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VOLATILITY AND PASS-THROUGH

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ABSTRACT

The response of inflation to nominal shocks varies across time: we use confidential BLS micro data to show that there is a robust positive relationship between exchange rate pass-through and the dispersion of item-level price changes. Furthermore, we show that time-variation in price change dispersion is both large and countercyclical so that ignoring microeconomic dispersion leads to large time-varying bias when estimating pass-through. Why does pass-through vary with microeconomic dispersion? We estimate a quantitative price-setting model with various sources of heterogeneity in order to interpret our “model-free” empirical results and to better understand the economic forces that shape the response of prices to cost shocks at a moment in time. We conclude that strategic complementarities, which have received wide attention for their role in shaping average pass-through, also appear to vary dramatically across time. In contrast, we find no evidence supporting the "uncertainty" shock literature, which posits time-variation in underlying cost shocks as an explanation for time-variation in dispersion.

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1 Introduction

Understanding the nominal transmission mechanism is one of the most important and challenging questions in macroeconomics. This question is difficult to answer because we do not observe monetary shocks, and we do not observe the economic primitives that theory implies are important for shaping responses to these shocks. An important set of recent papers has used micro price-setting data in the open economy environment to address these challenges: exchange rate movements provide large and observable shocks to nominal costs, and micro data allows for identification of structural models that cannot be achieved with aggregate data.

These papers have shown that various firm-level characteristics are important for shaping the response of prices to exchange rate shocks. In particular, there is growing evidence that strategic complementarities are crucial for matching the low level of exchange rate pass-through observed in the data. However, these papers have largely focused on understanding what shapes the low level of pass-through that is observed on average across time.

In this paper we have two main contributions: 1) We show that there is a strong positive empirical relationship between the dispersion of item-level price changes and pass-through in the economy. Furthermore, price change dispersion fluctuates significantly over time which leads to systematic time-series variation in pass-through. For example, constant pass-through specifications are overstated by 50 percent during the mid-1990s and understated by 200 percent during the Great Recession. Since monetary policy is not conducted at random, we believe that documenting time-variation in pass-through is at least as important for policy making as measuring average pass-through in the data.

2) We use this empirical evidence to infer the nature of shocks across firms and time. In particular, we build a quantitative price-setting model with various sources of heterogeneity and formally estimate their importance in explaining the comovement between dispersion and pass-through. We find that variation in strategic complementarities (and thus firms' desired responses to cost shocks) drives our results. Current consensus in the exchange rate pass-through literature assigns a large role to such strategic complementarities in explaining why pass-through is low on average. Our results suggest that this mechanism is also crucial for driving variation in both pass-through and price change dispersion. In contrast, we find no evidence supporting the "uncertainty" shock literature, which posits time-variation in underlying cost shocks as an explanation for time-variation in dispersion.

Why do we study the relationship between price change dispersion and pass-through? Many recent papers have argued that the distribution of price changes can have important implications for aggregate price flexibility. For example, Midrigan (2011) and Alvarez, Le Bihan, and Lippi (2014) show that theory assigns a large role to the distribution of price changes in shaping the average response of inflation to nominal shocks. Vavra (2014) argues that increases in the dispersion of price changes during recessions should lead to increases in aggregate price flexibility. However, the existing literature drawing a connection between the distribution of price changes and price flexibility is largely theoretical. In this paper, we provide model free evidence that time-variation in price change dispersion strongly predicts time-variation in import price flexibility, and we provide additional insight into the forces that shape this relationship.

The first half of the paper concentrates on establishing our empirical fact that item-level price change dispersion strongly predicts exchange rate pass-through. We show that this positive relationship holds both at the item-level (cross-section) and at the month-level (time-series). That is: 1) Individual items with high dispersion of price changes have greater exchange rate pass-through. 2) During times when the cross-sectional dispersion of price changes across items is high, there is greater exchange rate pass-through.

We show that this fact is extremely robust. Broadly speaking, our results are robust to controlling for "micro" observables, "macro" observables and to a variety of measurement error related concerns. In addition, the fact that we find similar relationships in the cross-section and the time-series (as predicted by theory) provides some reassurance that the time-series relationship is not a spurious one driven by the failure to control for some confounding variable. Finally, we argue that our results cannot be explained by a mechanical relationship whereby higher exchange rate pass-through leads to higher item-level dispersion of price changes.

The set of controls in our empirical analysis is quite substantial. Amongst "micro" controls, we show that the positive relationship between pass-through and dispersion is not driven by differences in items' frequency of adjustment, frequency of product substitution, country of import, degree of differentiation, sector of production, the currency of the country from which the item is imported or the volatility of that currency. Our results are also robust to various macro controls. They continue to hold when excluding the 2008 recession, which is a large outlier in price change dispersion. In addition, controlling for the average frequency of price changes and product substitution, real GDP growth, exchange rate volatility, seasonality and long-run time trends does not alter our conclusions. We also run a battery of placebo checks, outlier robust regressions, alternative specifications for pass-through, and use a variety of simulations using our structural model to argue that our fact is not driven by small sample concerns or other measurement error related issues.

While the first half of the paper is purely empirical, the second half of the paper provides a structural interpretation of our new fact. There are many theoretical channels that can affect pass-through and price change dispersion. For example, the open-economy macro literature has argued that variable markups arising from various strategic complementarities can generate empirically relevant heterogeneity across firms in exchange rate pass-through by changing their desired responses to shocks. Up to now, this literature has focused on explaining the low average level of pass-through together with cross-sectional heterogeneity in pass-through across firms. However, if the strength of the variable markup channel varies across time, then firms' desired responses to shocks will also vary across time. Thus, simultaneous increases in dispersion and exchange rate pass-through could be explained by an increase in firms' responsiveness to shocks of constant magnitude.

In addition to this channel, it is also clear that time-variation in price change dispersion can arise from time-variation in the volatility of idiosyncratic shocks. A large literature including Bloom (2009), Bloom, Floetotto, Jaimovich, Saporta-Eksten, and Terry (2012), and Vavra (2014) have used such shocks to explain countercyclical dispersion of various economic variables such as firm growth, employment, and price changes. However, the role of such shocks for generating time-variation in measured exchange rate pass-through has heretofore been unexplored.

While the above channels have received the most attention in the literature, they are far from exhaustive. Differences in the size of adjustment costs, the sensitivity of firms' costs to exchange rates, changes in the volatility of exchange rates, or changes in the "commonality" of aggregate shocks all have the potential to affect price change dispersion and exchange rate pass-through.

We assess the importance of these various channels by using indirect inference to formally estimate the role of each in a quantitative menu cost model which builds on Gopinath and Itskhoki (2010) and Burstein and Gopinath (2013). While the basic modeling framework is intentionally standard, our emphasis is not on the model itself but is instead on what it can tell us about the underlying nature of firm-level shocks. The open economy environment is key to answering this question as it provides the source of identification in the model. For example, with only data on outcomes and no data on observable shocks, changes in "uncertainty" and changes in "responsiveness" are observationally equivalent: an increase in either implies an increase in the dispersion of observable firm decisions. However, these two theories have very different implications for how firms will respond to observable exchange rate shocks.

Using our identified quantitative model, we show that heterogeneous import shares, menu costs, changes in the volatility of exchange rates, and shocks to the "commonality" of aggregate shocks are all unable to explain our results.[1] We also find no evidence in support of the uncertainty shock channel. In contrast, our estimated model assigns an extremely important role to variation in firms' desired markups arising from strategic complementarities.[2]

Thus our model allows us to interpret our empirical results as evidence of time-varying markup adjustment. This provides further support for the type of mechanism emphasized by Gopinath and Itskhoki (2010). Their paper argues that variable markups (which generate heterogeneity in responsiveness) are necessary to explain cross-sectional differences in what they call long-run pass-through (LRPT). We extend the evidence for variable markups in two dimensions: 1) In addition to explaining cross-sectional patterns in LRPT, we show that variable markups help explain our new cross-sectional patterns in what Gopinath and Itskhoki (2010) call medium-run pass-through (MRPT). 2) More importantly, variable markups help explain time-series variation in MRPT.

MRPT measures the fraction of exchange rate movements passed-through into an item's price after one price adjustment whereas LRPT captures pass-through over an item's entire life. While much of the literature has moved towards the use of LRPT rather than MRPT, we focus on MRPT in this paper because it is the relevant pass-through concept for measuring time-varying price flexibility at business cycle frequencies. MRPT directly measures how shocks today are passed into price changes today whereas LRPT measures how shocks will transmit to prices potentially years

[1] We also model various sources of measurement error and show that these cannot explain our empirical patterns.

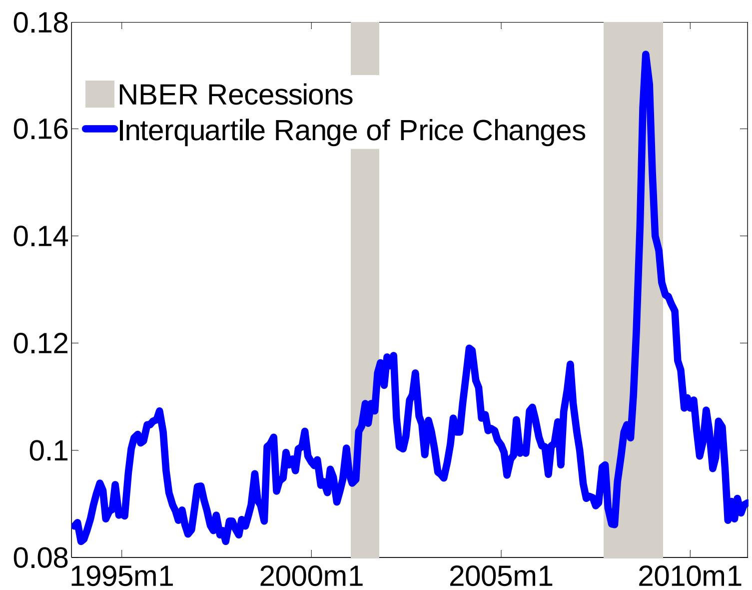
[2] For concreteness, we generate variable responsiveness channel using variable markups arising from Kimball demand, but this is largely for illustrative purposes. Other forms of strategic-complementarity have similar implications for responsiveness, price change dispersion, and exchange rate pass-through.

into the future. By construction, empirical measures of LRPT are not useful for measuring time-varying aggregate dynamics since LRPT is fixed across time for each item. Nonetheless, our focus on MRPT presents additional empirical challenges because potential biases induced by sampling error or mismeasured timing of price changes are much larger for MRPT than they are for LRPT. We address these measurement error issues explicitly in both our empirical and modeling sections.[3]

Despite its relevance for time-varying price flexibility, there is little work documenting heterogeneity in MRPT across items or time. Gopinath, Itskhoki, and Rigobon (2010) and Neiman (2010) are important exceptions. The former paper documents a strong relationship between dollar/non-dollar invoicing and MRPT while the latter paper demonstrates a robust relationship between whether transactions takes place within or between firms and MRPT. However, we believe our paper is the first to document that MRPT varies dramatically across time.[4]

While understanding why the average level of pass-through is so low is an important step in understanding the nominal transmission mechanism, we believe that documenting and understanding time-variation in pass-through is at least as important. If nominal stimulus occurred at random, then the average degree of price flexibility in the economy would be an important summary statistic. However, we show that import price flexibility rises dramatically with price dispersion during exactly the types of episodes when nominal stimulus is most likely. Figure 1 shows that the Great Recession was associated with an extremely large increase in microeconomic price churning.

Figure 1: Cross-Sectional IQR of Price Changes Across Time



[3]In our empirical section, we measure medium horizon pass-through using various alternative specifications and show that our results are unaffected. In the appendix to our modeling section we explicitly introduce various sources of measurement error and show they cannot explain our empirical results.

[4]In addition, our results show that even when restricting to dollar invoiced, inter-firm transactions there remains large cross-sectional variation in MRPT.

In addition to this model-free result, our evidence strongly supports alternative channels over "uncertainty shocks" as an explanation for countercyclical dispersion. In particular, time-variation in the competitive structure of markets or other shocks that induce time-variation in firm responsiveness appear more consistent with our data. Understanding the sources of time-varying dispersion is important for policy design as policies to reduce uncertainty almost certainly differ from policies designed to alter market structure and firms' responsiveness.

The remainder of the paper proceeds as follows: Section 2 contains our main empirical findings. Section 3 discusses the implications for time-varying pass-through. Section 4 lays out a basic flexible price model that demonstrates how primitives can potentially generate a positive relationship between pass-through to price change variance. Section 5 estimates a quantitative structural model to argue that variation in responsiveness best explains the data, and Section 6 concludes.

2 Empirical Results

2.1 Data Description

In this section we describe the data employed in this study. We use confidential micro data on import prices collected by the Bureau of Labor Statistics for the period 1994-2011. This data is collected on a monthly basis and contains information on import prices for very detailed items over time. This data set has previously been used by Clausing (2001), Gopinath and Rigobon (2008), Gopinath, Itskhoki, and Rigobon (2010), Gopinath and Itskhoki (2010), Berger, Faust, Rogers, and Steverson (2012) and Neiman (2010). Below, we provide a brief description of how the data is collected. See the IPP (Import Price Program) Data Collection Manual for a much more detailed description (U.S. Department of Labor, 2005).

The target universe of the price index consists of all items purchased from abroad by U.S. residents (imports). Sampling is designed at the entry level item (ELI) level, which in most cases corresponds to a 10-digit harmonized trade code. Within the 10-digit harmonized code, an item is defined as a unique combination of a firm, a product and the country from which a product is shipped. An example of the type of item that forms our unit of observation is "Lot # 12345, Brand X Black Mary Jane, Quick On/Quick Off Mary Jane, for girls, ankle height upper, TPR synthetic outsole, fabric insole, Tricot Lining, PU uppers, Velcro Strap."[5]

Price data are collected monthly for approximately 10,000 imported items. The BLS prefers to collect prices that, in the case of imports, are "free on board" (fob) at the foreign port of exportation before insurance, freight or duty are added. The prices collected are net (exclusive) of duties. Almost 90% of U.S. imports have a reported price in dollars.

The BLS collects prices using voluntary confidential surveys, which are usually conducted by mail. A reporting company is contacted for the transaction price on a monthly basis. Respondents are then asked to provide prices for actual transactions that occur as close as possible to the first day of the month. Typically a company specifies if a price has been contracted and the period for

[5] This example is taken from Gopinath and Rigobon (2008).

which it is contracted, including specifying the months in which actual trade will take place. For the periods when the price is contracted, the BLS will use the contracted price without contacting the firm directly and also enter a flag for whether the good is traded or not in those months.[6]

As with all surveys, there are some concerns about the quality of the IPP data. However, there are many reasons to believe that reporting is accurate. First, the BLS is very concerned with data quality and so works hard to make sure that the burden on the participating firms is not high. In the first step of data collection, a BLS agent negotiates with the company over the number of price quotes that the company is comfortable reporting so as not to place undue burden on the firm. The BLS also has a policy of contacting a respondent if the reported price has not changed or the item has not traded for 12 months. This quality check helps reduce misreporting. Second, Gopinath and Rigobon (2008) uses the Anthrax scare of 2001, which forced the IPP to conduct interviews by phone, as a natural experiment. They found almost no difference in reported price setting around these months, which helps reduce concerns about misreporting.

Nonetheless, in the appendix we explore the robustness of our quantitative results to four types of measurement error: sampling error in the price collection process, errors in reporting the correct size of the price change, unreported price changes, and variation in shipping lags of goods. We find that all of our conclusions are robust to various assumptions about the magnitude of these errors.

We focus on a subset of the data that satisfies the following three criteria: 1) We restrict attention to market transactions and exclude intrafirm transactions, as we are interested in price-setting driven mainly by market forces. 2) We require that a good have at least one price adjustment during its life. This is because the goal of the analysis is to relate the standard deviation of price changes to the price pass-through of the item and this requires observing at least one price change. This is the same sample restriction used by Gopinath and Itskhoki (2010) in their study of frequency and exchange rate pass-through. 3) We restrict attention to dollar-priced imports using all countries and all products, excluding petroleum. We restrict attention to dollar-priced items, so as to focus on the relationship between dispersion and MRPT after removing variation due to currency choice. The previous literature (Gopinath, Itskhoki, and Rigobon (2010)) has shown there are large differences in MRPT across goods which are invoiced in different currencies, but the vast majority of products in the database are invoiced in dollars. Our benchmark results include all countries and all products excluding petroleum so as to include the broadest possible sample.

Throughout the paper and appendices, we show that all of our results are robust to various alternative sample selections. In particular, we have explored a variety of subsamples and found that our results obtain for individual countries, for different mixes of products, and when restricting the analysis to items with several price changes.

[6]According to Gopinath and Rigobon (2008), the BLS contacted 87% of the items at least once every 3 months, with 45% of the items contacted on a monthly basis. 100% of the items are contacted at least once a year.

