-dependent industry (Printed Circuit Boards, SIC 3672) than in the slightly contract-

[21] Analogously, the magnification effects on imports for the case of the concentration index are from 5.5% to 8.2% (recession alone) and from 18.1% to 21.7% (recession and financial crisis). For the uncollectible index, such effects are from 5.5% to 7.8% (recession alone) and from 18.1% to 21.3% (recession and financial crisis).

Table 2.6
Amplification effects of contractual vulnerability on industry imports: differential impact of crises ($75^{th} - 25^{th}$ pctl)

	Recession alone (1)	Fin.Crisis alone (2)	Recession & Fin.Crisis (3)
Complexity	-0.053***	-0.014	-0.090***
	(0.010)	(0.015)	(0.015)
Concentration	-0.030***	-0.010	-0.043***
	(0.005)	(0.007)	(0.007)
Uncollectible	-0.039***	-0.009	-0.057***
	(0.007)	(0.011)	(0.013)

Notes: The calculations are based on the estimates from Table 2.5 (columns 2-4), with ***,** ,* respectively denoting significance at the 1%, 5% and 10% levels. Standard errors are reported in parentheses. The table shows the differential impact of crises on imports between highly and slightly contract-dependent industries, for each industry measure of contractual vulnerability. The calculations assume that a highly (slightly) contract-dependent industry is in the 75^{th} (25^{th}) percentile in the distribution of our measures of contractual vulnerability.

dependent industry (Steel Works, Blast Furnaces, and Rolling Mills, SIC 3312). For the uncollectible index, following a recession and financial crisis, the industry with high contract dependence exhibits a 5.5 percentage points ($exp(-0.057) - 1$) larger drop in imports than the industry with low dependence (Construction Machinery and Equipment, SIC 3531; and Paper Mills, SIC 2621, respectively). All the reported differences between the 75^{th} and the 25^{th} percentiles are statistically significant (at the 1% level), except when crises are measured as financial crises alone. Noticeably also, the amplification effects conditional on a recession and financial crisis (column 3) are larger than those associated with a recession alone (column 1), roughly by a factor of 1.5, on average.

2.4.2 Sensitivity analysis

To check the robustness of our findings we consider different exercises. We begin by dividing the sample of importing countries according to their rule of law index. We then verify the sensitivity of our results to changing the sample period, introducing additional controls, and using alternative specifications of the estimating equation. For the remainder of this section, all of the regressions include the same control variables as the baseline estimation, but to save space they are omitted in the tables.

2.4.2.1 Contract enforcement (rule of law) at the country level

We first test if the decline of imports among industries with higher contractual dependence is more pronounced in countries with lower *structural* levels of contractual enforcement. This would be a reasonable outcome if, independently of the industry, importing firms were more likely to default in countries with worse institutional quality (see, e.g., Schmidt-Eisenlohr, 2013); or alternatively, if poor institutions disproportionately exacerbated the risk of default of more contract-dependent industries. To measure a country's ability to enforce contracts we use the rule of law index from Kaufmann et al. (2010). We then split the sample according to whether importing countries are above or below the median value of this indicator. The results are shown in Table 2.7.

We find that, for the case of recessions and financial crises, the amplification effects due to contractual vulnerability at the industry level are indeed larger in

Table 2.7

Effects of crises and contract vulnerability on trade across countries and industries: contract enforcement at importing-country level. Dependent variable: Ln(bilateral sectoral imports)

	Below median (Low enforcement)				Above median (High enforcement)			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Recession alone	-0.135***	-0.130***	-0.132***	-0.134***	-0.030**	-0.028*	-0.029**	-0.027*
	(0.040)	(0.040)	(0.040)	(0.041)	(0.015)	(0.015)	(0.015)	(0.014)
Fin.Crisis alone	-0.037	-0.037	-0.036	-0.031	0.010	0.012	0.011	0.021
	(0.031)	(0.030)	(0.030)	(0.031)	(0.027)	(0.026)	(0.026)	(0.026)
Recession & Fin.Crisis	-0.251***	-0.245***	-0.245***	-0.250***	-0.126***	-0.122***	-0.123***	-0.102***
	(0.048)	(0.047)	(0.048)	(0.049)	(0.036)	(0.036)	(0.036)	(0.034)
Recession alone × Complexity		-0.165**				-0.123***		
		(0.070)				(0.028)		
Fin.Crisis alone × Complexity		-0.000				-0.088		
		(0.043)				(0.064)		
Recession & Fin.Crisis × Complexity		-0.244***				-0.202***		
		(0.057)				(0.049)		
Recession alone × Concentration			-0.231**				-0.279***	
			(0.115)				(0.050)	
Fin.Crisis alone × Concentration			-0.079				-0.107	
			(0.085)				(0.098)	
Recession & Fin.Crisis × Concentration			-0.452***				-0.328***	
			(0.105)				(0.082)	
Recession alone × Uncollectible				-0.270**				-0.298***
				(0.137)				(0.061)
Fin.Crisis alone × Uncollectible				0.034				-0.227*
				(0.115)				(0.131)
Recession & Fin.Crisis × Uncollectible				-0.454***				-0.384***
				(0.135)				(0.147)
Observations	1,694,895	1,694,895	1,694,895	661,887	3,844,000	3,844,000	3,844,000	1,415,701
R-squared	0.858	0.858	0.858	0.861	0.908	0.908	0.908	0.916
Importer-Exporter-Product FE	YES	YES	YES	YES	YES	YES	YES	YES
Year FE	YES	YES	YES	YES	YES	YES	YES	YES

Notes: See notes to Table 2.5. Standard errors are clustered by destination-year, with ***,**,* respectively denoting significance at the 1%, 5% and 10% levels. We use the rule of law index from Kaufmann et al. (2010) to compute country averages for the period 1996-2012. We then calculate the median value across countries and split the sample according to whether importing countries are below or above that median. The regressions include the same control variables as the baseline estimation (see Table 2.5).

countries with low contract enforcement. Illustratively, as observed in columns and 2 and 4, the coefficient attached to the interaction of the recession and financial crisis indicator and the complexity index is 21% higher (in absolute value) in the low-enforcement sample than in the high-enforcement sample (-0.244 and -0.202, respectively).

However, when crisis are measured by recessions alone, the evidence is mixed. The estimated coefficients on the interaction terms are higher (in absolute value) in the low-enforcement sample than in the high-enforcement sample for the case of the complexity index, but the opposite result holds when we use the concentration and the uncollectible indices. Lastly, in both country samples, the point estimates on the interaction terms with financial crisis alone are in most cases statistically not significant.

Note also that crises associated with recessions have much larger average impacts on sectoral imports of countries with low contract enforcement (columns 1 and 5). For example, in the wake of recessions alone, the estimated drop in imports for this group of countries is 12.6% ($exp(-0.135) - 1$), compared to 3.0% ($exp(-0.030) - 1$) for countries with high enforcement.

2.4.2.2 Extending the sample period

We next evaluate the robustness of our results to the use of data since 1980. The results are reported in Table 2.8. In contrast to our baseline results, the coefficient that captures the average effect of a financial crisis alone on industry imports

(α_2) is statistically significant at the 5% and 1% levels (columns 1-3 and column 4, respectively).

Table 2.8
Effects of crises and contract vulnerability on trade across countries and industries: extended sample period (1980-2000). Dependent variable: Ln(bilateral sectoral imports)

	(1)	(2)	(3)	(4)
Recession alone	-0.062***	-0.059***	-0.060***	-0.053***
	(0.016)	(0.016)	(0.016)	(0.016)
Fin.Crisis alone	-0.053**	-0.053**	-0.052**	-0.041*
	(0.023)	(0.023)	(0.023)	(0.023)
Recession & Fin.Crisis	-0.184***	-0.181***	-0.181***	-0.172***
	(0.028)	(0.027)	(0.027)	(0.027)
Recession alone × Complexity		-0.254***		
		(0.041)		
Fin.Crisis alone × Complexity		0.023		
		(0.048)		
Recession & Fin.Crisis × Complexity		-0.164***		
		(0.049)		
Recession alone × Concentration			-0.360***	
			(0.058)	
Fin.Crisis alone × Concentration			-0.065	
			(0.067)	
Recession & Fin.Crisis × Concentration			-0.290***	
			(0.066)	
Recession alone × Uncollectible				-0.711***
				(0.146)
Fin.Crisis alone × Uncollectible				0.179
				(0.145)
Recession & Fin.Crisis × Uncollectible				-0.190
				(0.161)
Observations	7,830,508	7,830,508	7,830,508	2,924,381
R-squared	0.855	0.855	0.855	0.859
Importer-Exporter-Industry FE	YES	YES	YES	YES
Year FE	YES	YES	YES	YES

Notes: See notes to Table 2.5. Standard errors are clustered by destination-year, with ***,** ,* respectively denoting significance at the 1%, 5% and 10% levels. The regressions include the same control variables as the baseline estimation (see Table 2.5).