2.2 Baseline Dispersion Results

2.2.1 Measuring Dispersion and Pass-through

We measure price change dispersion using two distinct but related empirical objects. First, we construct a measure of "item-level" dispersion. For each item j we define item-level dispersion as $DI_j = disp(\Delta p_{i,t}|i = j)$. That is, we calculate the dispersion of all non-zero price changes for item j across time. Since individual items typically have a small number of price changes, we measure item-level dispersion using the standard deviation of that item's price changes.

The second measure of dispersion we construct is "month-level" dispersion. For each month k we define month level-dispersion as $DM_k = disp(\Delta p_{i,t}|t = k)$. To calculate month-level dispersion, we fix a particular month and then calculate the dispersion of price changes across all items in that month.[7] Since there are thousands of price changes each month, we can calculate various different measures of dispersion including the standard deviation and interquartile range of price changes.

Summarizing our two measures of dispersion, "item-level" dispersion is calculated using a single item but all time-periods while "month-level" dispersion is calculated using all items but a single time-period. Since item-level dispersion varies across items rather than time, we refer to "cross-sectional" differences in item-level dispersion. Similarly, since month-level dispersion varies across time-periods rather than items, we refer to "time-series" variation in month-level dispersion.

We measure exchange rate pass-through using micro price data in a standard way. In particular, we focus on what Gopinath, Itskhoki, and Rigobon (2010) refer to as a medium-run pass-through (MRPT) and estimate the following regression on adjusting prices:

$$\Delta p_{i,t} = \beta \Delta_c e_t + Z'_{i,t}\gamma + \epsilon_{i,t} \tag{1}$$

Here, $\Delta p_{i,t}$ is item i's log price change, $\Delta_c e_t$ is the cumulative change in the bilateral exchange rate since item i's last price change, and $Z'_{i,t}$ is a vector of item and country level controls.[8] We estimate this regression with country and sector fixed effects.[9] The coefficient β measures the fraction of cumulated exchange rate movements "passed-through" to an item's price when adjusting.

We estimate (1) conditional on price adjustment. This implies that our pass-through estimates are not contaminated by nominal rigidities. If we did not condition on adjustment, pass-through would be very close to zero and dominated by variation in the frequency of adjustment. More importantly, conditioning on adjustment allows us to identify the presence of strategic complementarities since they affect how firms want to respond to shocks in the absence of nominal rigidities.

[7] Similar results obtain if we calculate month-level dispersion only within sectors.

[8] As usual, there are some concerns about interpreting exchange rate movements as exogenous, which is one reason for including controls for macro conditions. In addition, we are mainly interested in the relative ranking of pass-through across firms and time-periods rather than the absolute level, so endogeneity is less of a concern. Finally, our monthly data means we are identifying off of high frequency variation in exchange rate movements, which are hard to relate to anything observable.

[9] The sector fixed effects are at the primary strata lower (PSL) level, defined by the BLS as either the 2 or 4-digit harmonized tariff code. The other baseline controls are U.S. GDP and CPI and foreign country CPI.

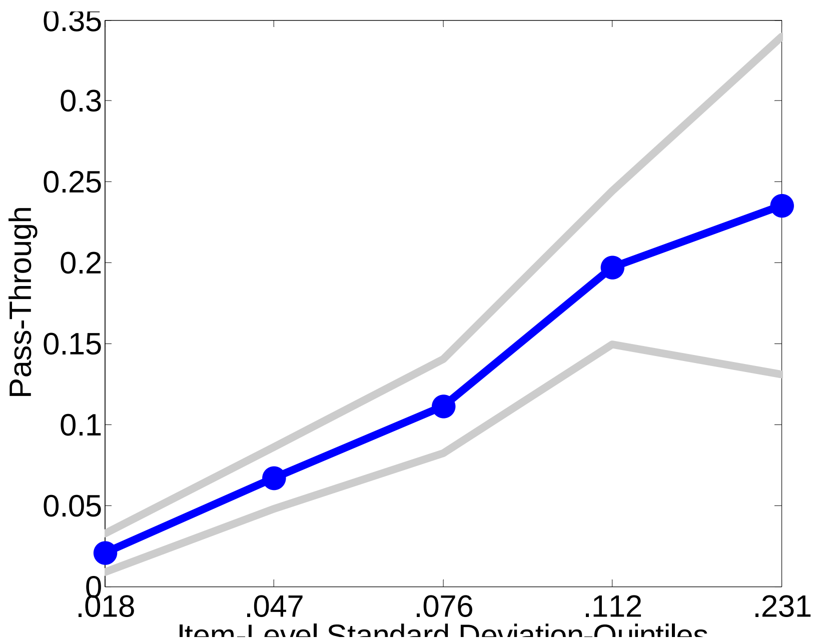
The results from estimating (1) for all price changes in our sample are shown in Table 1. Consistent with the previous literature, we find that average MRPT for dollar denominated items is low. Table 1 shows that when a price changes, it only passes through about 0.14% of a 1% change in the nominal exchange rate.[10]

2.2.2 Item-Level Dispersion Results

In this section, we document empirically that there is a strong relationship between medium-run pass-through and item-level price change dispersion.

Let $XSD_i = std(\Delta p_{i,t})$ be the standard deviation of item i's price changes (conditional on adjusting). To explore the relationship between MRPT and item-level dispersion, we split our sample into XSD_i quintiles and estimate equation (1) separately for each quintile. Figure 2 shows the baseline results with 95% confidence bands.Average pass-through increases from 2%

Figure 2: Medium-run pass-through across item-level XSD quintiles



in the lowest XSD_i quintile (with standard deviation equal to 0.016) to close to 25% for the highest quintile (with standard deviation equal to 0.213), an increase that is both economically and statistically significant. While we only show this baseline specification for a very broad set of countries and products and it includes no additional controls, in the following sections and appendices we show that this result is extremely robust and is not driven by other item-level features like the frequency of adjustment or degree of product differentiation.

[10]Existing papers typically find pass-through coefficients closer to 0.24. Our slightly lower number is due to the use of bilateral exchange rates, all countries rather than OECD countries, and the use of a moderately longer sample. Using trade-weighted currencies and OECD countries increases MRPT to close to 0.3.

2.2.3 Month-Level Dispersion Results

We now show that time periods characterized by greater price change dispersion also exhibit greater exchange rate pass-through. Our time-series evidence is of particular interest because it provides a direct test for time-varying price flexibility. Vavra (2014) argues for a positive relationship between price change dispersion and price flexibility but is unable to test for it empirically.

To test for a time-series relationship between price change dispersion and MRPT, we begin by calculating the cross-sectional interquartile range of price changes for each month in our sample. Then, just as we did for the item-level dispersion results, we sort our sample into quintiles by this month-level dispersion and calculate separate pass-through regressions in each quintile.

Figure 3: Medium-run pass-through across month-level IQR Quintiles

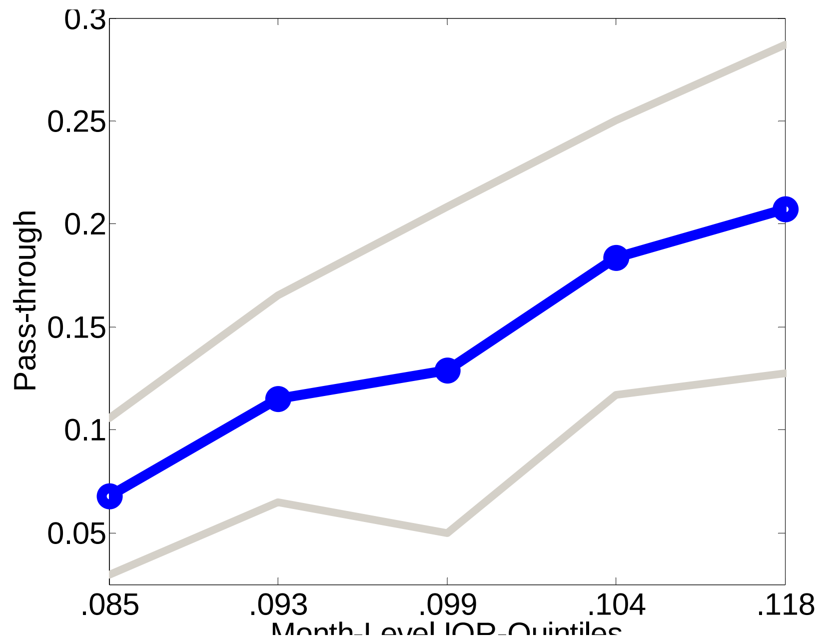


Figure 3 shows that pass-through more than triples from the lowest quintile of month-level dispersion to the highest quintile. This increase in pass-through is highly significant.[11] We assess relationship in more detail in the appendix and show that this same result obtains for various alternative measures of month-level dispersion, including the cross-sectional standard deviation of price changes as well as census level measures of dispersion computed in Bloom, Floetotto, Jaimovich, Saporta-Eksten, and Terry (2012). If we split the sample into deciles, we find even bigger variation across time, with pass-through in the highest dispersion months approaching 50%. In Section 3 we return to a detailed discussion of time-series variation in pass-through.

[11]Standard errors are larger than for the item-level relationships because our panel has a very large number of items but a much smaller number of time-periods.

2.3 Robustness of Item-Level Relationship

Figure 2 shows a strong positive relationship between item-level dispersion and MRPT, but it does not control for any other characteristics that vary across items. We now show that our item-level result is robust to a host of alternative explanations and controls. This is important, because when we move from empirics to theory, we will argue that our result reflects heterogeneity in fundamental economic primitives. In particular, we will argue that our result is driven by heterogeneity in the importance of strategic complementarities. In order to reach this strong conclusion, it is important to show that our result is not driven by other observable differences across items.

2.3.1 Ruling Out Mechanical Explanations

Looking at the MRPT specification in (1), there are two reasons to be concerned that the positive relationship between pass-through and dispersion in the cross-section might be picking up a mechanical relationship. First, note that the MRPT regression coefficient is equal to

$$\widehat{\beta} = \frac{cov(\Delta p_{i,t}, \Delta e_t)}{cov(\Delta e_{i,t}, \Delta e_t)} = \beta + cov(\epsilon_{i,t}, \Delta e_t).$$

where β is the desired response of prices to exchange rate movements. Taking the variance of both sides of (1) gives an expression for the variance of prices across time:

$$var\left(\Delta p_{i,t}\right) = \beta^2 var\left(\Delta e_t\right) + var\left(\epsilon_{i,t}\right) + 2\beta cov(\epsilon_{i,t}, \Delta e_t). \tag{2}$$

Thus, there are two mechanical channels by which increases in $\widehat{\beta}$ can lead to increases in the item-level dispersion of price changes: 1) If items differ in their β (perhaps due to heterogeneous sensitivity of costs to exchange rates), then there should be a positive relationship between β and $var\left(\Delta p_{i,t}\right)$. 2) If $cov(\epsilon_{i,t}, \Delta e_{i,t})$ differs across items, then this will induce a similar mechanical relationship. Since the number of price changes per item is small, this sample covariance can vary even though the true population covariance is zero.

While these are valid concerns, they do not explain our empirical result. The intuition is straightforward: to match the empirical variance of price changes, the variance of idiosyncratic shocks must be two orders of magnitude larger than that of exchange rate shocks. That is, $var\left(\epsilon_{i,t}\right) >> var\left(\Delta e_{i,t}\right)$. This implies that changing only β or $cov(\epsilon_{i,t}, \Delta e_{i,t})$ has negligible effects on $var\left(\Delta p_{i,t}\right)$.

Formally, we have empirical data on $var\left(\Delta p_{i,t}\right)$, $var\left(\Delta e_{i,t}\right)$, and β, so we can use equation 2 to measure the implied value of $var\left(\epsilon_{i,t}\right)$ under the null hypothesis that empirical differences across items can be solely explained by heterogeneity in β. Substituting empirical results from the data and using equation 2 yields $\beta = 0.144$, $var\left(\Delta e_{i,t}\right)$ =6.25e-4 and $var\left(\epsilon_{i,t}\right)$ =1.83e-2. Using these values for $var\left(\Delta e_{i,t}\right)$ and $var\left(\epsilon_{i,t}\right)$, if one varies β from 0.021 to 0.235 (as in the data) one can generate less than 0.1% of observed variation in dispersion.

A similar exercise also implies a negligible role for variation in $cov(\epsilon_{i,t}, \Delta e_{i,t})$ in explaining our result. Furthermore, in the following sections we show that our results hold even when restricting to items with a large number of price changes for which $var\left(\Delta p_{i,t}\right)$ and $cov(\epsilon_{i,t}, \Delta e_{i,t})$ will be less subject to measurement error. In addition, when we move to the structural interpretation of our results, we perform indirect inference on simulated data from the model that is consistent with sample sizes in the BLS, so any issues related to sampling error will be picked up in simulation.

Finally, it is also important to note that these mechanical explanations are only a concern for our item-level results and not for our month-level results since exchange rate movements are common to items within a given month. Thus, when calculating the variance of price changes across items instead of across time, $var\left(\Delta e_t\right)$ and $cov(\epsilon_{i,t}, \Delta e_t)$ are both identically zero so that the problematic terms are dropped from (2).

Thus, we conclude that the positive relationship between pass-through and dispersion is not driven by mechanical relationships whereby high pass-through necessarily implies high dispersion.

2.3.2 Item-Level Dispersion Interactions

Is the positive relationship between pass-through and dispersion driven by other observables? To explore this, we now run regressions on continuous measures of price change dispersion instead of the previous binned regressions. These more structured specifications allow us to include a variety of additional controls. Let the change in an item's price be given by:

$$\Delta p_{i,t} = \beta^{avg} \Delta_c e_{i,t} + \beta^{Vol} \left(Vol_i \times \Delta_c e_{i,t}\right) + \delta Vol_i + Z'_{i,t}\gamma + \epsilon_{i,t} \tag{3}$$

The coefficient β^{avg} captures the average pass-through in the sample and β^{Vol} gives the effect of item-level price change volatility on medium-run pass-through.[12] Results are shown in Table 2.

The first two rows show the results for our baseline sample which includes all countries and all items excluding petroleum products. Average exchange rate pass-through is 14%. β^{Vol} is significantly greater than zero, which means that items with higher price dispersion have higher MRPT. The price dispersion effect is economically meaningful: a one standard deviation increase in price dispersion implies a 37% (0.05/0.14) increase in average MRPT. Controlling for items' frequency of adjustment does not change this conclusion.[13] In addition, our results are not driven by a particular set of products or by a particular set of countries. Table 2 shows that restricting to OECD countries or manufactured items only strengthens the importance of item-level dispersion. In both subsamples, the price dispersion effect is economically and statistically significant.[14]

[12] In all specifications, the measure of item level price dispersion is the standard deviation of price changes (XSD) and robust standard errors are clustered by country and primary stratum lower (4 digit import type) pair. Measures of volatility in this and the following specifications are standardized to ease interpretation. This implies that β^{avg} is equal to the level of pass-through for an item with average dispersion, which is a natural measure of average pass-through.

[13] Also see the appendix where we show binned regressions by XSD and frequency for additional evidence that our results are not driven by heterogeneity in the frequency of adjustment.

[14] See the appendix for additional evidence. We find similar results when restricting to differentiated items. We can also allow for a sector specific interaction in the regression to fully control for sector composition. After controlling for sectoral differences, we continue to find a similar positive relationship between MRPT and item-level dispersion.

For brevity, we leave a variety of other item-level robustness checks for the appendix. In those results, we show that the positive relationship between MRPT and item-level dispersion holds for individual countries, using broad rather than bilateral exchange rates, is robust to controls for outliers and is not driven by small sample issues.

2.4 Robustness of Month-Level Relationship

In this section, we argue that the time-series relationship between month-level dispersion and pass-through is not driven by various other observables or confounding shocks. Since this time-series relationship is of direct consequence for policy making, it is particularly important to show that it is a deep relationship in the data rather than a spurious relationship driven by failing to control for some other observable. However, in contrast to the item-level data, which contain a limited set of covariates, we are able to show robustness results for a rich set of time-varying macro controls.

In order to control for other time-varying features of the data, we estimate an interaction specification between MRPT and month-level dispersion. More specifically, we run the regression

$$\Delta p_{i,t} = \beta^{avg}\Delta_c e_{i,t} + \beta^{IQR} IQR_t \times \Delta_c e_{i,t} + \lambda IQR_t + Z'_{i,t}\gamma + \epsilon_{i,t} \tag{4}$$

where IQR_t is the interquartile range of all (non-zero) price changes in month t and $Z'_{i,t}$ is the same vector of controls as in the cross-sectional regressions.[15] Table 3 shows that increasing IQR by one-standard deviation increases pass-through by 6 percentage points relative to an average pass-through of 14 percent. This positive relationship is highly significant, with a t-statistic of 7.01. We find similar effects when using the cross-sectional standard deviation instead of the interquartile range, as well as when restricting to OECD countries and manufactured items.

Using the continuous time-series specification also allows us to control for other things that might vary across time. Table 4 shows a number of these results. A long literature has argued that there may be secular changes in pass-through across time (e.g. Marazzi, Sheets, Vigfusson, Faust, Gagnon, Marquez, Martin, Reeve, and Rogers (2005)). If there are also trends in price change dispersion, our time-series results could be spurious. In addition, there may be seasonal patterns in price change dispersion and pass-through. The first robustness checks in Table 4 re-estimate regression 4 with a linear time-trend and month dummies. These controls do not affect our conclusions.

In addition to time-variation in price dispersion, both the frequency of adjustment and the frequency of product substitution vary across time.[16] However, controlling for frequency and product substitution does not affect the interaction between price change dispersion and pass-

[15] As in the cross-sectional regression we standardize all dispersion numbers to ease the interpretation of our results.

[16] Nakamura and Steinsson (2012) argue that missing price changes that occur at the time of product substitution can lead measured aggregate pass-through to be below true aggregate pass-through. Since our measure of MRPT conditions on price changes, the presence of product substitution is not directly relevant for our results. Nevertheless, we find that product substitution actually rises mildly with the dispersion of price changes. This means that the increase in pass-through we document probably understates the true increase in aggregate pass-through so that accounting for product substitution would, if anything, amplify our results.

through. Controlling for time-series variation in the volatility of exchange rates also does not affect our result.

In the following section, we will return to the implications of our results for time-varying pass-through. There we will show that both price dispersion and pass-through increased dramatically during the Great Recession in 2008. Are our time-series results then driven by this particularly dramatic period? Table 5 shows that they are not. Restricting our analysis to only data prior to 2008 changes none of our conclusions. More generally, price change dispersion tends to be countercyclical. Is our dispersion pass-through relationship just proxying for a recession pass-through relationship? Again, the answer is no. Table 4, shows that controlling for the state of the business cycle does not change our conclusions.

We also investigated the implications of these additional controls using various alternative samples and dispersion measures. In particular, using the standard deviation of price changes instead of the interquartile range and using different country and product mixes does not change the conclusion that these additional controls make little difference.[17]

2.5 Controls for Composition

While we have now shown that our empirical result holds under a variety of controls, these controls may not pick up compositional changes across time in our sample. In this section, we address various potential composition concerns with our results

2.5.1 2 Facts or 1 Fact?

First, is our item-level dispersion fact actually distinct from our month-level dispersion fact? Since items are periodically rotated out of the sample, we do not have a balanced panel. Thus, it is possible that the high dispersion time-periods in our data are driven by times when the sample contains items with unusually high price dispersion. To document that our two facts are indeed independent we first combine specifications (3) and (4) to allow for separate effects of cross-item and cross-month dispersion. That is, we estimate

$$\Delta p_{i,t} = \beta^{avg}\Delta_c e_{i,t} + \beta^{Vol}\left(XSD_i \times \Delta_c e_{i,t}\right) + \delta XSD_i + \beta^{IQR} IQR_t \times \Delta_c e_{i,t} + \lambda IQR_t + Z'_{i,t}\gamma + \epsilon_{i,t} \quad (5)$$

where XSD_i is the standard deviation of item i's price changes and IQR_t is the interquartile range of all price changes in month t. Table 6 shows that both the cross-item effects captured by XSD_i and the cross-month effects captured by IQR_t are highly significant. This conclusion is again robust to a variety of additional controls.

In addition to this double-interaction, the appendix shows results for a binned regression as in Figures (2) and (3). We split individual items into quintiles by their item-level dispersion of price changes, and then within each item-level quintile we run a time-series regression to estimate the

[17] In the interest of brevity we do not report these results, but all permutations of the above controls and subsamples are available upon request.

effect of month-level dispersion. Unlike the specification in (5) this "double-binned" regression does not impose linear effects of dispersion and allows the effect of controls to vary across bins. Nevertheless, we again find that both cross-item and cross-month dispersion effects are highly significant. Thus, simple changes in sample composition cannot jointly explain both facts.

2.5.2 Within or Between Sector Phenomena

In general, the positive relationship we observe between pass-through and month-level dispersion could be driven by either movements in dispersion across sectors or dispersion within a sector. While we believe either explanation would be interesting, they would have different implications for models. To address this, we decompose the month-level variance of price changes into a between and within-sector component: $VAR(dp_{i,t}) = VAR(dp_{i,t}^{\text{within sector}}) + VAR(dp_t^{\text{between sector}})$. We then separately interact pass-through with both between and within sector variance:

$$\Delta p_{i,t} = \beta^{avg}\Delta_c e_{i,t} + \beta^{VAR\_W} VAR\_W_t \times \Delta_c e_{i,t} + \beta^{VAR\_B} VAR\_B_t \times \Delta_c e_{i,t} + Z'_{i,t}\gamma + \epsilon_{i,t}$$

Table 7 displays results using both 2-digit and 4-digit sector definitions. For 2-digit sectors, only within-sector variance is significant while for 4-digit sectors both within and between sector variance are significant. We next perform a formal variance decomposition to describe how much of the variation in MRPT is accounted for by within vs. between sector changes in dispersion. The within-sector contribution is given by

$$W = \frac{\left(\beta^{VAR\_W}\right)^2 V\_W}{\left(\beta^{VAR\_W}\right)^2 V\_W + \left(\beta^{VAR\_B}\right)^2 V\_B},$$

where V_W is the time-series variance of within-sector price change dispersion and V_B is the time-series variance of between-sector price change dispersion. Using this decomposition, within-sector variance accounts for 99% of the time-series variation in pass-through using 2-digit sectors and 51% using 4-digit sectors. Thus, even for fairly narrow sectors, the time-series relationship between month-level dispersion and pass-through seem to be largely a within-sector phenomenon.

2.5.3 Are the Items Changing Prices During High Dispersion Months Special?

Are the items that change prices during high dispersion periods the same as the items that change prices during low dispersion periods? We now show that even when restricted to a balanced panel there is a positive relationship between month-level dispersion and MRPT. This means that pass-through for the same products rises with dispersion so our results cannot be explained by a time-varying product mix. Unfortunately, the BLS periodically rotates products in and out of the sample, so it is not feasible to construct a balanced panel that spans the entire length of our data. However, we do have enough data to construct a balanced panel that spans the Great Recession in 2008, which is the most important episode in our sample.

We restrict our analysis to a balanced panel that is in the sample continuously from 2007-2009 and then estimate separate pass-through regressions for 2007, 2008, and 2009. As in the full-sample, month-level dispersion rises dramatically in 2008. Pass-through also exhibits large variation, rising from 0.07 in 2007 to 0.64 in 2008 and falling back to 0.22 in 2009. Time-series variation in dispersion and pass-through for the balanced panel is even stronger and more significant than for our baseline unbalanced specification.

2.5.4 Exchange Rate Appreciations Vs. Depreciations

The 2008 recession was also characterized by an appreciation of the U.S. dollar against most major currencies. Are our pass-through results sensitive to the sign of exchange rate movements? We have re-run all of our empirical specifications restricting our regressions solely to price changes where $\Delta_c e_{i,t}$ is always positive or always negative. We find that both our month-level and our item-level dispersion MRPT relationships remain highly significant even when restricting only to exchange rate movements of a particular sign.[18] Thus, our results cannot be explained by changes across time in whether the dollar is appreciating or depreciating.

2.6 Measurement Error Concerns

2.6.1 Alternative Pass-through Specifications

All results thus far have relied on MRPT specifications of the form in (1). This specification directly measures the extent to which exchange rate movements are passed into current prices, so it is a natural measure of time-varying price flexibility. Nevertheless, there are some potential concerns with this specification. First, it is important that the timing of price changes be well-measured since mismeasured timing of price changes will lead to attenuation bias.[19]

Second, there may be heterogeneity across items in the number of price changes necessary to fully capture pass-through. That is, some items may fully pass-through exchange rate movements with a single price change while other items may take several price changes to achieve the same pass-through. If this is the case, then estimating pass-through conditional on a single price adjustment may provide a distorted picture of cross-item price flexibility (although it should still capture the degree of price flexibility at a given point in time).

With these measurement concerns in mind, we have estimated several alternative pass-through specifications.[20] First, we calculate a "fixed horizon" pass-through specification for various pass-through horizons. For these specifications, we calculate $\Delta p_{i,t}^K = p_{i,t+K} - p_{i,t}$ and $\Delta e_{i,t}^K = e_{i,t+K} - e_{i,t}$ for fixed horizons K. We then rerun specification (5) using this new measure of price and exchange rate changes. Crucially, this alternative specification does not condition on price adjustment, so

[18] For brevity we do not report these results, but they are available upon request.

[19] Mismatched timing of price changes and exchange rate movements can occur, e.g., when there is a lag between the time an item is purchased and the time it ships. In the appendix we show that controlling for shipping method does not alter our conclusions.

[20] In addition, in the appendix, we simulate various sources of measurement error and show that these cannot explain our results.

an individual item may have between 0 and K price changes occurring between t and $t+K$. Thus, this specification reflects the full extent of an item's pass-through over a fixed horizon, whether that pass-through occurs through one or several price changes. This measure of pass-through is analogous to the measure of life-long pass-through used in Gopinath and Itskhoki (2010), but it uses a fixed horizon for calculating pass-through rather than the observed life of each item. This fixed horizon pass-through has many of the attractive features of life-long pass-through but still allows us to calculate cross-month variation in pass-through.

In addition to fixed horizon regressions, we can also run our baseline MRPT regression allowing for lagged exchange rate movements to matter for current price changes. That is, we estimate:

$$\begin{aligned}\Delta p_{i,t} &= \beta_1^{avg}\Delta_c e_{i,t} + \beta_1^{Vol}\left(XSD_i \times \Delta_c e_{i,t}\right) + \beta_1^{IQR} IQR_t \times \Delta_c e_{i,t} \\ &\quad \beta_2^{ave}\Delta_c e_{i,t-1} + \beta_2^{Vol}\left(XSD_i \times \Delta_c e_{i,t-1}\right) + \beta_2^{IQR} IQR_t \times \Delta_c e_{i,t-1} \\ &\quad +\delta XSD_i + \lambda IQR_t + Z_{i,t}'\gamma + \epsilon_{i,t}.\end{aligned}$$

Table 8 provides results for these alternative pass-through specifications. In all cases, increases in both item-level and month-level dispersion lead to economically large and statistically significant increases in pass-through. Thus, our results are not sensitive to a particular measure of exchange rate pass-through. In the appendix we also show that life-long pass-through is increasing in item-level dispersion. However, since life-long pass-through is only measured once for each item we cannot measure time-variation in life-long pass-through.

2.6.2 Small Samples

Given the limited time span of our panel data and low average frequency of price adjustment, it is important to verify that the relationship between pass-through and dispersion is not being driven by small sample issues. In the previous section, we argued that small samples can lead to biased pass-through estimates by affecting $cov(\epsilon_{i,t}, \Delta e_{i,t})$, since the sample covariance can differ from zero even if the population covariance does not. This is only a concern for item-level results and not month-level relationships, and we argued that it was unlikely to play a quantitatively important role in our results. However, it is also straightforward to address these concerns empirically.

In the appendix, we show that our item-level results are robust to restricting the sample to items which have at least 3 or at least 5 price changes. We also run placebo regressions to see if our results are driven by spurious small sample issues. In particular, we substitute a count of an item's price changes or its price observations in place of XSD. These placebo regressions show that our results are not driven by a correlation between measured dispersion and item sample sizes.

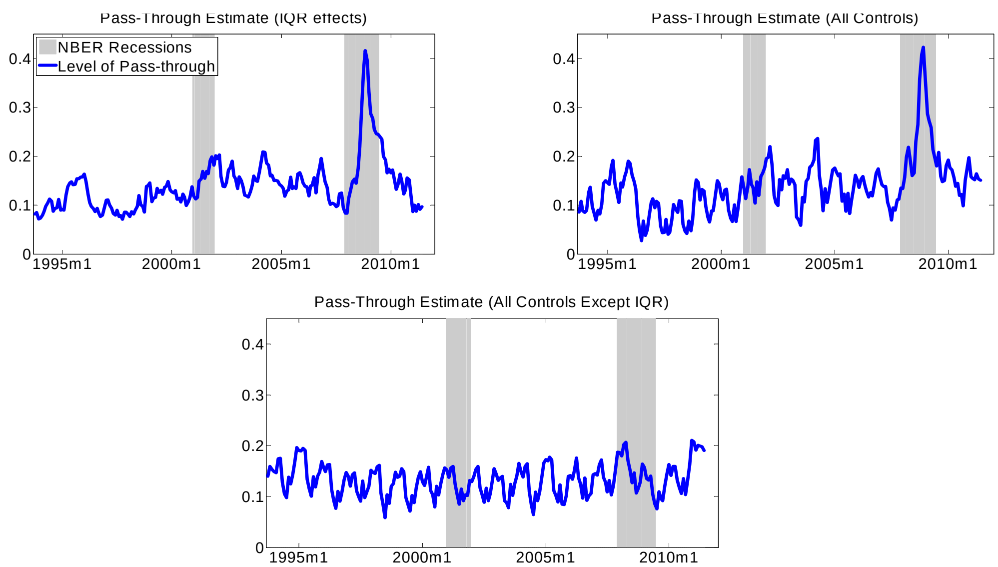
Finally, in our theoretical simulations, we simulate samples from the model that correspond to those in the BLS data. If small samples induced biases in our empirical specifications they should also appear in our regressions on simulated data, yet we find no such effects.

3 Time-Variation in Pass-Through

In the previous section we documented a robust link between exchange rate pass-through and microeconomic price change dispersion. In this section we show that this relationship generates economically significant variation in exchange rate pass-through at business cycle frequencies. In this sense our empirical results provide model-free evidence that looking at the microeconomic distribution of price changes is crucial for predicting inflation dynamics at a moment in time.

The results from the previous section allow us to construct implied time-series for exchange rate pass-through by multiplying observed variables by their estimated effects on exchange rate pass-through. For example, using Regression Specification 4 we estimate pass-through in each period t by computing $\widehat{MRPT}_t = \widehat{\beta}^{avg} + \widehat{\beta}^{IQR} IQR_t$.

Figure 4: Level of Exchange Rate Pass-through Across Time (Parametric Specifications)

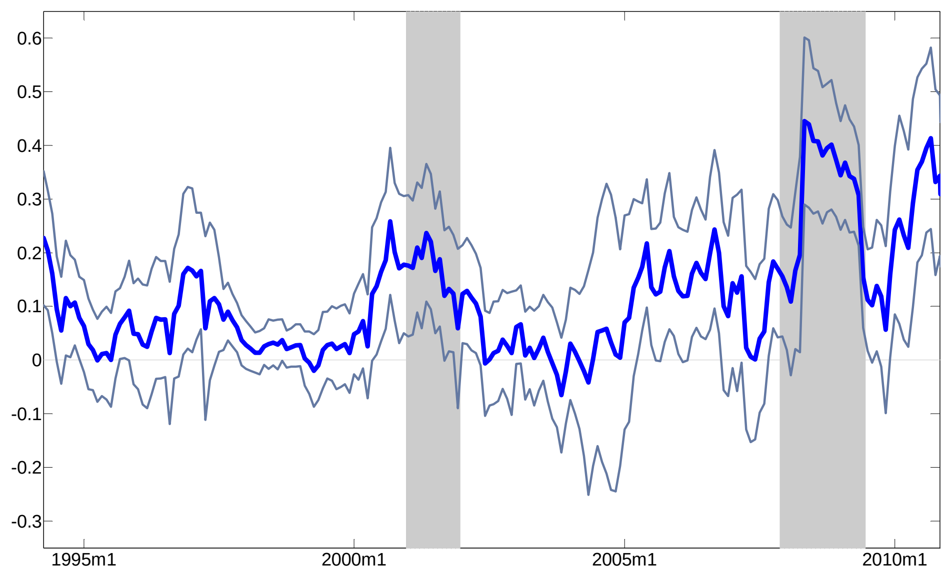


The identifying assumption in this specification is that the only thing that varies across time that affects exchange rate pass-through is IQR. The left hand panel of Figure 4 shows the resulting estimates for exchange rate pass-through under this specification. During the height of the Great Recession, this estimate of exchange rate pass-through rises to 44% relative to a low of approximately 7% during the late-1990s. The assumption that time-variation in exchange rate pass-through is solely driven by time-varying IQR is very strong, but it can easily be relaxed. In the right hand panel of Figure 4 we allow pass-through to vary with IQR, the frequency of adjustment, the frequency of product substitution, seasonal month dummies, and real GDP growth.

Allowing for these additional interactions does not change the conclusion that pass-through rose markedly during the 2008 recession. The main difference relative to the specification with only IQR is a large seasonal component. This can be seen most clearly in the bottom panel of Figure 4, which shows pass-through estimates for a specification with all controls except for IQR.