The coefficients on the interaction terms change in dissimilar ways relative to our baseline results. On one hand, the estimates on the interactions of a recession alone with our three industry measures of contractual vulnerability (β_1) are larger (in absolute value) when we use the extended sample period. On the contrary, the coefficients of the interactions involving a recession with financial crisis (β_3) are smaller (in absolute value) than the baseline estimates, and even not significant for the case of the uncollectible index. Overall, though, this exercise confirms that imports in more contract-dependent industries are disproportionately affected by recessions, either if these are accompanied by financial crises or not.

2.4.2.3 Controlling for financial vulnerability

Although our use of controls and fixed effects aims to mitigate concerns about omitted variables, we allow for the possibility that our industry measures of contractual vulnerability may pick up the effect of financial dependence. A financially dependent industry could be affected by credit constraints as a result of facing high fixed costs or significant working capital needs. In this exercise, we separately include interaction terms of our recession indicator with standard measures of financial dependence, which are based on data from Compustat's annual industrial files (period 1995–2012).

The first industry measure of financial vulnerability is the cash conversion cycle (CashCycle), a proxy for short term financial needs to cover net working capital, defined as the period between a firm's payment for materials and the collection of its sales (Raddatz, 2006). The second one is a measure of asset tangibility (TangAssets), namely the share of net property, plant and equipment in total book-value assets (Braun, 2003).[22] Industries with higher values of CashCycle and with lower values of

[22] We drop firm-year observations with negative or missing values on sales and assets from the

TangAssets are more financially dependent. Table 2.2 reports some summary statistics for these variables. Further, Table 2.3 shows that there are some statistically significant correlations between the contractual and the financial vulnerability indicators. Notably, the complexity index exhibits a somewhat high negative correlation with TangAssets and a positive correlation with CashCycle.

We confirm that the inclusion of the financial vulnerability indicators does not substantially change our main findings. The results obtained with CashCycle and TangAssets are presented in Table 2.9 and Appendix Table B.3, respectively.

The coefficients on the interactions of CashCycle and TangAssets with both recession alone and recession and financial crisis have the expected signs. These coefficients are statistically significant in all regressions, except in the ones that include the complexity index. This likely reflects the correlations between the complexity index and our two indicators of financial dependence, as mentioned above. Meanwhile, the coefficients on the interactions involving financial crisis alone tend to be close to zero and report large standard errors. In short, the bulk of our results imply that trade in more financially-dependent industries is more negatively affected during crises. This is consistent with the evidence in Chor and Manova (2012).

In unreported regressions, we also experiment with the Rajan and Zingales (1998) index of external finance dependence (i.e., the ratio of the difference between

---

Compustat sample. To reduce the effect of outliers, we first sum each firm's value of net property, plant and equipment over the sample period and then divide by the sum of assets over the sample period in order to construct TangAssets. An analogous procedure is followed to aggregate over time the ratios involved in the construction of CashCycle. We then trim both 1% tails of the firm distributions for each of the three measures and calculate industry medians. To gain observations, whenever a median value is not available for a SIC-4 industry, we impose the median value computed for the immediately higher level of aggregation (SIC-3).

Table 2.9
Effects of crises and contract vulnerability on trade across countries and industries: controlling for financial vulnerability.
Dependent variable: Ln(bilateral sectoral imports)

	(1)	(2)	(3)
Recession alone	-0.054***	-0.055***	-0.058***
	(0.014)	(0.014)	(0.015)
Fin.Crisis alone	-0.011	-0.011	-0.003
	(0.021)	(0.021)	(0.022)
Recession & Fin.Crisis	-0.194***	-0.194***	-0.191***
	(0.032)	(0.032)	(0.033)
Recession alone × Complexity	-0.130***		
	(0.026)		
Fin.Crisis alone × Complexity	-0.044		
	(0.036)		
Recession & Fin.Crisis × Complexity	-0.201***		
	(0.040)		
Recession alone × Concentration		-0.252***	
		(0.042)	
Fin.Crisis alone × Concentration		-0.099*	
		(0.054)	
Recession & Fin.Crisis × Concentration		-0.319***	
		(0.066)	
Recession alone × Uncollectible			-0.224***
			(0.052)
Fin.Crisis alone × Uncollectible			-0.105
			(0.072)
Recession & Fin.Crisis × Uncollectible			-0.343***
			(0.087)
Recession alone × CashCycle	-0.003	-0.022*	-0.051***
	(0.011)	(0.012)	(0.018)
Fin.Crisis alone × CashCycle	0.012	0.008	0.028
	(0.015)	(0.016)	(0.021)
Recession & Fin.Crisis × CashCycle	-0.032	-0.069***	-0.063***
	(0.019)	(0.019)	(0.025)
Observations	5,538,895	5,538,895	2,077,588
R-squared	0.897	0.897	0.903
Importer-Exporter-Industry FE	YES	YES	YES
Year FE	YES	YES	YES

Notes: See notes to Table 2.5. Standard errors are clustered by destination-year, with ***,**,* respectively denoting significance at the 1%, 5% and 10% levels. The regressions include the same control variables as the baseline estimation (see Table 2.5). 'CashCycle' is demeaned.

capital expenditures and cash flow over capital expenditures) as an additional measure of financial vulnerability. Our findings do not show a statistically significant role for this variable, while the rest of our key results remain unchanged.

2.4.2.4 Controlling for cyclicality

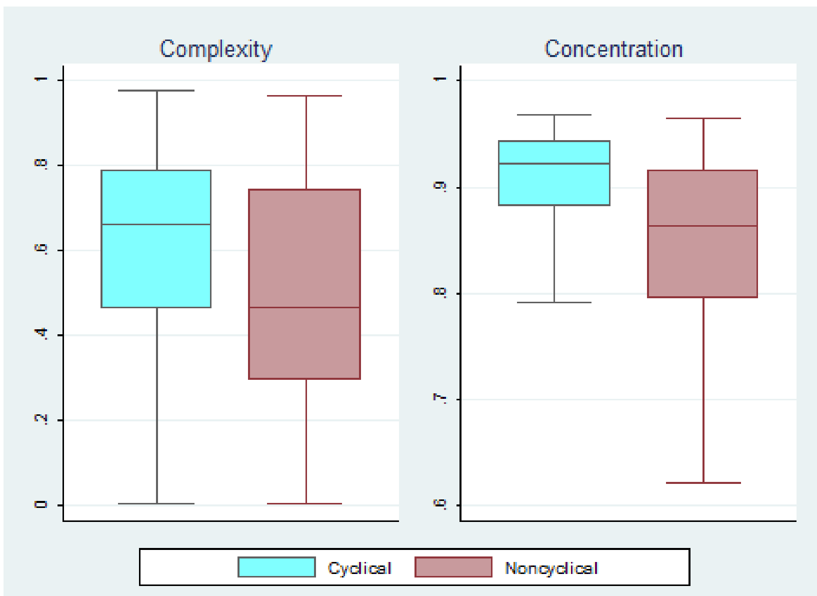
Compositional effects and durability play a role in explaining trade collapses in the aftermath of recessions and financial crises (see, e.g., Levchenko et al., 2010, and Eaton et al., 2011). This is because international trade is intensive in certain product categories, such as investment and durable consumption goods, that are more sensitive to cyclical fluctuations than other products. We next evaluate the robustness of our baseline results to the inclusion of interaction terms of the crisis indicators with dummy variables representing (loosely speaking) cyclical and noncyclical goods.

We construct two dummy variables using the mapping of 4-digit SIC industries to categories of final demand from Gomes et al. (2009). Our first dummy variable (labeled as "Cyclical (exc. NX)") takes the value of 1 (cyclical) if the industry is categorized by Gomes et al. (2009) as *durable consumption* or *investment*, and takes the value of 0 (noncyclical) if the industry is categorized as *nondurable consumption*, *government consumption and investment*, *consumption of services*, or *net exports*. Alternatively, a second dummy variable (labeled as "Cyclical (with NX)") is constructed in the same way except that the category *net exports* is included within the cyclical group.[23]

[23]Gomes et al's (2009) classification covers the majority of SIC-4 industries. However, to gain

To examine the relationship between contract dependence and cyclicality, we use our first dummy variable to split the sample of industries according to whether they are more or less cyclical. In Figure 2.3 we plot the distributions of the complexity and the concentration indices for each subsample. (Using the second dummy variable to split the sample yields relatively similar plots.)

Figure 2.3
Distribution of complexity index and concentration index, by cyclicality of industries



Notes: The box-and-whisker plots show the interquartile range, the median, and the most extreme values that are within 3/2 times the interquartile range of the 1st and 3rd quartiles. *Complexity* is the input relationship-specificity index from Nunn (2007). *Concentration* is the input concentration index from Levchenko (2007). Based on Gomes et al's (2009) classification of 4-digit SIC industries by final demand, our *cyclical* industries include durable consumption and investment goods; *noncyclical* industries include nondurable consumption, government consumption and investment, consumption of services, and net exports of goods and services.

It is visually apparent that cyclical industries tend to be more contractually vulnerable than noncyclical industries. This pattern is particularly strong if we

observations, whenever a certain SIC-4 industry is not categorized by them, we impose the category of final demand corresponding to the immediately higher level of aggregation (SIC-3).

observe the plots for the concentration index. More concretely, the median values of the complexity and the concentration indices are higher for cyclical industries than for noncyclical ones. The differences between medians are statistically significant at the 1% level according to the adjusted median chi-square and the Kruskal-Wallis tests.

The results with the first and the second dummy variable are displayed in Table 2.10 and Appendix Table B.4, respectively. We find that our baseline estimates are essentially unaffected when we account for the fact that crises can have a larger negative impact on more cyclical goods. If anything, the magnitudes of the interaction coefficients for the contractual-dependence measures decline slightly in absolute value. Moreover, when the coefficients on the interactions with the cyclicality dummies are statistically significant, recessions are found to disproportionately reduce trade in cyclical goods relative to noncyclical goods by 3 to 5 percentage points.[24] These results hold using either of our two dummy variables for cyclicality.

2.4.2.5 Controlling for product differentiation

As documented in Chapter 1, there is evidence that trade quantities of differentiated goods decline relatively more than those of nondifferentiated goods during recessions, even after controlling for features such as durability and end use. To rule out the possibility that our measures of contractual dependence may pick up the effect of product differentiation, we estimate a set of regressions that include

[24]This finding is consistent with previous evidence in the literature. For example, Abiad et al. (2014) report that the recent Great Trade Collapse caused an additional average drop of 10% in trade in capital and durable goods than in consumer nondurable goods.

Table 2.10
Effects of crises and contract vulnerability on trade across countries and industries: controlling for cyclicality. Dependent variable: Ln(bilateral sectoral imports)

	(1)	(2)	(3)
Recession alone	-0.054***	-0.055***	-0.057***
	(0.014)	(0.014)	(0.015)
Fin.Crisis alone	-0.011	-0.011	-0.004
	(0.021)	(0.021)	(0.022)
Recession & Fin.Crisis	-0.195***	-0.196***	-0.190***
	(0.032)	(0.032)	(0.033)
Recession alone × Complexity	-0.126***		
	(0.025)		
Fin.Crisis alone × Complexity	-0.037		
	(0.036)		
Recession & Fin.Crisis × Complexity	-0.223***		
	(0.037)		
Recession alone × Concentration		-0.264***	
		(0.045)	
Fin.Crisis alone × Concentration		-0.093	
		(0.063)	
Recession & Fin.Crisis × Concentration		-0.397***	
		(0.066)	
Recession alone × Uncollectible			-0.239***
			(0.053)
Fin.Crisis alone × Uncollectible			-0.075
			(0.081)
Recession & Fin.Crisis × Uncollectible			-0.419***
			(0.091)
Recession alone × Durables (exc. NX)	-0.028***	-0.029***	-0.047***
	(0.008)	(0.008)	(0.013)
Fin.Crisis alone × Durables (exc. NX)	0.009	0.009	0.008
	(0.009)	(0.009)	(0.013)
Recession & Fin.Crisis × Durables (exc. NX)	-0.003	-0.008	-0.010
	(0.012)	(0.012)	(0.019)
Observations	5,538,895	5,538,895	2,077,588
R-squared	0.897	0.897	0.903
Importer-Exporter-Product FE	YES	YES	YES
Year FE	YES	YES	YES

Notes: See notes to Table 2.5. Standard errors are clustered by destination-year, with ***,**,* respectively denoting significance at the 1%, 5% and 10% levels. The regressions include the same control variables as the baseline estimation (see Table 2.5).

additional interaction terms of the crisis indicators with a new dummy variable.

We construct a dummy variable (that we call "Differentiation") using the Rauch (1999) product classification. This variable takes the value of 1 if an industry is classified by Rauch as differentiated, and 0 if an industry is identified as nondifferentiated. The latter category includes goods traded in organized exchanges or with reference prices.[25]

Table 2.11 shows that recessions alone and recessions with financial crises generate a disproportional decline in the trade volumes of differentiated goods. According to our statistically significant results, the magnitude of this reduction is between 4% and 5%. Furthermore, the amplification effects induced by contractual vulnerability, captured by the interaction terms with our measures of contractual dependence, are moderately smaller relative to the baseline regressions. In column 2, we also note that these amplification effects are statistically significant for the case of financial crises alone, which stands in contrast to our baseline results.

2.4.2.6 Controlling for financial vulnerability, cyclicality, and product differentiation

Building on previous exercises, we perform a sensitivity check that simultaneously includes extra interaction terms in our baseline regressions. These variables aim to control for the potentially larger effect of crises on trade in industries with

[25]Since the Rauch (1999) classification is originally available at the SITC level, we previously map Rauch's categories into 4-digit SIC industries. To do so, we follow a similar procedure as the one employed in Chapter 1 for the classification of NAICS industries as differentiated and nondifferentiated (see Appendix A.1).

Table 2.11
Effects of crises and contract vulnerability on trade across countries and industries: controlling for product differentiation. Dependent variable: Ln(bilateral sectoral imports)

	(1)	(2)	(3)
Recession alone	-0.054***	-0.055***	-0.058***
	(0.014)	(0.014)	(0.015)
Fin.Crisis alone	-0.011	-0.011	-0.004
	(0.021)	(0.022)	(0.022)
Recession & Fin.Crisis	-0.195***	-0.196***	-0.191***
	(0.032)	(0.033)	(0.033)
Recession alone × Complexity	-0.123***		
	(0.023)		
Fin.Crisis alone × Complexity	-0.042		
	(0.029)		
Recession & Fin.Crisis × Complexity	-0.215***		
	(0.032)		
Recession alone × Concentration		-0.243***	
		(0.038)	
Fin.Crisis alone × Concentration		-0.103**	
		(0.048)	
Recession & Fin.Crisis × Concentration		-0.330***	
		(0.059)	
Recession alone × Uncollectible			-0.204***
			(0.047)
Fin.Crisis alone × Uncollectible			-0.060
			(0.072)
Recession & Fin.Crisis × Uncollectible			-0.368***
			(0.083)
Recession alone × Differentiation	-0.009	-0.019	-0.053***
	(0.010)	(0.012)	(0.015)
Fin.Crisis alone × Differentiation	0.008	0.007	-0.004
	(0.014)	(0.018)	(0.021)
Recession & Fin.Crisis × Differentiation	-0.009	-0.038*	-0.036
	(0.018)	(0.021)	(0.026)
Observations	5,538,895	5,538,895	2,077,588
R-squared	0.897	0.897	0.903
Importer-Exporter-Industry FE	YES	YES	YES
Year FE	YES	YES	YES

Notes: See notes to Table 2.5. Standard errors are clustered by destination-year, with ***,**,* respectively denoting significance at the 1%, 5% and 10% levels. The regressions include the same control variables as the baseline estimation (see Table 2.5).

higher working capital necessities (CashCycle), in cyclical industries (Cyclical (exc. NX)), and in differentiated goods industries (Differentiation).

Table 2.12 shows the results. We observe that the inclusion of the additional control variables somewhat reduces the magnitude (in absolute value), but not the statistical significance, of some of the key point estimates in our baseline regressions. On the other hand, the interactions between both the complexity index and the concentration index with financial crisis alone are now significant at the 10% level. We therefore find evidence that even without recessions, financial disruptions by themselves can lead to disproportionate declines in imports in more contractually vulnerable industries.