Essentially all the variation in pass-through at business cycle frequencies is captured by time-series variation in price change dispersion. Interestingly, there is large seasonality in pass-through, from a high of approximately 0.20 in December to a low of approximately 0.04 in June. Understanding these seasonal patterns is an interesting topic for future work, but the bottom line is that for understanding business cycle variation, looking at price change dispersion appears essential.[21]

Figure 5: Pass-through Estimate (Non-Parametric 12-month Rolling Window)



While the above results show that pass-through varies across time in a specification with a variety of controls, there is always concern that omitted variables might undo this time-series variation. That is, there may be additional variables we are not controlling for that affect pass-through and would undo the time-series variation we have found. We can assess this concern by allowing pass-through to vary across time non-parametrically. Ideally, we could re-estimate the baseline pass-through regression (1) with a full set of month dummies. However, small sample sizes make such regressions infeasible. Instead, we estimate the baseline regression using a rolling 12-month window. That is, our estimate of pass-through for period t is given by running a regression including only data from 6 months before and after period t:

$$\Delta p_{i,\tau} = \beta_t \Delta_c e_{i,\tau} + Z'_{i,\tau}\gamma + \epsilon_{i,\tau} \mid t-6 \leq \tau \leq t+6.$$

This allows us to construct a monthly measure of β_t that varies fully non-parametrically across time. Figure (5) shows the resulting estimates together with 90% confidence intervals. Overall the results are quite similar to the parametric specification, and again there is strong variation

[21] Seasonality is unlikely to be explained purely by a spike in the frequency of adjustment at the end of the year. This is because our measure of pass-through conditions on adjustment, so we are finding variation in how much adjusting prices respond to exchange rate movements over the season that are unlikely to be explained purely by frequency.

in pass-through at business-cycle frequencies. Running annual pass-through or 6-month pass-through regressions instead of using overlapping rolling windows produces very similar results.[22] This specification shows that being completely agnostic about what drives pass-through movements across time delivers quite similar results to our benchmark specifications.

Overall our results show that exchange rate pass-through varies dramatically across time, with microeconomic price change dispersion. This means that estimating average pass-through regressions without looking at micro data induces a significant time-varying bias, with pass-through substantially understated during periods of microeconomic churning. While a large literature tries to understand average pass-through and its implications for the nominal transmission mechanism, the above evidence shows that pass-through is not a single number and that concentrating on average pass-through may be misleading for how prices will respond to nominal shocks at a moment in time.

Beyond providing direct empirical evidence that the distribution of price changes matters for predicting pass-through, we now show that our empirical results provide additional identification that is useful for understanding the nature of heterogeneity and aggregate shocks in the economy.

4 Basic theoretical framework

4.1 Flexible price model

In this section we lay out a simple framework following Burstein and Gopinath (2013) that shows how economic primitives shape the relationship between exchange rate pass-through and price change dispersion. Consider the problem of a foreign firm selling items to U.S. importers. The firm has perfectly flexible prices, set in dollars. The optimal flexible price (in logs) of item i at the border is the sum of the gross markup (μ_i) and the dollar marginal cost ($mc_i(e, \eta_i)$) which depends on both the exchange rate (e) as well as an item-specific component, orthogonal to the exchange rate (η_i):[23]

$$p_i = \mu_i + mc_i(e, \eta_i). \tag{6}$$

Taking the total derivative of equation (6) gives:

$$\Delta p_i = -\Gamma_i(\Delta p_i - \Delta p) + \alpha_i \Delta e + \epsilon_i \tag{7}$$

where $\Gamma_i \equiv -\frac{\partial \mu_i}{\partial(\Delta p_i - \Delta p)}$ is the elasticity of a firm's optimal markup with respect to its relative price. We refer to this parameter as markup "responsiveness". It captures the classic pricing to market channel of Dornbusch (1987) and Krugman (1987), where firms may adjust markups in response to cost shocks, leading to incomplete pass-through. This channel implies a negative relationship between markups and relative prices, $p_i - p$, which Burstein and Gopinath (2013)

[22] Quarterly results (available from authors on request) are also similar although small sample sizes mean that the standard errors become extremely large and estimates are quite noisy.

[23] In the appendix, we consider a more general model which includes GE effects and scale-dependent marginal cost.

show is a robust implication of various strategic complementarities that generate incomplete pass-through. $\alpha_i \equiv \frac{\partial mc_i}{\partial e}$ is the partial elasticity of the dollar marginal cost to the exchange rate, e. We refer to this as the "import intensity" channel. Finally, $\epsilon_i = \Delta\eta_i$ captures the innovation of idiosyncratic marginal cost.[24] Rearranging this equation gives an explicit expression for the direct effect (that is when $\Delta p = 0$) of a change in the exchange rate on prices at the border:[25]

$$\frac{\Delta p_i}{\Delta e} = \frac{\alpha_i}{1 + \Gamma_i} \tag{8}$$

This expression for exchange rate pass-through is intuitive. The first factor affecting pass-through is the fraction of marginal cost denominated in dollars. If marginal cost is entirely denominated in dollars ($\alpha_i = 0$), then fluctuations in the exchange rate are irrelevant for the foreign firm's optimal dollar price and pass-through is zero. In general, exchange rate pass-through is increasing in import intensity.

The second factor affecting pass-through is the response of the foreign firm's optimal markup to changes in its relative price. If $\Gamma_i = 0$ (the CES case) the firm's optimal markup does not change as its price deviates from its competitors and pass-through is at its maximum. If $\Gamma_i > 0$, then as the price of the firm increases relative to its competitors, the elasticity of its demand rises, lowering its optimal markup. Similarly, when the firm's price is relatively low its optimal markup rises. Thus, if $\Gamma_i > 0$, the foreign firm will move its price less than one-for-one in response to cost shocks.

Since lowering Γ_i means that firms will be more responsive to all cost shocks, we refer to lowering Γ_i as increasing total "responsiveness". That is, firms with low Γ_i will respond strongly to both idiosyncratic shocks as well as exchange rate shocks. In contrast, firms with high α_i will respond more to exchange rate shocks but not to idiosyncratic cost shocks. We use the term responsiveness to differentiate general cost pass-through from exchange rate specific pass-through.

In addition to its implications for pass-through, we can also use equation (7) to show how α and Γ affect the variance of Δp_i. Solving for Δp_i and computing its variance gives:

$$var(\Delta p_i) = \left(\frac{\alpha_i}{1 + \Gamma_i}\right)^2 var(\Delta e_i) + \left(\frac{1}{1 + \Gamma_i}\right)^2 var(\epsilon_i), \tag{9}$$

where we have used the fact that exchange rate and idiosyncratic shocks are uncorrelated.

Intuitively, the variance of the firm's optimal price is larger if it faces a more volatile exchange rate or idiosyncratic shocks. In addition, using equation (9), it follows that factors that increase exchange rate pass-through ($\alpha_i \uparrow$, $\Gamma_i \downarrow$) also increase the variance of price changes. Moreover, using equation (9) it can be shown that for empirically relevant values of α_i and Γ_i, changing Γ_i has much

[24] Since we do not observe this shock, it is without loss of generality to normalize the price response to η to be one.
[25] We also set the innovation of the idiosyncratic shock to its average value (zero).

larger effects on price change variance than changing α_i.[26] The intuition, as previously discussed in Section 2.3.1, is that empirical estimates of $var(\epsilon_i)$ greatly exceed $var(\Delta e_i)$. In addition, α_i is typically small. This means that the first term contributes little to the overall variance of price changes, so changing its size also has little effect. In the quantitative modeling section, we show that this simple intuition survives in a realistic model. That is, the mechanical link between heterogeneity in α_i and heterogeneity in $var(\Delta p_i)$ is not empirically important.

4.2 Modeling Price Stickiness

Price stickiness is a pervasive feature of micro price data. For example, Gopinath and Rigabon (2008) find that the median price duration for imports to the U.S. is 10.6 months. More importantly, the price adjustment mechanism can have direct effects on measured pass-through. For example, in menu cost models, where price adjustment is endogenous, conditioning on price adjustment will induce a selection bias in MRPT estimates.[27] In contrast, this bias is absent in Calvo pricing models where price adjustment is exogenous.

To understand how the primitives of a menu cost model can affect measured pass-through, it is useful to examine our baseline MRPT regression shown in equation (1). By definition, the estimated MRPT regression coefficient is equal to:

$$\widehat{\beta} = \frac{cov(\Delta p_{i,t}, \Delta e_t)}{cov(\Delta e_{i,t}, \Delta e_t)} = \beta + \underbrace{cov(\epsilon_{i,t}, \Delta e_t)}_{\text{selection bias}}$$

where β is the "true" responsiveness of desired prices to exchange rate movements.[28] Menu cost models induce $cov(\epsilon_{i,t}, \Delta e_t) > 0$ for firms that choose to adjust, even if the unconditional covariance is zero. This is because in a menu cost model, firms are more likely to choose to adjust when the idiosyncratic shock and the exchange rate movement reinforce each other. Thus, $cov(\epsilon_{i,t}, \Delta e_{i,t}) > 0$, for adjusters. This implies that estimated pass-through conditional on price adjustment, $\widehat{\beta}$, is biased upward relative to true desired pass-through, β.[29]

Higher menu costs lead firms to adjust less often and by larger amounts (which increases the dispersion of price changes) as firms economize on the number of times they adjust prices. Since increases in the menu cost lead to a wider range of inaction, the importance of selection effects rises, which increases $cov(\epsilon_{i,t}, \Delta e_t)$.This increase in the selection bias causes estimated MRPT to

[26] More formally, combine the two formulas in elasticity form to get:

$$\frac{\left|\left(\frac{\partial var(\Delta p_i)}{\partial \Gamma}\frac{\Gamma}{var(\Delta p_i)}\right)\right|}{\left(\frac{\partial var(\Delta p_i)}{\partial \alpha}\frac{\alpha}{var(\Delta p_i)}\right)} = \frac{\Gamma}{1+\Gamma}\left(1+\frac{1}{\alpha^2}\frac{var(\Delta \eta_i)}{var(\Delta e_i)}\right)$$

Substituting calibrated values from the modeling section yields a ratio of approximately 200.

[27] We are not the first people to notice this bias. See the brief discussion of the bias in footnotes 7 and 26 of Gopinath, Itskhoki, and Rigobon (2010).

[28] This underlying β will be determined by α and Γ, as shown in the previous section.

[29] It's worth noting that this is only a "bias" if one is interested in measuring desired pass-through in the population. But if one is interested in measuring how much actual prices will respond to exchange rate movements, the relevant object is $\widehat{\beta}$ not β.

increase when the size of the menu costs increase.

At the same time, increasing the variance of idiosyncratic cost shocks lowers MRPT because the magnitude of the selection bias is decreasing in the size of these shocks. The intuition is simple: as the size of the idiosyncratic shocks increases, firms are more likely to adjust their prices for purely idiosyncratic reasons, which lowers $cov(\epsilon_{i,t}, \Delta e_{i,t})$, conditional on adjustment. At the same time, larger shocks mean larger price dispersion. Thus variation in the size of idiosyncratic shocks induces a counterfactual negative correlation between MRPT and dispersion.

Reviewing the conclusions from this and the previous section, it follows that heterogeneity in α, Γ or in the size of menu costs should generate a positive relationship between measured MRPT and the dispersion of price changes. However, we now show that only the "responsiveness" channel arising from variation in Γ is quantitatively successful.

5 Quantitative Model

We now formally assess the theoretical link between price change dispersion and exchange rate pass-through in an estimable quantitative model. The model allows for all the theoretical channels discussed in the previous section and also includes indirect equilibrium effects that the simple model in Section 4.1 ignored. The main model we explore builds heavily on the menu cost model of Gopinath and Itskhoki (2010). This model has been successful at matching a variety of cross-sectional and steady-state empirical facts. We build on it by formally estimating various forms of heterogeneity in the cross-section as well as by adding aggregate shocks to explain our month-level dispersion evidence. We intentionally build on this workhorse model of incomplete pass-through to show that it implies extremely tight links between our empirical facts and underlying economic primitives. The model features heterogeneity in import sensitivity, idiosyncratic volatility, and responsiveness with less than full responsiveness driven by the strategic complementarities which arise under Kimball demand.[30]

5.1 Model Description and Calibration

5.1.1 Industry Demand Aggregator

The industry is characterized by a continuum of varieties indexed by j. There is a unit measure of domestic varieties and a measure $\omega < 1$ of foreign varieties available for domestic consumption, which captures the idea that not all varieties are traded internationally.

We generate variable markups by utilizing a Kimball (1995) style aggregator:

$$\frac{1}{|\Omega|} \int_{\Omega} \Psi \left(\frac{|\Omega| \, C_j}{C} \right) dj = 1 \tag{10}$$

with $\Psi(1) = 1, \Psi'(.) > 0$ and $\Psi''(.) < 0$. C_j is the quantity demanded of variety $j \in \Omega$, where

[30] In the modeling appendix we also show that a Calvo model delivers similar conclusions about the importance of responsiveness but overall fits the data less well.

Ω is the set of all varieties available domestically. Ω has measure $1+\omega$. Individual varieties are aggregated into a final consumption good C. This intermediate aggregator contains the CES specification as a special case. The demand function for C_j implied by equation (10) is:

$$C_j = \varphi\left(D\frac{P_j}{P}\right)\frac{C}{|\Omega|}, \text{ where } \varphi(.) \equiv \Psi'^{-1}(.) \tag{11}$$

Here P_j is the price of variety j, P is the sectoral price index and $D \equiv \left[\int_\Omega \Psi'\left(\frac{|\Omega|C_j}{C}\right)\frac{C_j}{C}dj\right]$. P is defined implicitly by the following equation

$$PC = \int_\Omega P_j C_j dj$$

5.1.2 Firm's problem

Consider the problem of a firm producing variety j. Foreign and domestic firms face symmetric problems and we label foreign variables with asterisks. The firm faces a constant marginal cost:[31]

$$MC_{jt} = \frac{W_t^{1-\alpha}(W_t^*)^\alpha}{A_{jt}}$$

where W_t is the domestic wage and the parameter α is the share of foreign inputs in the firm's cost function. A_{jt} denotes idiosyncratic productivity, which follows an AR(1) in logs:

$$\log(A_{jt}) = \rho_A \log(A_{j,t-1}) + \mu_{jt} \quad \text{with} \quad \mu_{jt} \sim iid\ N(0, \sigma_A)$$

Combining yields firm profits from selling variety j in the domestic market:

$$\Pi_{jt} = \left[P_{jt} - \frac{W_t^{1-\alpha}(W_t^*)^\alpha}{A_{jt}}\right] C_{jt}$$

Firms are price-setters but face a menu cost κ when adjusting prices. Let the state vector of firm j be $S_{jt} = (P_{j,t-1}, A_{jt}; P_t, W_t, W_t^*)$ where $P_{j,t-1}$ and A_{jt} are idiosyncratic state variables and P_t, W_t, and W_t^* are aggregate state variables. The value of a firm selling variety j is characterized by the following Bellman equation:

$$\begin{aligned}
V^N(S_{jt}) &= \Pi_{jt}(S_{jt}) + E\{Q(S_{jt+1})V(S_{jt+1})\} \\
V^A(S_{jt}) &= \max_{P_{jt}} \{\Pi_{jt}(S_{jt}) + E\{Q(S_{jt+1})V(S_{jt+1})\}\} \\
V(S_{jt}) &= \max\{V^N(S_{jt}), V^A(S_{jt}) - \kappa\}
\end{aligned}$$

where $V^N(.)$ is the value function if the firm does not adjust its price, $V^A(.)$ is the value function if it adjusts, and $V(.)$ is the value of making the optimal price adjustment decision. $Q(S_{jt+1})$ is the stochastic discount factor. Each period the firm chooses whether to adjust its price by comparing

[31] This cost function can be derived from a CRS production function in domestic and foreign inputs.

the value of not adjusting to the value of adjusting net of the menu cost.

5.1.3 Sectoral equilibrium

We define $e_t \equiv \ln(W_t^*/W_t)$ as the log real exchange rate. Sectoral equilibrium is characterized by a path of the sectoral price level, $\{P_t\}$, consistent with optimal pricing policies of firms given the exogenous idiosyncratic productivity process and wage rates in the two countries. This sectoral equilibrium allows for indirect effects that we shut down in Section 4.1 but explore in our model appendix. Following Krusell and Smith (1998) and its open economy implementation in Gopinath and Itskhoki (2010), assume that $E_t \ln P_{t+1} = \gamma_0 + \gamma_1 \ln P_t + \gamma_2 e_t$. We then solve the firm's Bellman equation for a given conjecture for $\boldsymbol{\gamma}$, simulate the model and iterate to convergence. As in Gopinath and Itskhoki (2010), this forecasting rule is highly accurate in equilibrium.