2.4.2.7 Time-variant multilateral trade resistance

By including exporter-importer-industry fixed effects, our baseline specification captures the time-invariant dimension of the multilateral trade resistance term. In another set of regressions, we capture the time-varying component of multilateral resistance under two alternative specifications. First, we incorporate exporter-industry and importer-industry fixed effects, letting them vary by 6-year periods. This method allows for the possibility that the time-variant multilateral resistance term be industry-specific. Second, we consider a specification that omits the year fixed effects but includes exporter-year and importer-year fixed effects. Under this methodology, the newly introduced fixed effects subsume the average impact of crises on industry imports, and hence we are no longer able to identify the parameter α_k.

Table 2.12
Effects of crises and contract vulnerability on trade across countries and industries: controlling for financial vulnerability, cyclicality, and differentiation. Dependent variable: Ln(bilateral sectoral imports)

	(1)	(2)	(3)
Recession alone	-0.054***	-0.055***	-0.058***
	(0.014)	(0.014)	(0.015)
Fin.Crisis alone	-0.011	-0.011	-0.003
	(0.022)	(0.022)	(0.022)
Recession & Fin.Crisis	-0.195***	-0.194***	-0.192***
	(0.032)	(0.032)	(0.033)
Recession alone × Complexity	-0.122***		
	(0.024)		
Fin.Crisis alone × Complexity	-0.050*		
	(0.029)		
Recession & Fin.Crisis × Complexity	-0.197***		
	(0.034)		
Recession alone × Concentration		-0.227***	
		(0.037)	
Fin.Crisis alone × Concentration		-0.109**	
		(0.043)	
Recession & Fin.Crisis × Concentration		-0.284***	
		(0.059)	
Recession alone × Uncollectible			-0.151***
			(0.049)
Fin.Crisis alone × Uncollectible			-0.093
			(0.067)
Recession & Fin.Crisis × Uncollectible			-0.320***
			(0.077)
Recession alone × CashCycle	0.005	-0.011	-0.012
	(0.011)	(0.011)	(0.015)
Fin.Crisis alone × CashCycle	0.010	0.004	0.034*
	(0.016)	(0.015)	(0.020)
Recession & Fin.Crisis × CashCycle	-0.030	-0.062***	-0.052*
	(0.021)	(0.020)	(0.027)
Recession alone × Durables (exc. NX)	-0.028***	-0.028***	-0.041***
	(0.007)	(0.007)	(0.012)
Fin.Crisis alone × Durables (exc. NX)	0.008	0.008	0.003
	(0.009)	(0.009)	(0.012)
Recession & Fin.Crisis × Durables (exc. NX)	-0.001	-0.001	0.001
	(0.012)	(0.012)	(0.020)
Recession alone × Differentiation	-0.008	-0.014	-0.045***
	(0.010)	(0.011)	(0.014)
Fin.Crisis alone × Differentiation	0.006	0.006	-0.013
	(0.015)	(0.018)	(0.021)
Recession & Fin.Crisis × Differentiation	-0.005	-0.023	-0.022
	(0.018)	(0.021)	(0.027)
Observations	5,538,895	5,538,895	2,077,588
R-squared	0.897	0.897	0.903
Importer-Exporter-Product FE	YES	YES	YES
Year FE	YES	YES	YES

Notes: See notes to Table 2.5. Standard errors are clustered by destination-year, with ***,**,* respectively denoting significance at the 1%, 5% and 10% levels. The regressions include the same control variables as the baseline estimation (see Table 2.5). CashCycle' is demeaned.

The results are summarized in Table 2.13. Columns 1-3 correspond to the first specification, and columns 4-6 refer to the second specification. We confirm the majority of our baseline findings, although some differences are worth noting. In particular, in columns 1-3 we show that financial crises alone generate a statistically significant average decline in imports (of around 5%). Additionally, as in some previous robustness checks, we find that the amplification effects of contractual vulnerability conditional on a financial crises alone are statistically significant (columns 1-2 and 4-6). Finally, under the first specification, the amplification effects associated with recessions alone are smaller than in the baseline model, as implied by the point estimates of β_1.

2.5 Conclusions

In this paper we provide evidence on a mechanism that has been ignored in the existing literature on crises and trade. We document empirically that when countries experience recessions and financial disruptions, their imports fall disproportionately in more contract-dependent industries. Put differently, contractual imperfections at the product level exacerbate the negative impact of crises on international trade. This mechanism operates on top of other relevant sources of heterogeneity across industries, such as financial dependence, degree of cyclicality, and product differentiation. Moreover, the estimated amplification effect of contractual vulnerability on sectoral imports appears to strengthen if the crisis country has weak rule of law.

We argue that these findings can be rationalized by two considerations. First,

Table 2.13

Effects of crises and contract vulnerability on trade across countries and industries: controlling for time-variant multilateral resistance. Dependent variable: Ln(bilateral sectoral imports)

	(1)	(2)	(3)	(4)	(5)	(6)
Recession alone	-0.041***	-0.041***	-0.043***			
	(0.013)	(0.013)	(0.013)			
Fin.Crisis alone	-0.052**	-0.052**	-0.046**			
	(0.020)	(0.020)	(0.020)			
Recession & Fin.Crisis	-0.189***	-0.190***	-0.178***			
	(0.032)	(0.032)	(0.032)			
Recession alone × Complexity	-0.041*			-0.146***		
	(0.024)			(0.007)		
Fin.Crisis alone × Complexity	-0.065**			-0.041***		
	(0.032)			(0.008)		
Recession & Fin.Crisis × Complexity	-0.204***			-0.250***		
	(0.038)			(0.011)		
Recession alone × Concentration		-0.103**			-0.280***	
		(0.041)			(0.014)	
Fin.Crisis alone × Concentration		-0.105*			-0.115***	
		(0.057)			(0.017)	
Recession & Fin.Crisis × Concentration		-0.298***			-0.406***	
		(0.066)			(0.023)	
Recession alone × Uncollectible			-0.107**			-0.324***
			(0.054)			(0.029)
Fin.Crisis alone × Uncollectible			-0.130			-0.097***
			(0.082)			(0.033)
Recession & Fin.Crisis × Uncollectible			-0.388***			-0.471***
			(0.091)			(0.045)
Observations	5,538,895	5,538,895	2,077,588	5,538,895	5,538,895	2,077,588
R-squared	0.905	0.905	0.911	0.901	0.901	0.907
Importer-Exporter-Industry FE	YES	YES	YES	YES	YES	YES
Importer-Industry FE (6-year periods)	YES	YES	YES	NO	NO	NO
Exporter-Industry FE (6-year periods)	YES	YES	YES	NO	NO	NO
Importer-Year FE	NO	NO	NO	YES	YES	YES
Exporter-Year FE	NO	NO	NO	YES	YES	YES
Year FE	YES	YES	YES	NO	NO	NO

Notes: See notes to Table 2.5. Robust standard errors (clustered by destination-year in columns 1-3), with ***,** ,* respectively denoting significance at the 1%, 5% and 10% levels. The regressions include the same control variables as the baseline estimation (see Table 2.5).

a large share of cross-border transactions rely on post-shipment payment. Second, the risk of default of importers is affected by macroeconomic conditions and worsens in industries which are more sensitive to the quality of contracting institutions.

Our quantitative findings may seem surprising in light of some considerations that could downplay the risk of default faced by exporters. First, although post-shipment terms are the most prevalent payment method in trade, other safer alternatives include cash in advance and bank-intermediated financing terms. Second, the use of intermediaries for exporting is relatively widespread. Third, exporters may purchase insurance against defaults based on insolvency. Fourth, for the substantial fraction of trade that occurs among related parties, contractual frictions are presumably less relevant. In spite of all of these arguments, our results robustly indicate that contract enforcement at the importing-country and industry level plays a significant role in shaping the response of trade to crises. We would require finer data to evaluate the importance of each of the aforementioned considerations. In addition, it would be interesting to disentangle the roles of the extensive margin and the intensive margin of trade in our story. We see our paper as a first step towards tackling those issues.

Appendix A: Appendix for Chapter 1

A.1 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 A.1), I use the 4-digit SITC to 10-digit HS (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% 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. For illustrative purposes, Table A.1 displays the complete classification of 4-digit NAICS industries. I checked that the results are not affected by using data on U.S. trade flows from other years (in particular, 2001 and 1996).