We assume that all prices are set in the domestic currency, since our empirical analysis is restricted to dollar prices. Following Gopinath and Itskhoki (2010), we assume that $W_t = 1$ and that all fluctuations in the real exchange rate arise from fluctuations in W_t^*. In economic terms, these assumptions derive from assuming that the value of the domestic currency and real wage are stable relative to the exchange rate. These are good assumptions for the U.S.

5.1.4 Calibration

While there are a number of strategic complementarities that can generate variable markups (and thus incomplete pass-through), the specific form we explore in our quantitative results is the Klenow and Willis (2006) specification of the Kimball aggregator (equation 10):

$$\Psi = \left[1 - \varepsilon \ln\left(\frac{\sigma x_j}{\sigma - 1}\right)\right]^{\frac{\sigma}{\varepsilon}}, \quad \text{where} \quad x_j \equiv D\frac{P_j}{P}$$

This demand specification is governed by two parameters: $\sigma > 1$ and $\varepsilon > 0$. The elasticity and the super-elasticity of demand are given by:

$$\widetilde{\sigma}(x_j) = \frac{\sigma}{1 - \varepsilon \ln\left(\frac{\sigma x_j}{\sigma-1}\right)} \quad \text{and} \quad \widetilde{\varepsilon}(x_j) = \frac{\varepsilon}{1 - \varepsilon \ln\left(\frac{\sigma x_j}{\sigma-1}\right)}$$

Under these assumptions the markup is given by

$$\widetilde{\mu} = \frac{\sigma}{\sigma - 1 + \varepsilon \ln\left(\frac{\sigma x_j}{\sigma-1}\right)}$$

so that when $\varepsilon \longrightarrow 0$, we get a CES demand structure with an elasticity of substitution equal to σ and a markup equal to $\frac{\sigma}{\sigma-1}$. The price elasticity of desired markups is given by

$$\Gamma \equiv -\frac{\partial \ln \widetilde{\mu}}{\partial \ln P_j} = \frac{\varepsilon}{\sigma - 1 + \varepsilon \ln\left(\frac{\sigma x_j}{\sigma-1}\right)}.$$

This implies that responsiveness is decreasing in ε and increasing in σ (if $\varepsilon > 0$). Since we do not directly observe either σ or ε we cannot separately identify heterogeneity in these two parameters. For simplicity and following Gopinath and Itskhoki (2010), we assume that variation in Γ is driven solely by ε but note that variation in σ would also yield similar results. We return to this point in the discussion of our model results.

The calibrated values for all parameters are reported in Table 9. The period in our model is one month so we calibrate the discount rate to generate an annual 4% real interest rate ($\beta = 0.96^{1/12}$). We set the elasticity of demand, σ, equal to 5. This implies a steady-state markup of 25%, which is the middle of the range estimated by Broda and Weinstein (2006) using U.S. import data from 1990-2001. We assume that the log of the real exchange rate, e, follows a random walk in logs. Empirically this series is highly persistent. We set the mean increment of the innovation of the real exchange rate equal to 2.5% following Gopinath and Itskhoki (2010). To calibrate the share of imports, $\frac{\omega}{1+\omega}$, we use the share of imports as a percentage of GDP from the Bureau of Economic Analysis. The four year average (2008-2011) of this import share for the U.S. is 16.5%, which implies that $\omega = 0.2$.We set the persistence of the idiosyncratic shock process, ρ_A, to be equal to 0.85, which is in between the values used by Gopinath and Itskhoki (2010) and Nakamura and Steinsson (2008), and we set κ to target a frequency of 16%.

Finally, the parameters α, ε,and σ_A are jointly calibrated to match three moments of the data: the average level of pass-through, the R^2 from our medium run pass-through regression and the mean standard deviation of item level price changes. To get intuition for why these moments separately identify our parameters, it is useful to remember the intuition from our simple model and our baseline MRPT regression:

$$\Delta p_{i,t} = \beta \Delta e_{i,t} + \epsilon_{i,t} \tag{12}$$

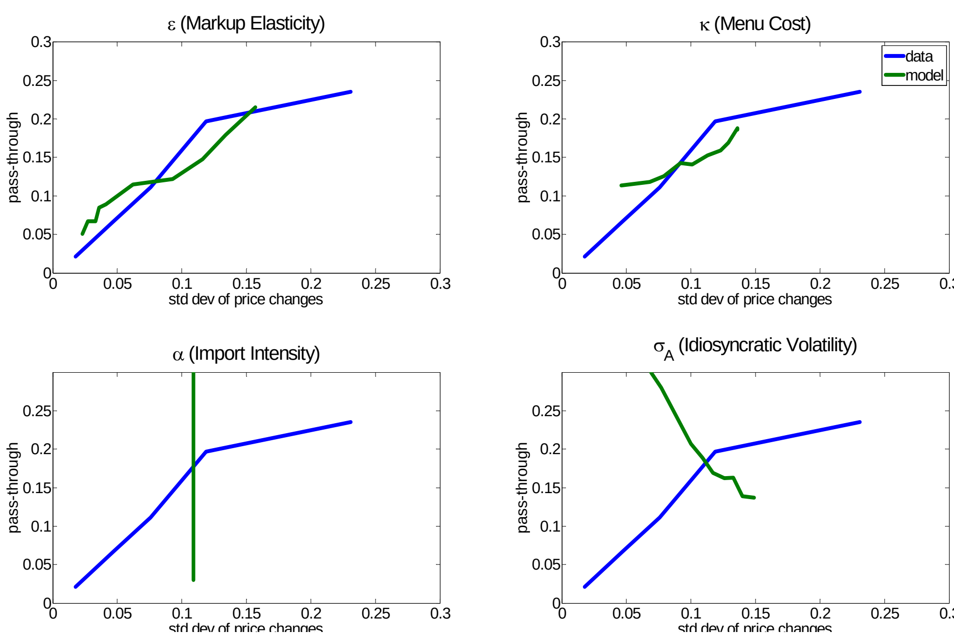
Decreasing ε means that firms respond more to both exchange rate movements and idiosyncratic shocks when adjusting prices. This increases the average level of pass-through and the standard deviation of price changes but has a negligible effect on the R^2 from estimating equation (12). This is because lowering ε increases both explained variance coming from $\Delta e_{i,t}$ and unexplained variance coming from $\epsilon_{i,t}$ by roughly equal amounts so that the ratio of the residual sum of squares to the total sum of squares remains unchanged. Increasing σ_A leads to a large increase in the variance of price change and a decrease in estimated pass-through since the selection bias conditional on price adjustment is decreasing in σ_A. Increasing σ_A also leads to a large decrease in R^2, since amplifying $\epsilon_{i,t}$ increases the residual sum of squares. Finally, increasing α leads to large increases in measured pass-through but has little effect on the variance of price changes since the variance of price changes is almost entirely driven by idiosyncratic shocks. At the same time, increasing α leads to an increase in R^2 since it increases the signal to noise ratio in the pass-through regression.

Thus, movements in these three parameters produce distinctly different effects on the average level of pass-through, the R^2 from our MRPT regression, and the mean standard deviation of item level price changes so that these three moments allow us to identify our parameters of interest. We find that the best fit parameters for α, ε,and σ_A are 0.18, 2.5 and 0.07, respectively.

5.2 Simple Comparative Statics

To understand the role of various channels in explaining the empirical relationship between MRPT and the dispersion of price changes, we begin with a simple comparative statics exercise. Each panel of Figure 6 shows results when we fix three of $\varepsilon, \kappa, \alpha$ and σ_A at their steady state values and vary the fourth parameter. For each set of parameters, we simulate a panel of firms with the same number of observations as in the BLS data and compute MRPT and the standard deviation of price changes exactly as in Section 2. For comparison, the empirical relationship between the standard deviation of price changes and MRPT that we documented in the IPP microdata is shown in blue.

Figure 6: Menu Cost Comparative Statics



The top-left panel of Figure 6 shows the results from varying ε from 0 to 100. It is apparent that variation in ε generates a strong positive correlation between price change dispersion and MRPT. Moreover, the quantitative fit is quite good: the model matches the slope, level and much of the quantitative variation of this relationship. The bottom-left panel Figure 6 shows what happens when we vary α from 0 to 1. This leads to large changes in MRPT but negligible movements in the variance of price changes. This is consistent with the discussion in Section 2.3.1.

The top-right panel shows the model-simulated results when we vary κ from 0 to 0.2. Consistent with the discussion in the previous section, variation in κ generates a modest positive relationship between MRPT and the standard deviation of price changes. This positive correlation results from the fact that increases in menu costs lead firms to tolerate wider price imbalances before adjusting and strengthen the importance of selection effects. This in turn increases the dispersion of price changes as well as measured pass-through.

Finally, the bottom-right side panel shows the results when we vary the standard deviation of the idiosyncratic shock from 0 to 0.2. Variation in σ_A generates a strong negative relationship between MRPT and the standard deviation of price changes in model-simulated data for similar reasons: larger σ_A generates more price change dispersion and implies that firms are more likely to adjust their prices for purely idiosyncratic reasons (smaller selection bias in estimated pass-through).

Thus, the comparative statics imply that variation in ε or κ can potentially replicate the observed relationship between MRPT and the standard deviation of price changes. However, explaining the positive relationship between MRPT and dispersion through variation in κ has grossly counterfactual implications for the frequency of adjustment. In the data there is a mild positive correlation between the frequency of adjustment and the dispersion of price changes. In contrast to the data, variation in κ induces an almost perfect negative correlation between dispersion and frequency: as menu costs rise, the inaction region widens and frequency falls while price change dispersion rises.

While we view this comparative statics exercise as highly informative, it has several weaknesses: 1) In the data, we are sorting firms into bins by the standard deviation of price changes. Since our comparative statics exercise instead computes results for a series of models that vary by a single parameter, we are implicitly sorting firms by this (unobserved) parameter rather than by the standard deviation of price changes. Thus, there is not a perfect match between our comparative statics simulations and our empirical exercise. 2) In the data, firms are likely to differ along many dimensions simultaneously so that heterogeneity is unlikely to be well-captured by a single parameter. 3) The comparative statics exercise is intrinsically qualitative and informal. For example, both κ and ε generate positive relationships between MRPT and dispersion and there is little formal guidance for which is a better fit even along this single moment.

We now turn to a formal estimation strategy that squarely addresses each of these weaknesses.

5.3 Indirect Inference

In this section, we allow for permanent firm heterogeneity, which we assume is unobserved by the econometrician. We then formally estimate the importance of different forms of heterogeneity in explaining our empirical results using indirect inference.[32] Motivated by tractability as well as the results from the comparative statics exercise, we allow for three dimensions of heterogeneity across firms. In particular, we allow firms to differ by κ, ε and σ_A.[33] We assume that each parameter takes on one of two values uniformly distributed around the previous mean.[34] For example, we assume that for a particular firm, κ is either equal to $\kappa_h = .043 + \kappa_\Delta$ or $\kappa_l = .043 - \kappa_\Delta$ where

[32] See Collard-Wexler (2013) and Keane and Anthony (2003) for examples of indirect inference in similar problems.

[33] While it would also be possible to allow for heterogeneity in α, our comparative statics exercise suggests that this parameter plays no role in explaining the relationship between MRPT and dispersion or in the relationship between dispersion and frequency. However, shutting down heterogeneity along this dimension substantially reduces the computational burden involved in estimation.

[34] When relevant, we bound the value of $\kappa_l, \varepsilon_l, \sigma_l$ at 0.

κ_Δ is a parameter to be estimated which governs the degree of menu cost differences across firms. We allow for a similar two point symmetric distribution for each source of heterogeneity so that we have three parameters which must be estimated: $\theta = (\kappa_\Delta, \sigma_\Delta, \varepsilon_\Delta)$.

Fixing $\kappa_\Delta, \sigma_\Delta, \varepsilon_\Delta$ there are then eight different types of firms in our model (taking on high or low values for each parameter).[35] After solving for the sectoral equilibrium with these eight firm types we simulate a firm panel, which we sample exactly as in the BLS microdata to account for any small sample issues which might arise in our empirical specification. From this firm panel we calculate an auxiliary model that consists of fifteen reduced form moments $g(\theta)$ which capture essential features of the data. We then try to match these simulated moments to their empirical counterparts.

This indirect inference estimation procedure explicitly addresses the concerns identified with the comparative statics exercise: Simulated and actual data are treated identically and we use no information from simulated data that is not available in actual data. In addition, we explicitly allow for the presence of multiple sources of heterogeneity and formally assess their relative importance.

To construct our empirical moments, we begin by sorting firms into five bins by their standard deviation. We then calculate the relative standard deviation of price changes across each standard deviation bin, the relative MRPT across each bin, and the relative frequency across each bin.[36] The first five moments reflect the model's ability to capture the heterogeneity in price change dispersion observed in the data. The second five moments capture the relationship between this dispersion and pass-through. The final five moments capture the relationship between dispersion and frequency, which we previously argued are useful for separately identifying heterogeneity in menu costs from heterogeneity in responsiveness.[37]

Given these 15 moments, we then pick our 3 parameters to solve $\widehat{\theta} = \arg\min_\theta g(\theta)' W(\theta) g(\theta)$ where $W(\theta)$ is a positive definite weight-matrix.[38] Table 10 displays resulting parameter estimates as well as several measures of model fit. The first take-away from Table 10 is that the estimated level of ε heterogeneity is large and significant. In contrast, heterogeneity in σ_A is significant but not strongly so, and a null hypothesis that there is no heterogeneity in κ cannot be rejected. We can also assess the overall model fit. Using standard over-identification tests our full model cannot be rejected at 99% confidence levels. We can also investigate restricted models that turn off various sources of heterogeneity. The results for the restricted models show that the model

[35] While it would be desirable to allow for more than a 2-point distribution of heterogeneity for each parameter, allowing for a 3-point distribution would require solving the model for 27 different types of firms while allowing for a 4-point distribution would require 64 firm types, so it is clear that the problem rapidly rises in difficulty. Since we want to estimate the model, we must resolve it for a large number of $\kappa_\Delta, \sigma_\Delta, \varepsilon_\Delta$ which rapidly becomes infeasible.

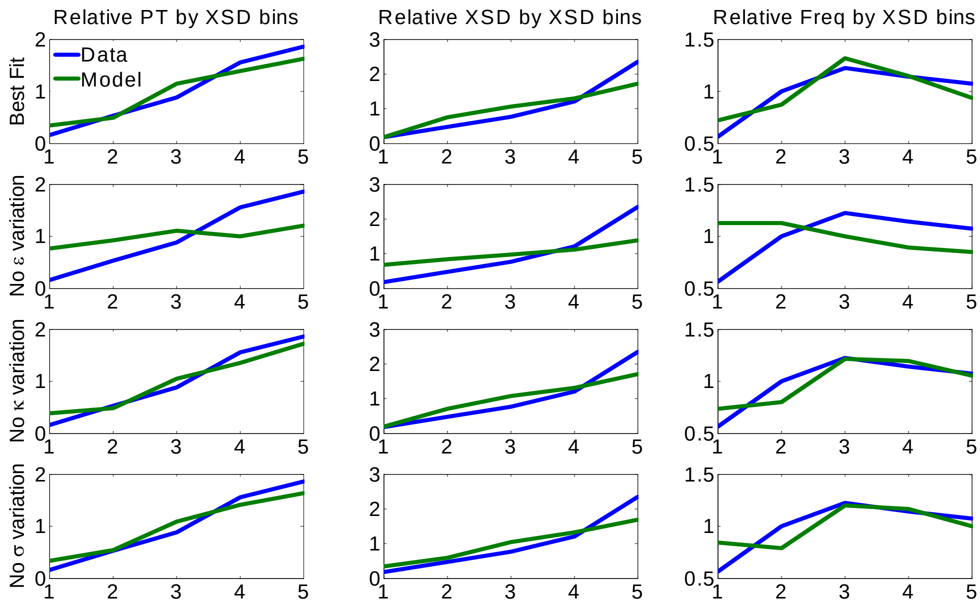
[36] We concentrate on the relative values rather than the absolute values because our benchmark calibration is not perfectly able to match the level of XSD, MRPT and freq. We think of both our empirical exercise and our exercise with heterogeneity largely as being about matching the relative differences across firms. Nevertheless, redoing the results using absolute rather than relative moments did not qualitatively change the conclusions.

[37] As is standard in indirect inference, our auxiliary model need not have any structural interpretation. For example, we have already noted that our OLS MRPT regression will pick up both direct effects of parameters on β as well as indirect effects on covariance terms.

[38] We pick $W(\theta)$ to be the standard efficient weight matrix so that we can apply asymptotic formulas for standard errors but using an identity weight matrix did not change our qualitative conclusions.

with no heterogeneity in ε can easily be rejected while models with no heterogeneity in κ or no heterogeneity in σ_A cannot be rejected in favor of the full model.

Figure 7: Full and Restricted Model Fits



The numerical results can be seen more easily in Figure 7, which shows the model fit to all fifteen moments as well as the fit of restricted models which shut down various sources of heterogeneity.

The main take-away from this visual inspection is that the fit in the second row is dramatically worse than the fit in the first row. Turning off heterogeneity in ε means the next-best model fit does not generate enough heterogeneity in price change dispersion, fails to generate enough of a positive relationship between price change dispersion and pass-through, and it implies a negative rather than positive correlation between dispersion and pass-through. In contrast, turning off heterogeneity in menu costs or in volatility has only negligible effects on the model fit.

5.4 Aggregate Shocks

5.4.1 Baseline Results

Now that we have shown that variation in responsiveness is crucial for explaining the empirical relationship between item-level price change dispersion and pass-through, we turn to understanding our month-level dispersion results. In general, we find that our conclusions about the role of various economic primitives in the cross-section transfer almost directly to the time-series. For this reason, our discussion of these simulations is intentionally brief.

In the previous section, we assumed that there was heterogeneity across firms that was constant across time. Instead, we now assume that firms are identical but are subject to various aggregate shocks.[39] We consider aggregate shocks to each of our parameters in turn. For expositional purposes we describe only ε shocks, but we treat our other shocks analogously.

For simplicity, we assume that ε_t follows a two-state Markov process with transition probabilities $\begin{bmatrix} \Pi_{11} & \Pi_{12} \\ \Pi_{21} & \Pi_{22} \end{bmatrix}$. We also allow the Krusell-Smith forecast for the sectoral price level to depend on ε_t.[40] We have little guidance on either the size or the persistence of our aggregate shocks, so rather than taking a strong stand on this process, we simply report results for a range of aggregate shocks. In particular, we report results for two different shock sizes. Under the "small" shock, ε_t moves between $(1+.6)\,\overline{\varepsilon}$ and $\frac{1}{1+.6}\,\overline{\varepsilon}$ where $\overline{\varepsilon}$ is the previous baseline calibration. This shock produces time-series variation that is one-fifth of the cross-sectional variation explored in the previous section. In addition, we consider a "large" shock calibration that moves ε_t between $4\overline{\varepsilon}$ and $\frac{1}{4}\,\overline{\varepsilon}$. This large shock produces time-series variation that is comparable to the cross-sectional variation in the previous section. We have computed results for both a low monthly shock persistence of $\Pi_{11} = \Pi_{22} = 0.90$ and a high monthly persistence of $\Pi_{11} = \Pi_{22} = 0.975$. Changing the persistence barely affected our results, so for brevity we report only the high persistence case.

Table 11 shows results for the model with different aggregate shocks. In all cases, we divide months in thirds by their month-level dispersion and calculate pass-through in high and low dispersion months. Overall aggregate shocks to ε_t are most consistent with our empirical time-series results. Increases in ε reduce pass-through and the standard deviation of price changes. Under the "large shock" calibration, variation in ε can explain roughly the same amount of variation in MRPT observed in the data. However, the movements in price change dispersion across time are somewhat too large: the model produces a cross-sectional standard deviation of price changes that ranges from 0.05 to 0.13. In the data, the comparable range is 0.12 to 0.15. (As previously mentioned, our baseline calibration mildly underpredicts the average standard deviation of price changes in the data). In ongoing work we plan to explore whether asymmetric or continuous shocks that relax the binary assumption can provide a better fit to the data. Nevertheless, shocks to ε produce variation in pass-through and dispersion that is relatively consistent with the data.

In contrast, shocks to σ_A induce the wrong correlation between the standard deviation of price changes and pass-through. In addition, they produce time-series variation in both the dispersion of price changes and in frequency that are substantially too large relative to the data.

Shocks to κ do a reasonable job of matching the time-series relationship between dispersion and pass-through, but they do a terrible job of matching the time-series relationship between frequency and pass-through. In the data, price change dispersion, pass-through and frequency all comove while with shocks to κ there is a strong negative relationship between frequency and price change dispersion. Furthermore, the time-series variation in frequency is much too large.

Shocks to α induce lots of movement in pass-through but almost no movement in the standard deviation of price changes. In addition, the small movement in price change dispersion induced

[39] Including item-level heterogeneity together with aggregate shocks did not alter our conclusions but makes the model somewhat more complicated.

[40] That is, we assume that $E_t \ln P_{t+1} = \gamma_0 + \gamma_1 \ln P_t + \gamma_2 e_t + \varepsilon_t \times [\gamma_4 + \gamma_5 \ln P_t + \gamma_6 e_t]$. Again we find that the Krusell-Smith forecasting rule is highly accurate.

by shocks to α goes in the wrong direction. As α rises, pass-through rises but the cross-sectional standard deviation of price changes falls. That is because a large α essentially increases the size of the exchange rate shock relative to the size of idiosyncratic shocks. Since the exchange rate shock is common to all firms, this reduces the cross-sectional dispersion of price changes. Thus, as in the cross-sectional results, only shocks to ε do a reasonable job of reproducing the empirical evidence.

5.4.2 Additional Shocks

In addition to the above aggregate shocks which were also explored in the cross-section, we have also modeled two additional aggregate shocks which are more applicable to the time-series. First, we allowed the volatility of exchange rates to change across time since the 2008 recession was also associated with greater exchange rate volatility. However, we found that even large increases in exchange rate volatility only have mild quantitative effects, and qualitatively have the wrong sign relative to the empirical evidence. That is, increases in the volatility of exchange rates mildly increase pass-through, but they (very mildly) decrease the degree of month-level dispersion. This is for the same reason that increases in α decrease the dispersion of price changes.

In addition to greater exchange rate volatility, it is also possible that the large degree of pass-through observed during the Great Recession was in part driven by the fact that the recession was a large shock which affected many firms. If a shock is common to more firms, then it might have greater general equilibrium effects and thus lead to greater pass-through. To assess the role of the "commonness" of shocks, we introduced time-variation in the fraction of firms that are sensitive to the exchange rate, ω. As ω rises, exchange rate shocks affect more firms and general equilibrium effects should increase in importance. However, we find that the quantitative effect of changes in ω on pass-through is relatively small and that there are no effects of ω on the dispersion of price changes. Increasing ω from 0.2 to 0.9 only increases pass-through from roughly 16% to 23% and has no effect on dispersion. Thus, general equilibrium effects in our model cannot account for the empirical relationship between month-level dispersion and exchange rate pass-through.

5.4.3 Interpreting ε shocks

Does time-variation in ε have a natural economic interpretation? While it is difficult to provide direct evidence of a time-varying super elasticity of demand, Gopinath and Itskhoki (2010) argued for the importance of such variation in the cross-section. It is easy to imagine that the forces which drive differences in super elasticity across firms might also vary across time. Furthermore, it is important to note that time-variation in the elasticity of substitution rather than the super elasticity also delivers similar quantitative results. The reason is that on average, $\Gamma = \frac{\varepsilon}{\sigma - 1}$, so that shocks to σ also generate time-variation in responsiveness as long as firms exhibit some incomplete markup adjustment ($\varepsilon > 0$).

What is crucial for explaining the time-series relationship between pass-through and dispersion is reduced-form time-variation in Γ; the structural source of the time-variation is less important. For example, business cycle shocks to "market competitiveness" would equally well explain our time-series patterns through the same Γ channel. If certain periods of time such as recessions are characterized by increased competition, with larger σ and lower markups, they will also be times of greater responsiveness and price change dispersion.[41] This story is intuitively appealing and is consistent with the fact that both pass-through and dispersion are countercyclical.

While we model time-varying responsiveness using a Kimball demand framework, this mechanism is much more general. Any shock to strategic complementarities across time that affects markup adjustment (Γ) will deliver similar predictions. As shown in Section 4 of Burstein and Gopinath (2013), a number of other mechanisms can also generate less than perfect responsiveness including desire to maintain market share (Atkeson and Burstein (2008)), customer shopping concerns (Paciello, Pozzi, and Trachter (2013)) or local distribution costs. We believe that better understanding the source of "responsiveness" shocks and time-varying strategic complementarities is an interesting avenue for future research. While these shocks help fit the data, we think there is much work to be done exploring their plausibility, size and implications for business cycles more generally.

5.5 What Drives Dispersion?

Our paper joins a long literature documenting countercyclical dispersion of various economic variables. At the same time, theory has emerged trying to match this empirical evidence and explore its macroeconomic implications. This theoretical work has focused almost exclusively on "uncertainty" or "volatility" shocks that raise the variance of shocks hitting agents in the economy.

While the theoretical literature has focused on volatility shocks, our models show that this is not the only possibility consistent with existing empirical evidence. Looking at equation 9 shows that increasing $var(\epsilon_i)$ or lowering Γ both increase the cross-sectional dispersion of price changes. That is, greater dispersion of outcomes could be explained by greater volatility of shocks and constant responsiveness, or it could be explained by greater responsiveness and constant volatility. Bachmann and Moscarini (2011) demonstrate this point in a model of firm learning where firms endogenously change prices more aggressively during recessions. By looking only at data on the variance of outcomes, it is impossible to differentiate greater volatility from greater responsiveness.[42]

In contrast to the existing empirical literature, in our open economy environment we can separately identify changes in volatility from changes in responsiveness. Identifying the source of our empirical pass-through-volatility relationship was precisely the point of the previous sections. Those results showed that our import price data strongly supports time-variation in responsiveness rather than volatility shocks to explain countercyclical dispersion. Increases in volatility are unable to explain an increase in pass-through. In contrast, greater responsiveness increases both price change dispersion and exchange rate pass-through in a manner consistent with the data. This

[41] If σ is bigger in recessions this implies that the average markup is strongly procyclical. While there is a lot of debate about the cyclicality of markups, recent work by Nekarda and Ramey (2013) shows that average markups are procyclical.

[42] Note that this same argument applies to dispersion of any outcome that includes an endogenous component, for example revenue TFP dispersion depends on firms' price decisions.

result holds across a variety of price-setting environments whether price adjustment is frictionless, time-dependent or state-dependent.[43] Thus, we believe it is quite general.

Together, our results suggest that the literature studying countercyclical dispersion may have embraced time-varying volatility too quickly. Time-variation in the strength of strategic complementarities (and thus responsiveness) appears to be more relevant, at least for import price-setting. Understanding the generalizability and empirical relevance of our results for other sectors of the economy and other economic outcomes is an important avenue for future research.

6 Conclusions

In this paper, we exploit the open economy environment to provide evidence on the response of inflation to cost shocks at a moment in time. We start by documenting a robust positive relationship between exchange rate pass-through and the dispersion of item-level price changes. Through a battery of robustness checks, we argued that this relationship was not driven by other observables or confounding variables. Furthermore, price change dispersion varies dramatically across time. Ignoring this variation induces a huge, time-varying bias when estimating pass-through. In other words, pass-through is not a single number and ignoring time-variation produces a misleading picture of how prices will respond to shocks at a particular point in time.

This result has important implications for the nominal transmission mechanism more generally. One reason for studying exchange rate pass-through in the first place is because it can shed light on the nominal transmission mechanism using a large, observable shock to nominal cost. Our results provide "model-free" evidence that the transmission mechanism varies systematically across time with the distribution of price changes in the economy. We believe this result is extremely important for policy making, because monetary policy is not conducted at random times. Monetary easing is particularly likely to occur during times when we have shown that price flexibility is systematically higher than average.

After documenting the empirical relationship between dispersion and pass-through, we estimated a quantitative price-setting model to understand it. Our estimated model strongly prefers variation in strategic complementarities to variation in volatility, menu costs, or import intensity. Thus, our empirical evidence does not support volatility shocks as a source of countercyclical dispersion, but it does suggest a promising alternative. In particular, time-variation in responsiveness appears to better fit the data. A large existing literature has argued that imperfect responsiveness is important for explaining the average behavior of pass-through, and we find the idea that this responsiveness might also vary across time to be quite plausible.

While we showed that "responsiveness" matters, there are many different mechanisms that map into responsiveness as a reduced form. We believe that trying to disentangle these mechanisms is an interesting avenue for future research. Our BLS data has quite limited firm-level covariates which might be used to shed light on underlying mechanisms, but alternative data sets with more

[43] See the appendix for results for the Calvo model.

firm-level characteristics exist. Together our empirical and modeling results suggest that exploring time-variation in the competitive structure of markets or in strategic-complementarities and trying to test these ideas in alternative data is a promising research topic.

Our results have both obvious and more subtle implications for policy. Most obviously, if policy makers want to understand how prices are likely to respond to exchange rate changes or predict the real responses to nominal exchange rate changes, they cannot ignore individual price-setting behavior.[44] More subtly, if policy makers want to mitigate the adverse effects of price change volatility then it is important to understand what leads to this volatility. At least in the context of import prices, our results suggest that policy makers should focus on policies that affect market structure and firms' responsiveness rather than on policies that reduce uncertainty and volatility.

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[44] One might wonder if our empirical observations are implementable for actual predictions. We note that in our BLS data, dispersion is much more precisely estimated at high frequencies than is pass-through. In addition, new online data sets such as those in the Billion Prices Project introduced by Cavallo (2012) can potentially calculate daily measures of price change volatility.

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Table 1: Average medium-run pass-through

β	$se(\beta)$	t-stat	N_{obs}	R^2
0.144	0.014	10.17	95284	0.067

Table 2: Interaction Specification: Item-Level Volatility

	Average pass-through		Volatility (Item-Level)		Frequency		N_{obs}	R^2
	β^{avg}	$se(\beta^{avg})$	β^{Vol}	$se(\beta^{Vol})$	β^{freq}	$se(\beta^{freq})$		
All countries, all items ex petroleum								
	0.14	0.01	0.05	0.02			95284	0.07
	0.14	0.01	0.05	0.02	0.02	0.01	95284	0.07
OECD countries, all items ex petroleum								
	0.18	0.02	0.09	0.03			53469	0.08
	0.19	0.02	0.08	0.03	0.07	0.02	53469	0.08
All countries, all manufacturing items								
	0.14	0.01	0.06	0.02			78439	0.09
	0.13	0.01	0.06	0.02	0.03	0.01	78439	0.09

Note: Robust standard errors clustered by country*PSL pair. Primary strata lower (PSL), defined by the BLS, represents 2 to 4-digit sectoral harmonized codes. In all specifications, volatility is measured using the item-level standard deviation of price changes.

Table 3: Interaction Specification: Month-Level Volatility

	Average pass-through		Volatility (Month-Level)		Frequency		N_{obs}	R^2
	β^{avg}	$se(\beta^{avg})$	β^{Vol}	$se(\beta^{Vol})$	β^{freq}	$se(\beta^{freq})$		
All countries, all items ex petroleum								
- Vol=Month-level IQR	0.14	0.01	0.06	0.01			95284	0.07
	0.14	0.01	0.06	0.01	0.01	0.01	95284	0.07
- Vol=Month-level XSD	0.13	0.01	0.05	0.01			95284	0.07
	0.13	0.01	0.05	0.01	0.03	0.01	95284	0.07
OECD countries, all items ex petroleum								
- Vol=Month-level IQR	0.17	0.01	0.05	0.01			53469	0.08
	0.17	0.01	0.05	0.01	-0.01	0.01	53469	0.08
- Vol=Month-level XSD	0.17	0.01	0.06	0.01			53469	0.08
	0.18	0.01	0.07	0.01	-0.03	0.02	53469	0.08
All countries, all manufacturing items								
- Vol=Month-level IQR	0.13	0.01	0.05	0.01			78437	0.09
	0.13	0.01	0.05	0.01	0.00	0.01	78437	0.09
- Vol=Month-level XSD	0.13	0.01	0.05	0.01			78437	0.09
	0.13	0.01	0.05	0.01	0.00	0.01	78437	0.09

Note: Robust standard errors clustered by country*PSL pair. Primary strata lower (PSL), defined by the BLS, represents 2 to 4-digit sectoral harmonized codes.

Table 4: Interaction Specification: Month-Level Dispersion Robustness

	Average pass-through		Volatility (Month-Level)		Frequency/Subs		N_{obs}	R^2
	β^{avg}	$se(\beta^{avg})$	β^{Vol}	$se(\beta^{Vol})$	β^{freq}	$se(\beta^{freq})$		
Controls								
- Time trend + Month	.135	.025	.058	.012			95284	.075
- GDP growth	.146	.013	.054	.009			95284	.071
- Erate SD	.156	.015	.056	0.010			95284	.072
- Frequency	.140	.013	.063	.010	.011	.012	95284	.072
- Product subs	.143	.013	.062	.010	.0004	.011	95284	.071
- Time trend + Month + Frequency	.128	.025	.054	.012	.011	.014	95284	.076
+ GDP growth + Erate SD								

Note: Robust standard errors clustered by country*PSL pair. Primary strata lower (PSL), defined by the BLS, represents 2 to 4-digit sectoral harmonized codes. In all specifications month-level price change volatility is measured using the interquartile range. The sample is all countries, all items ex petroleum. Erate SD is the std dev of the exchange rate in a 12-month rolling window around current date

Table 5: Interaction Specification: Month-Level Dispersion Robustness: Pre-2008

	Average pass-through		Volatility (Month-Level)		Frequency/Subs		N_{obs}	R^2
	β^{avg}	$se(\beta^{avg})$	β^{Vol}	$se(\beta^{Vol})$	β^{freq}	$se(\beta^{freq})$		
Controls								
- Time trend + Month	.115	.024	.028	.008			73736	.073
- Frequency	.106	.014	.030	.008	.014	.012	73736	.070
- Product subs	.109	.014	.027	.008	.012	.013	73736	.070
- Time trend + Month + Frequency	.098	.024	.029	.011	.021	.014	73736	.074
- Time trend + Month + Product subs	.115	.024	.028	.011	.005	.012	73736	.073

Note: Robust standard errors clustered by country*PSL pair. Primary strata lower (PSL), defined by the BLS, represents 2 to 4-digit sectoral harmonized codes. In all specifications month-level price change volatility is measured using the interquartile range.