Table A.1
Classification of manufacturing industries by product differentiation at the 4-digit NAICS level

Code	Classific.	Industry Description
3111	N	Animal Foods
3112	N	Grain and Oilseed Milling Products
3113	U	Sugar and Confectionery Products
3114	U	Fruit and Vegetable Preserves and Specialty Foods
3115	N	Dairy Products
3116	N	Meat Products and Meat Packaging Products
3117	N	Seafood Products Prepared, Canned and Packaged
3118	D	Bakery and Tortilla Products
3119	D	Foods, NESOI
3121	N	Beverages
3122	N	Tobacco Products
3131	N	Fibers, Yarns, and Threads
3132	D	Fabrics
3133	D	Finished and Coated Textile Fabrics
3141	D	Textile Furnishings
3149	D	Other Textile Products
3151	D	Knit Apparel
3152	D	Apparel
3159	D	Apparel Accessories
3161	D	Leather and Hide Tanning
3162	D	Footwear
3169	D	Other Leather Products
3211	D	Sawmill and Wood Products
3212	N	Veneer, Plywood, and Engineered Wood Products
3219	D	Other Wood Products
3221	N	Pulp, Paper, and Paperboard Mill Products
3222	U	Converted Paper Products
3231	D	Printed Matter and Related Product, NESOI

Table A.1 (continued)

3241	U	Petroleum and Coal Products
3251	N	Basic Chemicals
3252	U	Resin, Synthetic Rubber, & Artificial & Synthetic Fibers & Filiment
3253	N	Pesticides, Fertilizers and Other Agricultural Chemicals
3254	D	Pharmaceuticals and Medicines
3255	D	Paints, Coatings, and Adhesives
3256	D	Soaps, Cleaning Compounds, and Toilet Preparations
3259	D	Other Chemical Products and Preparations
3261	D	Plastics Products
3262	D	Rubber Products
3271	D	Clay and Refractory Products
3272	D	Glass and Glass Products
3273	U	Cement and Concrete Products
3274	D	Lime and Gypsum Products
3279	D	Other Nonmetallic Mineral Products
3311	U	Iron and Steel and Ferroalloy
3312	U	Steel Products From Purchased Steel
3313	N	Alumina and Aluminum and Processing
3314	N	Nonferrous Metal (Except Aluminum) and Processing
3315	U	Foundries
3321	D	Crowns, Closures, Seals and Other Packing Accessories
3322	D	Cutlery and Handtools
3323	U	Architectural and Structural Metals
3324	U	Boilers, Tanks, and Shipping Containers
3325	D	Hardware
3326	U	Springs and Wire Products
3327	D	Bolts, Nuts, Screws, Rivets, Washers and Other Turned Products
3329	D	Other Fabricated Metal Products
3331	U	Agriculture and Construction Machinery
3332	D	Industrial Machinery
3333	D	Commercial and Service Industry Machinery
3334	D	Ventilation, Heating, Air-Conditioning, and Commercial Refrigeration Equipment
3335	D	Metalworking Machinery
3336	D	Engines, Turbines, and Power Transmission Equipment
3339	D	Other General Purpose Machinery
3341	U	Computer Equipment
3342	D	Communications Equipment
3343	D	Audio and Video Equipment
3344	D	Semiconductors and Other Electronic Components
3345	D	Navigational, Measuring, Electromedical, and Control Instruments
3346	D	Magnetic and Optical Media
3351	D	Electric Lighting Equipment
3352	D	Household Appliances and Miscellaneous Machines, NESOI
3353	D	Electrical Equipment
3359	D	Electrical Equipment and Components, NESOI
3361	D	Motor Vehicles
3362	D	Motor Vehicle Bodies and Trailers
3363	D	Motor Vehicle Parts
3364	U	Aerospace Products and Parts
3365	U	Railroad Rolling Stock
3366	D	Ships and Boats

Table A.1 (continued)

3369	D	Transportation Equipment, NESOI
3371	D	Household and Institutional Furniture and Kitchen Cabinets
3372	D	Office Furniture (Including Fixtures)
3379	D	Furniture Related Products, NESOI
3391	D	Medical Equipment and Supplies
3399	U	Miscellaneous Manufactured Commodities

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").

Construction of measures of sectoral trade prices and volumes. As explained in section 1.3.3, 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% 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 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% 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% 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% 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. I use the import share estimates for nondurable *major* PCE categories reported by McCulley (2011) for 2009. These major PCE categories are: 'Food and beverages purchased off-premises consumption', 'Clothing and footwear', 'Gasoline and other energy goods', and 'Other Nondurables'.

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.[1] Due to the lack of disaggregated information, I assume the same im-

[1]The NIPA tables for PCE by type of product published by the BEA imply the following

port shares for both the nondifferentiated and the differentiated content within each type of major product.

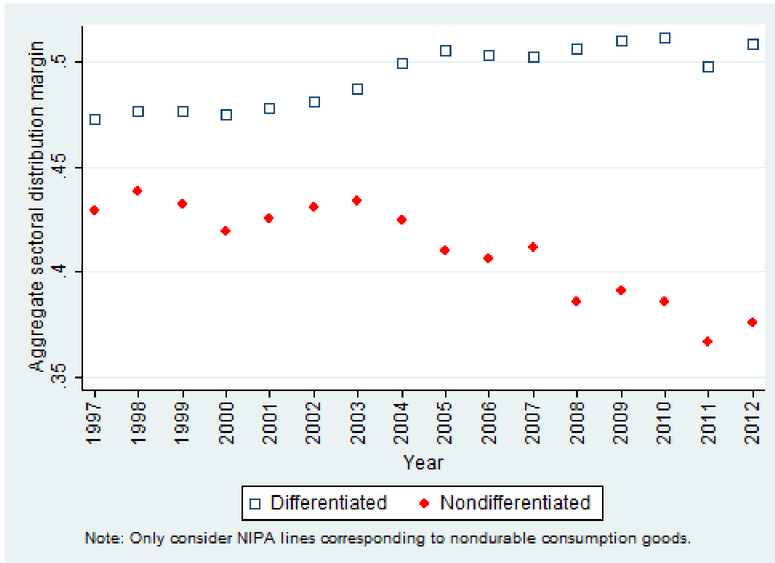
For each major PCE category, the relative weight of sector-s goods in the sectoral consumption basket is calculated as the ratio of the relative weight of sector-s goods in PCE, to the relative weight of the overall sector-s consumption basket in PCE. As part of these calculations I use the shares of PCE by major type of product reported by McCulley (2011).

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 A.1. For this exercise, the summary-level PCE categories corresponding to the following nondurable ma-

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.

Figure A.1
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.

jor 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 major products are coded as differentiated: 'Clothing and footwear', 'Pharmaceutical and other medical products', 'Household supplies', 'Personal care products', and 'Magazines, newspapers, and stationery'.[2]

[2]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

A.2 Model: market clearing conditions

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

$$L^z_{H,t} = \left(C^z_{H,t} + D^z_{H,t}\right) \int_0^1 \left(\frac{P^z_{H,t}(i)}{P^z_{H,t}}\right)^{-\xi} di,$$

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:

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

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

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

Finally, market clearing in the aggregate labor market requires:

$$L_{H,t} = \sum_{s=N,D} L_{Hs,t} + L^z_{H,t}.$$

An analogous set of goods and labor market clearing conditions hold for the Foreign country.

---

differentiated content, as revealed by my Rauch-type classification of PCE categories at the detail level.

A.3 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.[3] 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 A.2, respectively. To facilitate comparison with previous findings, the first four columns of this Table replicate Table 1.4.

In columns 5 and 6, we observe that regardless of the specific values assumed

[3]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 $\Gamma_D = 1.22$.

Table A.2

Summary measures of theoretical and empirical impulse responses to contractionary nominal spending shock: relative prices and quantities, Home country (%)

	Empirical (1)	Baseline model (2)	Constant markups (3)	Calvo param. Diff. (4)	Dist.Marg. Diff. (5)	Dist.Marg. Nondiff. (6)	Calvo param. Nondiff. (7)
Maximum response with expected sign (1 year after shock)							
Rel. PPI	-0.27	-0.19	-0.25	-0.03	-0.17	-0.19	-0.09
Rel. Production	-0.29	-0.32	-0.39	-0.04	-0.28	-0.31	-0.15
Rel Import price	-0.97	-0.19	-0.25	-0.03	-0.16	-0.18	-0.11
Rel Import volume	-0.38	-0.34	-0.43	-0.04	-0.30	-0.32	-0.18
Average response 1 year after shock							
Rel. PPI	-0.14	-0.16	-0.19	-0.02	-0.14	-0.15	-0.07
Rel. Production	-0.15	-0.28	-0.32	-0.02	-0.25	-0.27	-0.11
Rel Import price	-0.32	-0.16	-0.19	-0.02	-0.14	-0.15	-0.09
Rel Import volume	0.08	-0.30	-0.36	-0.03	-0.26	-0.28	-0.15
Response 6 months after shock							
Rel. PPI	-0.14	-0.16	-0.19	-0.02	-0.14	-0.15	-0.07
Rel. Production	-0.15	-0.28	-0.32	-0.02	-0.25	-0.27	-0.11
Rel Import price	-0.32	-0.16	-0.19	-0.02	-0.14	-0.15	-0.09
Rel Import volume	0.08	-0.30	-0.36	-0.03	-0.26	-0.28	-0.15

Notes: See notes to Table 1.4. Columns 1-4 replicate the columns from Table 1.4. 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.

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 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.

Appendix B: Appendix for Chapter 2

B.1 Data sources

Table B.1
Data sources

Variable	Source
Trade and country-level data	
Variable	Source
World export and import data	The Center for International Data
US export data	The Center for International Data
World export price index	IFS database
Real GDP in US Dollars	WDI database
Bilateral real exchange rate	Penn World Table 8.1
Free trade agreements	de Sousa (2012)
Bilateral geographic distance	CEPII distance database
Private credit to GDP ratio	Financial Development and Structure database
Rule of law	Worldwide Governance Indicators database
Banking crisis dates	Laeven and Valencia (2013)
Sovereign debt crisis dates	Laeven and Valencia (2013)
Industry data	
Variable	Source
Complexity index	Chor (2010) (based on Nunn, 2007)
Concentration index	Chor (2010) (based on Levchenko, 2007)
Collection Effectiveness Index	Credit Research Foundation
Cash conversion cycle	Compustat
Asset tangibility	Compustat (based on Braun, 2003)
Cyclicality	Durability classification by Gomes et al. (2009)
Differentiation	Product classification by Rauch (1999)

B.2 Data from the National Summary of Domestic Trade Receivables

The National Summary of Domestic Trade Receivables (NSDTR) data are in readable PDF format, so we first transcribe these files to machine-readable format. The NSDTR's Collection Effectiveness Index (CEI) is constructed as follows:

$$\text{CEI} = \frac{\text{Beginning total receiv.} + (\text{Quarterly credit sales}/3) \text{ - Ending total receiv.}}{\text{Beginning total receiv.} + (\text{Quarterly credit sales}/3) \text{ - Ending current receiv.}}$$

where:

'Beginning (Ending) total receiv.': Receivables balance at beginning (end) of 3-month period being reported. Considers all domestic open invoices and notes receivable, deferred billings or datings, past-due billings, credits, unapplied cash, suspense accounts, charge backs, invoice deductions, bankruptcies, claims, disputes, litigation and accounts placed for collections.

'Quarterly credit sales': Total invoiced receivable for the 3-month period reported. Includes freight, taxes, and containers.

'Ending current receiv.': Portion of receivables (domestic open accounts and notes) not yet due as of end of period according to terms, including datings and deferred items. We take median values across quarters by 4-digit SIC industry. On each quarter, the survey includes only industries that report a minimum of 3 responding firms. For more detailed information about the NSDTR, see `http://www.crfonline.org/surveys/surveys.asp`.

Table B.2 summarizes the 10 most and 10 least collection-effective industries.

Table B.2

Uncollectible index: the ten highest and lowest ranked industries

Best collection performance			**Worst collection performance**		
SIC	Description	1-CEI	SIC	Description	1-CEI
2873	Nitrogenous Fertilizers	0.02	3678	Electronic Connectors	0.28
2421	Sawmills and Planing Mills, General	0.02	3585	Refrigeration and Heating Equipment	0.30
2842	Speciality Clean., Polish., and Sanitary Preparations	0.02	3829	Measuring and Controlling Devices, NEC	0.30
2032	Canned Specialties	0.03	3569	General Industrial Machinery and Equipm., NEC	0.31
2676	Sanitary Paper Products	0.03	3069	Fabricated Rubber Products, NEC	0.32
2874	Phosphatic Fertilizers	0.03	3563	Air and Gas Compressors	0.33
2834	Pharmaceutical Preparations	0.03	3273	Ready-Mixed Concrete	0.33
3275	Gypsum Products	0.04	3444	Sheet Metal Work	0.39
2021	Creamery Butter	0.04	2399	Fabricated Textile Products, NEC	0.41
2023	Dry, Condensed, and Evaporated Dairy Products	0.04	3561	Pumps and Pumping Equipment	0.49

Notes: The uncollectible index is constructed as $1 - CEI$. See text for further details.

B.3 Sensitivity analysis: additional results

Table B.3
Effects of crises and contract vulnerability on trade across countries and industries: controlling for financial vulnerability.
Dependent variable: Ln(bilateral sectoral imports)

	(1)	(2)	(3)
Recession alone	-0.054***	-0.055***	-0.055***
	(0.014)	(0.014)	(0.015)
Fin.Crisis alone	-0.011	-0.011	-0.004
	(0.021)	(0.021)	(0.022)
Recession & Fin.Crisis	-0.195***	-0.195***	-0.189***
	(0.032)	(0.032)	(0.032)
Recession alone × Complexity	-0.127***		
	(0.021)		
Fin.Crisis alone × Complexity	-0.045		
	(0.032)		
Recession & Fin.Crisis × Complexity	-0.220***		
	(0.032)		
Recession alone × Concentration		-0.242***	
		(0.038)	
Fin.Crisis alone × Concentration		-0.091*	
		(0.054)	
Recession & Fin.Crisis × Concentration		-0.337***	
		(0.059)	
Recession alone × Uncollectible			-0.249***
			(0.051)
Fin.Crisis alone × Uncollectible			-0.066
			(0.079)
Recession & Fin.Crisis × Uncollectible			-0.394***
			(0.093)
Recession alone × TangAssets	0.012	0.081**	0.156***
	(0.026)	(0.033)	(0.044)
Fin.Crisis alone × TangAssets	-0.027	-0.005	-0.000
	(0.027)	(0.039)	(0.053)
Recession & Fin.Crisis × TangAssets	0.010	0.144***	0.140*
	(0.036)	(0.047)	(0.076)
Observations	5,538,895	5,538,895	2,077,588
R-squared	0.897	0.897	0.903
Importer-Exporter-Industry FE	YES	YES	YES
Year FE	YES	YES	YES

Notes: See notes to Table 2.5. Standard errors are clustered by destination-year, with ***,**,* respectively denoting significance at the 1%, 5% and 10% levels. The regressions include the same control variables as the baseline estimation (see Table 2.5). 'TangAssets' is demeaned.

Table B.4
Effects of crises and contract vulnerability on trade across countries and industries: controlling for cyclicality. Dependent variable: Ln(bilateral sectoral imports)

	(1)	(2)	(3)
Recession alone	-0.054***	-0.055***	-0.057***
	(0.014)	(0.014)	(0.015)
Fin.Crisis alone	-0.011	-0.011	-0.003
	(0.021)	(0.021)	(0.022)
Recession & Fin.Crisis	-0.195***	-0.196***	-0.190***
	(0.032)	(0.032)	(0.033)
Recession alone × Complexity	-0.125***		
	(0.025)		
Fin.Crisis alone × Complexity	-0.036		
	(0.036)		
Recession & Fin.Crisis × Complexity	-0.221***		
	(0.037)		
Recession alone × Concentration		-0.262***	
		(0.045)	
Fin.Crisis alone × Concentration		-0.092	
		(0.062)	
Recession & Fin.Crisis × Concentration		-0.391***	
		(0.066)	
Recession alone × Uncollectible			-0.249***
			(0.054)
Fin.Crisis alone × Uncollectible			-0.071
			(0.082)
Recession & Fin.Crisis × Uncollectible			-0.410***
			(0.094)
Recession alone × Durables (with NX)	-0.028***	-0.029***	-0.039***
	(0.007)	(0.007)	(0.011)
Fin.Crisis alone × Durables (with NX)	0.005	0.005	0.003
	(0.008)	(0.008)	(0.012)
Recession & Fin.Crisis × Durables (with NX)	-0.012	-0.017*	-0.016
	(0.010)	(0.010)	(0.016)
Observations	5,538,895	5,538,895	2,077,588
R-squared	0.897	0.897	0.903
Importer-Exporter-Product FE	YES	YES	YES
Year FE	YES	YES	YES

Notes: See notes to Table 2.5. Standard errors are clustered by destination-year, with ***,**,* respectively denoting significance at the 1%, 5% and 10% levels. The regressions include the same control variables as the baseline estimation (see Table 2.5).