Table 6: Interaction Specifications with both Cross-Item and Cross-Month Dispersion

	Average pass-through		Item-Level Volatility		Month-Level Volatility		N_{obs}	R^2
	β^{avg}	$se(\beta^{avg})$	β^{XSD}	$se(\beta^{XSD})$	β^{IQR}	$se(\beta^{IQR})$		
Controls								
- No additional controls	.141	.013	.043	.017	.060	.009	95284	.072
- Item level frequency	.139	.013	.041	.017	.060	.009	95284	.072
- Aggregate frequency	.137	.013	.041	.017	.060	.009	95284	.073
- Time trend + Month	.137	.024	.042	.017	.055	.012	95284	.076
- All above controls	.125	.024	.042	.017	.055	.012	95284	.077

Note: Robust standard errors clustered by country*PSL pair. Primary strata lower (PSL), defined by the BLS, represents 2 to 4-digit sectoral harmonized codes. The sample is all countries, all items ex petroleum.

Table 7: Within and Between

Sector Definition	β^{avg}	β^{VAR_W}	t-stat W	β^{VAR_B}	t-stat B
2-digit	.141	.056	5.95	.010	0.82
4-digit	.141	.036	3.29	.034	2.59

Within is month-level variance of price changes within sectors. Between gives variance of inflation rates across sectors

Table 8: Alternative Pass-through Specifications

	Average pass-through		Item-Level Volatility		Month-Level Volatility			
Fixed Horizon:	β^{avg}	$se(\beta^{avg})$	β^{XSD}	$se(\beta^{XSD})$	β^{IQR}	$se(\beta^{IQR})$	N_{obs}	R^2
1 Month	.027	.007	.034	.013	.023	.006	496060	.018
3 Month	.054	.011	.048	.023	.026	.007	448400	.049
6 Month	.085	.016	.069	.033	.026	.009	384827	.098
12 Month	.113	.018	.093	.022	.023	.009	282572	.169
Lagged Specification:								
Current Ex. Rate (β_1)	.146	.015	.040	.020	.063	.010		
Previous Ex Rate (β_2)	.082	.010	.040	.017	.054	.010	83043	.082

Note: Robust standard errors clustered by country*PSL pair.

Table 9: Parameter Values

Parameter	Symbol	Menu Cost Model	Source
Discount Factor	β	$0.96^{1/12}$	Annualized interest rate of 4%
Fraction of imports	$\omega/(1+\omega)$	16.5%	BEA input-output table
Cost sensitivity to ER shock			
Foreign firms	α^*	0.18	Estimation (see text)
U.S. firms	α	0	
Menu cost	κ	4.3%	Estimation (see text)
markup elasticity	ε	2.5	Estimation (see text)
Demand elasticity	σ	5	Broda and Weinstein (2006)
Std. dev. Exchange rate shock, e_t	σ_e	2.5%	Match bilateral RER
Idiosyncratic productivity process, a_t			
Std. dev. of shock	σ_A	7.0%	Estimation (see text)
Persistence of shock	ρ_A	0.85	Gopinath and Itshkoki (2010)

Table 10: Estimated Parameters and Fit

Parameter	Estimate	95% Confidence Interval
ε_Δ	10	(8.14,11.86)
σ_Δ	.03	(.0035,.0565)
κ_Δ	.014	(-.0125,.0405)

Models	Wald-Statistic/Likelihood Ratio	95% Critical Value	99% Critical Value
Unrestricted Model	25.76	21.03	26.22
$\varepsilon_\Delta = 0$	46.9	3.84	5.63
$\sigma_\Delta = 0$	3.23	3.84	5.63
$\kappa_\Delta = 0$	3.57	3.84	5.63

Asymptotic s.e.'s for parameters in parantheses. Unrestricted model Wald-Statistic: $g\left(\widehat{\theta}\right)' W\left(\widehat{\theta}\right)' g\left(\widehat{\theta}\right) \sim \chi^2(12)$

Restricted models: $2\left[g\left(\widehat{\theta}_r\right)' W\left(\widehat{\theta}_u\right)' g\left(\widehat{\theta}_r\right) - g\left(\widehat{\theta}_r\right)' W\left(\widehat{\theta}_u\right)' g\left(\widehat{\theta}_r\right)\right] \sim \chi^2(1)$

Table 11: Aggregate Shocks

	Data (Low XSD)			Data (High XSD)		
	XSD	MRPT	FREQ	XSD	MRPT	FREQ
	0.12	0.08	0.14	0.15	0.17	0.18
	High ε			Low ε		
	XSD	MRPT	FREQ	XSD	MRPT	FREQ
Small	0.08	0.15	0.11	0.11	0.20	0.13
Large	0.05	0.12	0.08	0.13	0.22	0.15
	Low σ			High σ		
	XSD	MRPT	Freq	XSD	MRPT	FREQ
Small	0.07	0.19	0.08	0.13	0.14	0.21
Large	0.07	0.20	0.08	0.23	0.11	0.45
	High κ			Low κ		
	XSD	MRPT	FREQ	XSD	MRPT	FREQ
Small	0.10	0.21	0.08	0.09	0.15	0.16
Large	0.10	0.30	0.04	0.07	0.13	0.29
	High α			Low α		
	XSD	MRPT	FREQ	XSD	MRPT	FREQ
Small	0.09	0.29	0.12	0.09	0.13	0.12
Large	0.08	0.66	0.14	0.09	0.06	0.13

Small = ± 0.6 factor, Large = ± 3.0 factor. Binary agg shock has persistence .975

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Appendix Materials - Not For Publication

7 Empirical Appendix - Not for Publication

In this empirical appendix, we provide a number of additional robustness checks that extend the baseline results in the body of the text.

7.1 Additional Item-Level Results

In this section we perform a variety of robustness checks for our item-level results. We begin by further discussing the role of adjustment frequency, then return to compositional issues. Finally, we discuss a battery of additional robustness checks, alternative samples, and placebo regressions for our baseline results.

In the main text we estimate interaction specifications to argue that differences in frequency across items do not explain our relationship between cross-item dispersion and MRPT. However, that specification assumes that the effects of dispersion are linear and does not allow for the effects of other controls to vary with item-level characteristics. In this robustness check we provide further evidence that the relationship between cross-item dispersion and MRPT is not driven by frequency. This is an important concern to address because Gopinath and Itskhoki (2010) showed that there is a robust relationship between LRPT and the frequency of adjustment. In order to further address this concern, we split items first into equal weighted frequency quintiles then examine the relationship between MRPT and item-level dispersion within each frequency quintile. In other words, we examine the relationship between pass-through and dispersion holding the frequency of adjustment (roughly) constant. The results are shown in Figure 8.

The relationship between pass-through and dispersion is increasing within each frequency quintile, and the magnitude of the increase is substantial. Average pass-through increases from 3% to 20% as we move from the lowest to highest XSD quintile. This complements the evidence in the text that the relationship between MRPT and price dispersion does not seem to be driven by differences in frequency across items.[45]

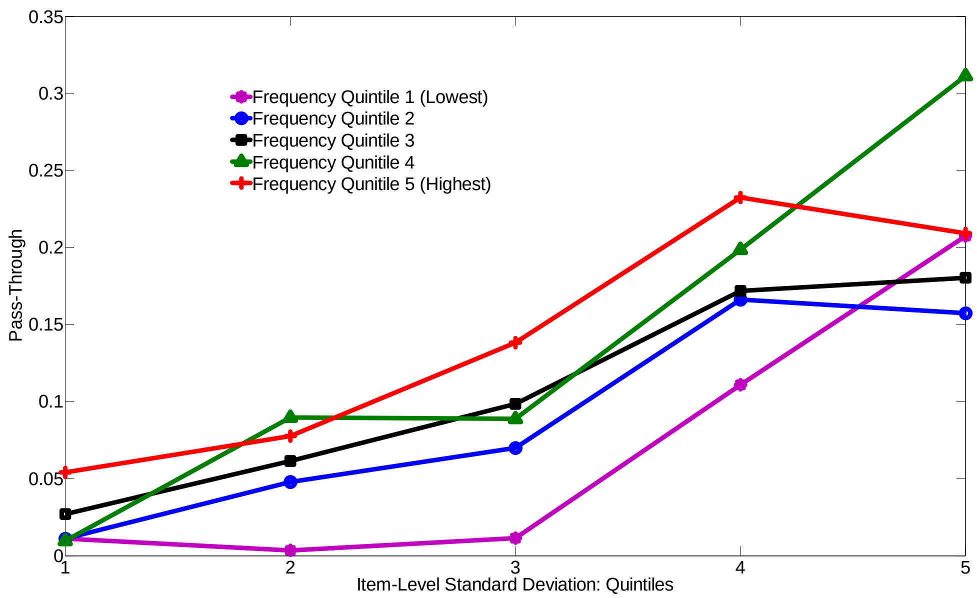
We next address whether our results are driven by choice of which items we sampled. Our baseline results utilize all of the items in the IPP micro data excluding petroleum. Is the strong relationship between pass-through and dispersion affected if we split by other observable product characteristics? To address this question we first examine the sub-sample of goods that can be classified as differentiated, following Rauch's classification, as well as the sample of goods that are manufactured.[46] For differentiated goods, Figure 9 shows that moving from the lowest to highest-dispersion quintile raises MRPT from 2% to 27%. Similar results obtain when using all manufactured goods. In all cases, the difference in pass-through across XSD bins is strongly statistically significant.

In addition to splitting by product type, we can also split our sample by country of origin. Perhaps our results are driven by compositional differences in the behavior of items coming from

[45] This is not surprising since Gopinath and Itskhoki (2010) document a significant relationship between LRPT and the frequency of price adjustment but find no relationship between MRPT and the frequency of price adjustment.

[46] Items are classified as manufacturing items if their 1-digit SIC 1987 codes begin with a 2 or a 3.

Figure 8: Medium-run passthrough and XSD controlling for the frequency of price adjustment



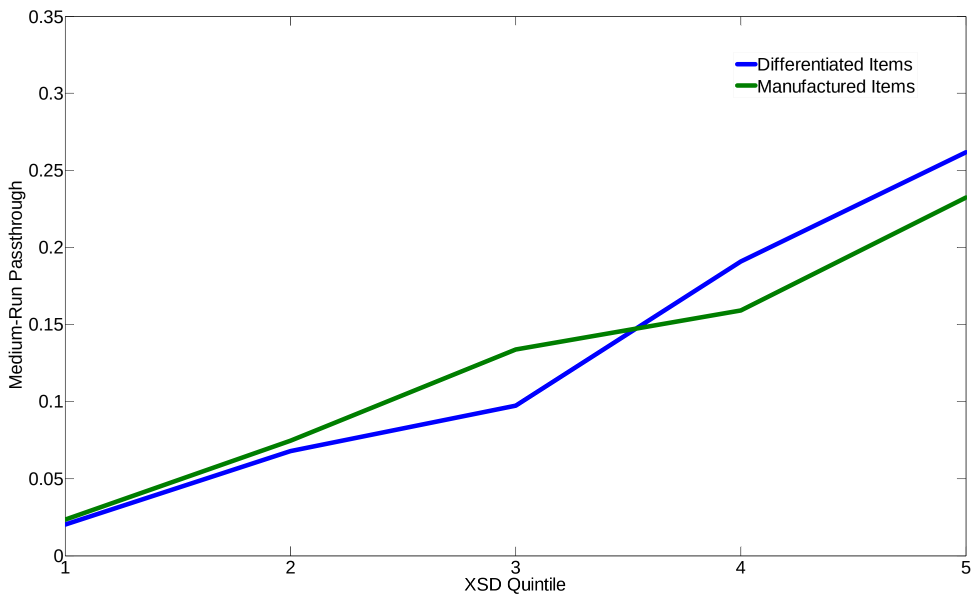
different countries. Figure 10 shows this is not the case. For all countries and country groups with greater than 5000 price observations there is a strong upward sloping relationship between item-level dispersion and MRPT. While the relationship is insignificant for Mexico, the sample size is small relative to more aggregated country groups and Canada. Among countries with at least 5000 observations, only Japan has fewer observations (and Japan exhibits a significant upward sloping relationship).[47]

In addition to these alternative binned regressions, Table A1 shows the results from estimating equation 3 for a variety of alternative sub-samples and alternative specifications. The first robustness check only uses items which have at least 3 changes. It is difficult to precisely measure dispersion for items with few price changes, so there is some concern that our baseline specifications might be polluted by outliers and small sample issues. However, the first two rows of Table A1 show that our results are essentially unchanged when restricted to items with at least 3 price changes. We have also restricted the analysis to items with at least 5 price changes and arrived at similar results.

In our second and third robustness checks, we use trade-weighted exchange rates (the broad and major currency one respectively) instead of the relevant bilateral exchange rate. As rows 3-6 show, the price dispersion is again both economically and statistically significant. A one standard deviation increase in price dispersion causes MRPT to increase relative to average pass-through by

[47] While it would be desirable to run this specification for each country rather than aggregating to country groups, our empirical specification requires splitting the data into fifths and then estimating second moments of price-setting on these bins, so using smaller countries becomes infeasible.

Figure 9: Medium-run passthrough by XSD



over 50%.[48]

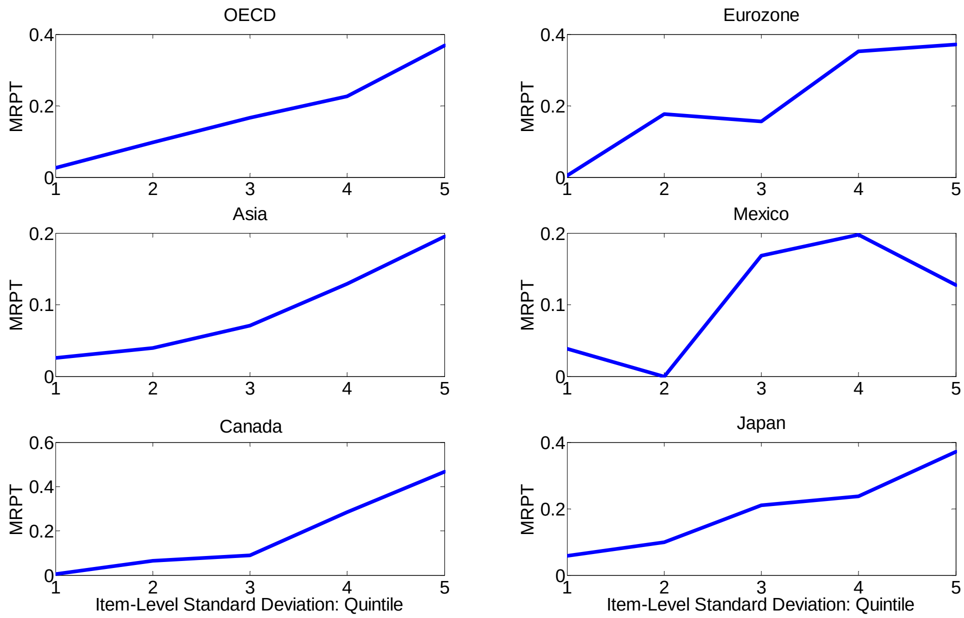
Rows 7-10 show the results from our fourth and fifth robustness checks. In these robustness checks, we run placebo regressions to see whether our results are spuriously driven by small sample issues. In these placebo regressions, when estimating equation 3, we substitute the total number of price changes observed for an item or the number price observations respectively for XSD. These placebo regressions test whether our results are driven by a correlation between measured dispersion and item sample sizes. Table 5 shows that, as desired, the coefficient on β^{XSD} is not significant when we replace XSD with placebos. This suggests that the relationship between MRPT and price dispersion is not being driven by sampling error. Finally, rows 11 and 12 show the results from estimating equation 3 using a median regression rather than OLS. Median regressions are more robust to the presence of outliers. Once again, the price dispersion effect is strongly significant.