Bibliography

Abiad, A., Mishra, P., and Topalova, P. (2014). How Does Trade Evolve in the Aftermath of Financial Crises? *IMF Economic Review*, 62:2, 213-247.

Ahn, J. (2014). Understanding Trade Finance: Theory and Evidence from Transaction-Level Data. Unpublished. International Monetary Fund.

Alessandria, G., Kaboski, J., and Midrigan, V. (2010). The Great Trade Collapse of 2008-09: An Inventory Adjustment? *IMF Economic Review*, 58:2, 254-294.

Amiti, M., Itskhoki, O., and Konings, J. (2014). Importers, Exporters, and Exchange Rate Disconnect. *American Economic Review*, 104:7, 1942-1978.

Anderson, J., and van Wincoop, E. (2003). Gravity And Gravitas: A Solution To The Border Puzzle. *American Economic Review*, 93:1, 170-192.

Antràs, P. (2015). Global Production. Firms, Contracts and Trade Structure. Unpublished. Harvard University.

Antràs, P., and Foley, F. (forthcoming). Poultry in Motion: A Study of International Trade Finance Practices. *Journal of Political Economy*.

Arouri, M., Jawadi, F., and Nguyen, D.K. (2013). What Can We Tell about Monetary Policy Synchronization and Interdependence over the 2007-2009 Global Financial Crisis? *Journal of Macroeconomics*, 36, 175-187.

Asmundson, I., Dorsey, T., Khachatryan, A., Niculcea, I., and Saito, M. (2011). Trade and Trade Finance in the 2008-09 Financial Crisis. IMF Working Paper 11/16, International Monetary Fund.

Auboin, M., and Engemann, M. (2014). Testing the Trade Credit and Trade Link: Evidence from Data on Export Credit Insurance. *Review of World Economics*, 150:4, 715-743.

Balke, Nathan., and Wynne, M. (2007). The Relative Price Effects of Monetary Shocks. *Journal of Macroeconomics*, 29:1, 19-36.

Barsky, R., House, C., and Kimball, M. (2007). Sticky-Price Models and Durable Goods. *American Economic Review*, 97:3, 984-998.

Baumeister, C., Liu, P., and Mumtaz, H. (2013). Changes in the Effects of Monetary Policy on Disaggregate Price Dynamics. *Journal of Economic Dynamics and Control*, 37:3, 543-560.

Beck, T., Demirguc-Kunt, A., and Levine, R. (2000). A New Database on Financial Development and Structure. *World Bank Economic Review*, 14, 597-605.

Bems, R., Johnson, R., and Yi, K. (2013). The Great Trade Collapse. *Annual Review of Economics*, 5.

Bergin, P. and Corsetti, G. (2014). International Competitiveness and Monetary Policy. Unpublished. UC Davis and Cambridge University.

Berkowitz, D., Moenius, J., and Pistor, K. (2006). Trade, Law, and Product Complexity. *The Review of Economics and Statistics.*, 88:2, 363-373.

Berman, N., de Sousa, J., Martin, P., and Mayer, T. (2012). Time to Ship during Financial Crises. NBER WP 18274.

Bils, M. and Klenow, P. (2004). Some Evidence on the Importance of Sticky Prices. *Journal of Political Economy*, 112:5, 947-985.

Bils, M., Klenow, P., and Kryvtsov, O. (2003). Sticky Prices and Monetary Policy Shocks. *Federal Reserve Bank of Minneapolis Quarterly Review*, 27:1, 2-9.

Boissay, F. and Gropp, R. (2013). Payment Defaults and Interfirm Liquidity Provision. *Review of Finance*, 17:6, 1853-1894.

Boivin, J., Giannoni, M., and Mihov, I. (2009). Sticky Prices and Monetary Policy: Evidence from Disaggregated U.S. Data. *American Economic Review*, 99:1, 350-384.

Bouakez, H. and Normandin, M. (2010). Fluctuations in the Foreign Exchange Market: How Important are Monetary Policy Shocks? *Journal of International Economics*, 81:1, 139-153.

Bouakez, H., Cardia, E., and Ruge-Murcia, F. (2009). The Transmission of Monetary Policy in a Multisector Economy. *International Economic Review*, 50:4, 1243-1266.

Bouakez, H., Cardia, E., and Ruge-Murcia, F. (2014). Sectoral Price Rigidity and Aggregate Dynamics. *European Economic Review*, 65, 1-22.

Braun, M. (2003). Financial Contractibility and Asset Hardness. Unpublished. UCLA.

Braun, M. and Larrain, B. (2005). Finance and the Business Cycle: International, Inter-Industry Evidence. *Journal of Finance*, 60:3, 1097-1128.

Bricongne, J., Fontagne, L., Gaulier, G., Taglioni, D., and Vicard, V. (2012). Firms and the global crisis: French exports in the turmoil. *Journal of International Economics*, 87:1, 134-146.

Broda, C. and Weinstein, D. (2006). Globalization And The Gains From Variety. *The Quarterly Journal of Economics*, 121:2, 541-585.

Burstein, A. and Gopinath, G. (2014). International Prices and Exchange Rates. In: Gopinath, G., Helpman, E., and Rogoff, K. (eds.), *Handbook of International Economics*, 4th ed.

Burstein, A., Neves, J., and Rebelo, S. (2003). Distribution Costs and Real Exchange Rate Dynamics during Exchange-Rate-Based Stabilizations. *Journal of Monetary Economics*, 50:6, 1189-1214.

Calvo, G. (1983). Staggered Prices in a Utility-Maximizing Framework. *Journal of Monetary Economics*, 12:3, 383-398.

Carvalho, C. (2006). Heterogeneity in Price Stickiness and the Real Effects of Monetary Shocks. *Frontiers in Macroeconomics*, 2:1, 1-56.

Carvalho, C. and Lee, J. (2011). Sectoral Price Facts in a Sticky-Price Model. Staff Reports 495, Federal Reserve Bank of New York.

Carvalho, C. and Nechio, F. (2011). Aggregation and the PPP Puzzle in a Sticky-Price Model. *American Economic Review*, 101:6, 2391-2424.

Chari, V., Kehoe, P., and McGrattan, E. (2000). Sticky Price Models of the Business Cycle: Can the Contract Multiplier Solve the Persistence Problem? *Econometrica*, 68:5, 1151-1180.

Chor, D. (2010). Unpacking Sources of Comparative Advantage: A Quantitative Approach. *Journal of International Economics*, 82:2, 152-167.

Chor, D. and Manova, K. (2012). Off the Cliff and Back? Credit Conditions and International Trade during the Global Financial Crisis. *Journal of International Economics* , 87, 117-133.

Christiano, L.J., Eichenbaum, M., Evans, C. (1999). Monetary Policy Shocks: What Have We Learned and to What End? In: Taylor, J.B., Woodford, M. (eds.), Handbook of Macroeconomics, Vol. 1A, pp. 65-148.

Corsetti, G. and Dedola, L. (2005). A Macroeconomic Model of International Price Discrimination. *Journal of International Economics*, 67:1, 129-155.

Corsetti, G., Dedola, L., and Leduc, S. (2008). International Risk Sharing and the Transmission of Productivity Shocks. *Review of Economic Studies*, 75:2, 443-473.

Corsetti, G. and Pesenti, P. (2009). The Simple Geometry of Transmission and Stabilization in Closed and Open Economies. In: Clarida, R. and Giavazzi, F. (eds.), *NBER International Seminar on Macroeconomics 2007*, NBER.

Cravino, J. (2014). Exchange Rates, Aggregate Productivity and the Currency of Invoicing of International Trade. Unpublished. University of Michigan.

Credit Research Foundation. *National Summary of Domestic Trade Receivables*, various issues.

Cuñat, A. and Melitz, M. (2012). Volatility, Labor Market Flexibility, and The Pattern of Comparative Advantage. *Journal of the European Economic Association*, 10:2, 225-254.

De Loecker, J., Goldberg, P., Khandelwal, A., and Pavcnik, N. (2012). Prices, Markups and Trade Reform. NBER WP 17925.

Demir, B. and Javorcik, B. (2014). Grin and Bear It: Producer-Financed Exports from an Emerging Market. Unpublished. Bilkent University and University of Oxford.

de Sousa, J. (2012). The Currency Union Effect on Trade is Decreasing Over Time. *Economic Letters*, 117:3, 917-920.

Devereux, M., Shi, K., and Xu, J. (2007). Global Monetary Policy under a Dollar Standard. *Journal of International Economics*, 71:1, 113-132.

Eaton, J., Kortum, S., Neiman, B., and Romalis, J. (2011). Trade and the Global Recession. NBER WP 16666.

Eichenbaum, M., and Evans, C. (1995). Some Empirical Evidence on the Effects of Monetary Policy Shocks on Exchange Rates. *Quarterly Journal of Economics*, 110:4, 1975-1010.

Erceg, C., and Levin, A. (2006). Optimal Monetary Policy with Durable Consumption Goods. *Journal of Monetary Economics*, 53:7, 1341-

Feenstra, R., Inklaar, R., and Timmer, M. (forthcoming). The Next Generation of the Penn World Table. *American Economic Review*.

Feenstra, R., Lipsey, R., Deng, H., Ma, A., and Mo, H. (2005). World Trade Flows: 1962-2000. NBER WP 11040.

Feenstra, R., Romalis, J., and Schott, P. (2002). U.S. Imports, Exports and Tariff Data, 1989 to 2001. NBER WP 9387.

Fitzgerald, D. and Haller, S. (2014). Pricing-to-Market: Evidence From Plant-Level Prices. *Review of Economic Studies*, 81:2, 761-786.

Goldberg, L., and Campa, J. (2010). The Sensitivity of CPI to Exchange Rates: Distribution Margins, Imported Inputs, and Trade Exposure. *Review of Economics and Statistics*, 92:2, 392-407.

Goldberg, P. and Hellerstein, R.(2008). A Structural Approach to Explaining Incomplete Exchange Rate Pass-through and Pricing-To-Market. *American Economic Review Papers and Proceedings*, 98:2, 423-429.

Gomes, J., Kogan, L., and Yogo, M. (2009). Durability of Output and Expected Stock Returns. *Journal of Political Economy*, 117:5, 941-986.

Gopinath, G. and Itskhoki, O. (2010). Frequency of Price Adjustment and Pass-through. *Quarterly Journal of Economics*, 125:2, 675-727.

Gopinath, G. and Itskhoki, O. (2011). In Search of Real Rigidities. In: Acemoglu, D., Rogoff, K., and Woodford, M. (eds.), *NBER Macroeconomics Annual 2010*, Vol. 25.

Gopinath, G., Itskhoki, O., and Rigobon, R. (2010). Currency Choice and Exchange Rate Pass-through. *American Economic Review*, 100:1, 304-336.

Gopinath, G., Itskhoki, O., and Neiman, B. (2012). Trade Prices and the Global Trade Collapse of 2008-2009. *IMF Economic Review*, 60:3, 303-328.

Gopinath, G. and Rigobon, R. (2008). Sticky Borders. *Quarterly Journal of Economics*, 123:2, 531-575.

Gordon, R. (1990). What Is New-Keynesian Economics? *Journal of Economic Literature*, XXVIII, 1115-1174.

Grossman, S. and Hart O. (1986). The Costs and Benefits of Ownership: A Theory of Vertical and Lateral Integration. *Journal of Political Economy*, XCIV, 691-719.

Haddad, M., Harrison, A., and Hausman, C. (2010). Decomposing the Great Trade Collapse: Products, Prices, and Quantities in the 2008-2009 Crisis. NBER WP 16253.

Hart, O., and Moore, J. (1999). Foundations of Incomplete Contracts. *Review of Economic Studies*, 66:1, 115-138.

Hoefele, A., Schmidt-Eisenlohr, T., and Yu, Z. (2013). Payment Choice in International Trade: Theory and Evidence from Cross-Country Firm Level Data. CESifo WP 4350.

Jacobson, T., Linde, J., and Roszbach, K. (2013). Firm Default and Aggregate Fluctuations. *Journal of the European Economic Association*, 11:4, 945-972.

Kaufmann, D., Kraay, A., and Mastruzzi, M. (2010). The Worldwide Governance Indicators: A Summary of Methodology, Data and Analytical Issues. World Bank Policy Research Working Paper No.5430.

Kaufmann, D. and Lein, S. (2013). Sticky Prices or Rational Inattention – What Can We Learn from Sectoral Price Data? *European Economic Review*, 64, 384-394.

Kim, S. (2001). International transmission of U.S. monetary policy shocks: Evidence from VARs. *Journal of Monetary Economics* , 48:2, 339-372.

Krishna, P. and Levchenko, A. (2012). Comparative Advantage, Complexity, and Volatility. *Journal of Economic Behavior and Organization*, 94, 314-329.

Laeven, L. and Valencia, F. (2013). Systemic Banking Crises Database. *IMF Economic Review*, 61:2, 225-270.

Levchenko, A. (2007). Institutional Quality and International Trade. *Review of Economic Studies*, 74:3, 791-819.

Levchenko, A., Lewis, L., and Tesar, L. (2010). The Collapse of International Trade During the 2008-2009 Crisis: In Search of the Smoking Gun. *IMF Economic Review*, 58:2, 214-253.

Manova, K. (2013). Credit Constraints, Heterogeneous Firms and International Trade. *Review of Economic Studies*, 80:2, 711-744.

McCulley, C. (2011). Trends in Consumer Spending and Personal Saving, 1959-2009. *Survey of Current Business*, Bureau of Economic Analysis, June.

Mendoza, E. (1991). Real Business Cycles in a Small Open Economy. *American Economic Review*, 81:4, 797-818.

Mora, J. and Powers, W. (2011). Global Perspectives in the Decline of Trade Finance. In: Chauffour, J. and Malouche, M. (editors), *Trade Finance during the Great Trade Collapse*, The World Bank.

Mussa, M. (1986). Nominal Exchange Rate Regimes and the Behavior of Real Exchange Rates: Evidence and Implications. *Carnegie-Rochester Conference Series on Public Policy*, 25:1, 117-214.

Nadiri, M. I. (1969). The Determinants of Trade Credit in the U.S. Total Manufacturing Sector. *Econometrica*, 37:3, 408-423.

Nakamura, E. and Steinsson, J. (2008). Five Facts About Prices: A Reevaluation of Menu Cost Models. *Quarterly Journal of Economics*, 123:4, 1415-1464.

Neiman, B. (2011). A State-Dependent Model of Intermediate Goods Pricing. *Journal of International Economics*, 85:1, 1-13.

Nunn, N. (2007). Relationship-Specificity, Incomplete Contracts, and the Pattern of Trade. *The Quarterly Journal of Economics*, 122:2, 569-600.

Nunn, A. and Trefler, D. (2014). Domestic Institutions as a Source of Comparative Advantage. In: Gopinath, G., Helpman, E., and Rogoff, K. (eds.), *Handbook of International Economics*, 4th ed.

Peneva, E. (2011). Some Evidence on Factor Intensity and Price Rigidity. *Journal of Economic Dynamics and Control*, 35:10, 1652-1658.

Raddatz, C. (2006). Liquidity Needs and Vulnerability to Financial Underdevelopmen. *Journal of Financial Economics*, 80:3, 677-722.

Rauch, J. (1999). Networks versus Markets in International Trade. *Journal of International Economics*, 48:1, 7-35.

Rajan, R. and Zingales, L. (1998). Financial Dependence and Growth. *American Economic Review*, 88:3, 559-586.

Ravn, M. and Uhlig, H. (2002). On Adjusting the HP-Filter for the Frequency of Observations. *The Review of Economics and Statistics*, 84:2, 371-375.

Rodrik, D. (2000). How Far Will International Economic Integration Go?. *Journal of Economic Perspectives*, 14:1, 177-186.

Schmidt-Eisenlohr, T. (2013). Towards a Theory of Trade Finance. *Journal of International Economics*, 91:1, 96-112.

Scotti, C. (2011). A Bivariate Model of Federal Reserve and ECB Main Policy Rates. *International Journal of Central Banking*, September.

Seiden, Martin H. (1964). *The Quality of Trade Credit*. National Bureau of Economic Research.

Stockman, A. and Tesar, L. (1995). Tastes and Technology in a Two-Country Model of the Business Cycle: Explaining International Comovements. *American Economic Review*, 85:1, 168-185.

Williamson, O. (1979). Transaction Cost Economics: The Governance of Contractual Relations. *Journal of Law and Economics*, XXII, 233-261.

Williamson, O. (1985). *The Economic Institutions of Capitalism*. The Free Press.

Wu, C. and Xia, F. (2014). Measuring the Macroeconomic Impact of Monetary Policy at the Zero Lower Bound. NBER WP 20117.