7.2 Additional Month-Level Results

In this section, we perform a variety of robustness checks for our baseline month-level results. First, in the main text, we showed that MRPT was increasing in IQR quintile. However, the standard errors for this bin approach were large due to limited sample sizes. Thus we want to more formally test for the presence of a time-series relationship between price change dispersion and MRPT. We

[48] Consistent with what was found in Nakamura and Steinsson (2012), average passthrough is significantly higher when we use broader exchange rates measures. The much larger response of prices to the trade-weighted exchange rate suggests that items respond to exchange rates beyond the bilateral one, presumably due to the role of intermediate inputs and strategic complementarities in pricing.

Figure 10: Medium-run Passthrough by XSD (Country Level Results)



begin by calculating the cross-sectional interquartile range of price changes for each month in our sample. We then split our sample in thirds by the interquartile range. Let I_t^{high} be an indicator for the one-third of months with the highest interquartile range in our sample. Similarly, let I_t^{low} be an indicator for the one-third of months with the lowest interquartile range in our sample. Our baseline time-series specification is then:

$$\Delta p_{i,t} = \left[\beta^{high} \Delta_c e_{i,t} + Z'_{i,t} \gamma^{high}\right] I_t^{high} + \left[\beta^{low} \Delta_c e_{i,t} + Z'_{i,t} \gamma^{low}\right] I_t^{low} + \epsilon_{i,t}.$$

Table A2 shows that during high dispersion months, MRPT is 21% while in low dispersion months MRPT is only 8%. This difference is both economically and statistically significant, with pass-through more than doubling between low and high dispersion months. Table A2 also shows that these differences remain significant for alternative sample selections as well as alternative measures of cross-sectional dispersion. In addition to the interquartile range, we sort months by the standard deviation of price changes. The interquartile range is more robust to outliers, so we view it as a more reliable benchmark, but using the standard deviation does not change our results. We also split our sample using Census based measures of cross-sectional TFP dispersion from Bloom, Floetotto, Jaimovich, Saporta-Eksten, and Terry (2012). When splitting by census based dispersion measures our results become even more significant, with estimated MRPT more than quadrupling between low and high dispersion months. Finally, Table A2 shows that the difference between high and low dispersion results remains significant when using alternative country restrictions as well as when restricting to a more narrow set of products.

We next show that our month-level relationships show up even when using purely aggregate data to estimate pass-through. Our benchmark pass-through specifications calculate the average response of individual prices to exchange rate movements, but there is a long-literature computing aggregate pass-through by looking at the relationship between the aggregate import price index and aggregate exchange rate measures. Once we have measured month-level dispersion using microdata, we can include it as a control in a purely aggregate pass-through specification. To show that month-level dispersion still matters, we implement a natural generalization of our baseline continuous medium-run pass-through regression. In particular, we estimate:

$$\Delta p_t = \alpha + \sum_{j=0}^{2} \beta_j^{AVE} \Delta e_{t-j} + \sum_{j=0}^{2} \beta_j^{VOl} \left(VOL_t \times \Delta e_{t-j} \right) + Z'_{j,t}\gamma + \varepsilon_t$$

where Δp_t is the ex-petroleum import prices, Δe_{t-j} is the trade-weighted exchange rate and $Z'_{j,t}$ is a vector of controls. We show results for two measures of month-level volatility: the cross-sectional standard deviation of price changes (XSD) and the interquartile range (IQR). $\sum \beta_j^{Vol}$ is the main object of interest as it shows how increases in volatility affect exchange rate pass-through. The results are shown in Table A3 and in all specifications, $\sum \beta_j^{Vol}$ is both economically and statistically significant. Increasing the IQR (XSD) by one standard deviation increases aggregate pass-through by 50% (80%) relative to average pass-through in our baseline specification. These results show that even if one is interested in running only aggregate regressions, there is significant time-variation in pass-through and that in order to accurately measure the level of pass-through at a point in time one must condition these regressions on the level of month-level volatility (which requires micro data). In addition Table A3 also shows results for similar aggregate pass-through regressions where we also control for frequency as well as for a linear time trend and quarter dummies to control for seasonality. This does not change the conclusion that aggregate pass-through rises with month-level dispersion.

7.3 More on Month-Level and Item-Level Dispersion Relationship

In the body of the paper we discussed concerns that our month-level and item-level dispersion facts might not actually be distinct since we do not have a balanced panel. Table 5 showed results for interaction specifications that simultaneously controlled for item-level and month-level dispersion to show that both are jointly significant. However, that specification imposes linearity in the effects of both item-level and month-level dispersion and it did not allow for the effects of controls to differ with the level of dispersion. We now take a different approach to argue that the month-level and item-level dispersion facts are indeed two facts and not one. In particular, we run a "double-binned" regression as in Figures (2) and (3). We split individual items into quintiles by their item-level dispersion of price changes, and then within each item-level quintiles we run a time-series regression to estimate the effect of month-level dispersion. Table A4 shows that pass-through is increasing in item-level dispersion during both high and low month-level dispersion periods. In addition, within item-level dispersion bins there is a significant time-series relationship between month-level dispersion and pass-through. These regressions show that simple changes in sample composition cannot jointly explain both facts.

7.4 Long-run Pass-through Relationships

In this section, we examine the relationship between long-run pass-through (LRPT) and item-level volatility. As described in the main text, to compute LRPT we regress the cumulative change in the price of an item of its life in the IPP sample, referred as its life-long price change, on the cumulative exchange rate movement over the same period. Studying the relationship between LRPT and item-level dispersion is important because our argument that greater responsiveness should lead to greater price change dispersion and exchange rate pass-through is not specific to MRPT. Firms that are more responsive should have greater price change dispersion and greater pass-through at all horizons, so if we did not see an empirical relationship between item-level dispersion and LRPT, it would be problematic for our theoretical model.

In addition, despite our focus on time-variation in pass-through at business cycle frequencies, there is a huge literature arguing that LRPT is interesting in its own right. For example, Gopinath and Itskhoki (2010) show that there is a strong relationship between LRPT and the frequency of adjustment. They also explain this relationship using variation in responsiveness, so finding no relationship between dispersion and LRPT would also be problematic for their theory. However, we now show that, reassuringly, there is indeed a strong relationship between item-level dispersion and LRPT.

We investigate the relationship between LRPT and item-level dispersion in two different ways. First, we take a non-parametric approach and examine the relationship between LRPT and item-level price change dispersion within quintiles. The results are shown in Figure 11. The blue line shows the strong positive relationship between price dispersion and LRPT. LRPT increases from 8% to 47% as one moves from the lowest to highest quintile of price dispersion. In addition, the green line confirms the result in Gopinath and Itskhoki (2010) that there is a strong positive relationship between frequency and LRPT.

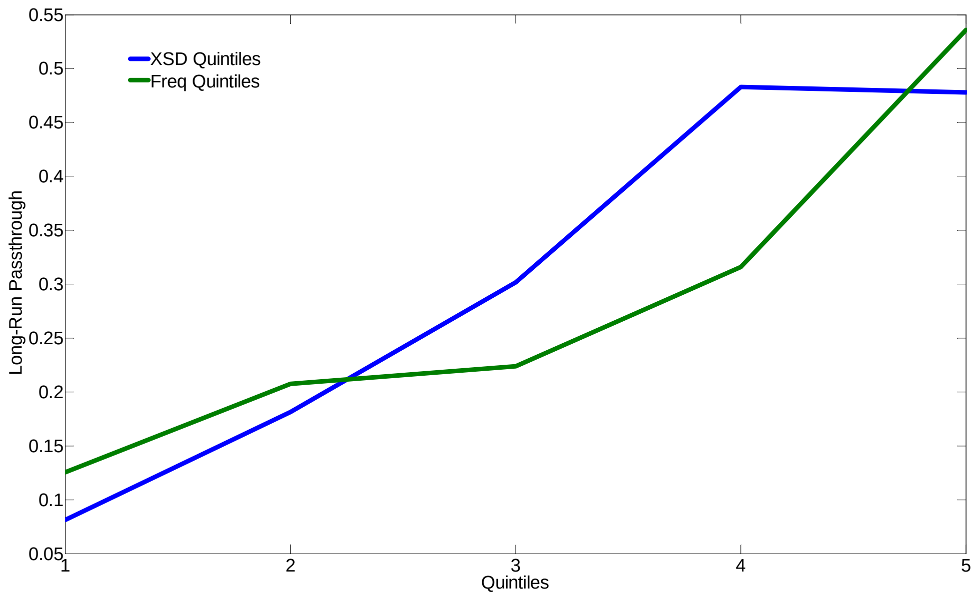
Second, we also take a more structured approach and estimate the long-run equivalent of equation (3), our continuous MRPT regression:

$$\Delta p_i = \beta^{avg}\Delta_c e_i + \beta^{Vol}\left(Vol_i \times \Delta_c e_i\right) + \delta Vol_i + Z_i'\gamma + \epsilon_i \tag{13}$$

The coefficient β^{avg} captures the average long-run pass-through in the sample and β^{Vol} estimates the effect of price change dispersion on long-run pass-through. The results are shown in Table A5. In all specifications, the measure of item level price dispersion is the standard deviation of price changes (XSD) and robust standard errors are clustered by country and PSL pair.

The first two rows show the results for our baseline sample which includes all countries and all items excluding petroleum products. Average exchange rate pass-through is 28%, which is significantly higher than what we found for MRPT. This result is indirect evidence for the presence of strategic complementarities as it suggests items respond more fully to exchange rate shocks in

Figure 11: The relationship between long-run passthrough and frequency/XSD



the long-run. β^{Vol} is greater than zero, which means that items with higher price dispersion have higher LRPT. This is true (and typically highly significant) across all specifications, including ones where we control for the item level frequency of adjustment. The price dispersion effect is economically meaningful: a one standard deviation increase in item-level price change dispersion implies a 43% (0.12/0.28) increase in average LRPT in our baseline sample. Again consistent with Gopinath and Itskhoki (2010), we also find a strong relationship between LRPT and the frequency of price adjustment. In row 3, we include both dispersion and frequency and show that both variables are significantly related to LRPT. Interestingly, the estimated size of the volatility effect is of similar magnitude to the estimated frequency effect. This suggests that variation in item-level volatility also explains a significant amount of the variation in LRPT. The last 6 rows show robustness checks that restrict the analysis to OECD countries and manufactured items. In both samples, the results are similar to our baseline results.

8 Modeling Appendix - Not For Publication

8.1 More General Flexible Price Results

In this section, we show that the intuition from our simple framework in Section 2, survives in a more general framework that allows for general equilibrium effects. Consider the problem of a foreign firm selling goods to importers in the U.S. The firm has perfectly flexible prices that are set in dollars. The optimal flexible price of good i at the border (in logs) can be written as the sum of the gross markup (μ_i), the dollar marginal cost (mc_i) and an idiosyncratic shock (ϵ_i):

$$p_i = \mu_i + mc_i(e_i,\eta_i)$$

Taking the total derivative of equation gives:

$$\Delta p_i = -\Gamma_i(\Delta p_i - \Delta p) + \alpha \Delta e_i + \Delta \eta_i$$

which can be rearranged to give:

$$\Delta p_i = \frac{1}{1+\Gamma_i}\left[\alpha \Delta e_i + \Gamma_i \Delta p + \Delta \eta_i\right]$$

In Section 2 we explored the case when all indirect GE effects were shut off ($\Delta p = 0$). Here, we include them to show that most of the simple intuition between about the positive relationship between MRPT and dispersion survives the introduction of GE effects. The above equation can be rearranged to give the simple pass-through equation:

$$\frac{\Delta p_i}{\Delta e_i} = \frac{\alpha_i}{1+\Gamma_i} + \frac{\Gamma_i}{1+\Gamma_i}\frac{\Delta p}{\Delta e_i} \tag{14}$$

We can do some comparative statics to see how parameters affect pass-through

$$\frac{\partial \frac{\Delta p_i}{\Delta e_i}}{\partial \alpha} = \frac{1}{1+\Gamma_i} > \ 0$$

$$\begin{aligned} \frac{\partial \frac{\Delta p_i}{\Delta e_i}}{\partial \Gamma_i} &= -\frac{\alpha_i}{(1+\Gamma_i)^2} + \frac{1}{(1+\Gamma_i)^2}\frac{\Delta p}{\Delta e_i} \\ &= \frac{\frac{\Delta p}{\Delta e_i} - \alpha_i}{(1+\Gamma_i)^2} < 0 \text{ if } \alpha_i > \frac{\Delta p}{\Delta e_i} \end{aligned} \tag{15}$$

As before, an upper bound on the level of pass-through is given by what fraction of marginal costs are denominated in units of the foreign currency, α_i. The higher this share, the higher the potential exchange rate pass-through. General equilibrium effects operating through the domestic price level do affect the comparative static with respect to the mark-up elasticity. All of things equal, if the mark-up elasticity is higher, then less of the exchange rate shock is passed into prices, which lowers $\frac{\Delta p_i}{\Delta e_i}$. This is the first term in equation (15). However, this is now an additional effect: a higher Γ_i means that individual prices are more sensitive to changes in the aggregate price level because strategic complementarities are higher. This is the second term in equation (15). This term is positive because $\frac{\Delta p}{\Delta e_i} > 0$ since increases in foreign marginal costs also raise the domestic price level. The total effect is ambiguous in general. However, for realistic cases (for instance all the parameter values we consider in our model), $\alpha_i > \frac{\Delta p}{\Delta e_i}$. To see this, remember that α_i is the fraction of marginal cost that is denominated in foreign currency. This gives an upper bound on the level of pass-through to individual prices from exchange rate shocks. It is hard to see how pass-through to the overall price level can be bigger than that effect since not all goods domestically are affected by the exchange rate shock and the overall-passthrough rate is affected by the level of strategic complementarities, Γ_i, which lowers the level of pass-through.

We now show that changes in parameters that increase pass-through also increase the variance of price changes. The variance of price changes is given by:

$$\begin{aligned} var(\Delta p_i) &= \left(\frac{\alpha_i}{1+\Gamma_i}\right)^2 var(\Delta e_i) + \left(\frac{\Gamma_i}{1+\Gamma_i}\right)^2 var(\Delta p) + \left(\frac{1}{1+\Gamma_i}\right)^2 var(\Delta \eta_i) \\ &+ \frac{\alpha_i \Gamma_i}{(1+\Gamma_i)^2} cov(\Delta e_i, \Delta p) + \frac{\alpha_i}{(1+\Gamma_i)^2} cov(\Delta e_i, \Delta \eta_i) + \frac{\Gamma_i}{(1+\Gamma_i)^2} cov(\Delta p, \Delta \eta_i) \end{aligned}$$

But the last terms are zero by assumption that idiosyncratic shocks are orthogonal to exchange rate shocks and will wash out in aggregate so that they do not affect the aggregate price level. This implies that

$$var(\Delta p_i) = \left(\frac{\alpha_i}{1+\Gamma_i}\right)^2 var(\Delta e_i) + \left(\frac{\Gamma_i}{1+\Gamma_i}\right)^2 var(\Delta p) + \left(\frac{1}{1+\Gamma_i}\right)^2 var(\Delta \eta_i) + \frac{\alpha_i \Gamma_i}{(1+\Gamma_i)^2} cov(\Delta e_i, \Delta p) \tag{16}$$

Using this expression, we get that

$$\frac{\partial var(\Delta p_i)}{\partial \Gamma_i} = -\frac{2\alpha_i^2}{(1+\Gamma_i)^3} var(\Delta e_i) + \frac{2\Gamma_i}{(1+\Gamma_i)^3} var(\Delta p) - \frac{2}{(1+\Gamma_i)^3} var(\eta_i) + \frac{\alpha_i(1-\Gamma_i)}{(1+\Gamma_i)^3} cov(\Delta e_i, \Delta p). \tag{17}$$

We now show that under a mild and empirically realistic restriction, the variance of price changes is declining in Γ_i. Empirically, we know that the variance of idiosyncratic price changes is an order of magnitude larger than the variance of aggregate price changes and exchange rate movements. With this in mind, we impose the restriction that

$$var(\Delta p_i) > var(\Delta e_i) + var(\Delta p).$$

We can substitute this restriction into (16) to get that

$$\left(\frac{\alpha_i}{1+\Gamma_i}\right)^2 var(\Delta e_i) + \left(\frac{\Gamma_i}{1+\Gamma_i}\right)^2 var(\Delta p) + \left(\frac{1}{1+\Gamma_i}\right)^2 var(\Delta \eta_i) + \frac{\alpha_i \Gamma_i}{(1+\Gamma_i)^2} cov(\Delta e_i, \Delta p) > var(\Delta e_i) + var(\Delta p)$$

or

$$var(\eta_i) > \left[(1+\Gamma_i)^2 - \Gamma_i^2\right] var(\Delta p) + \left[(1+\Gamma_i)^2 - \alpha_i^2\right] var(\Delta e_i) - \alpha_i \Gamma_i cov(\Delta e_i, \Delta p) \tag{18}$$

Using (17) we have

$$
\begin{aligned}
\frac{\partial var(\Delta p_i)}{\partial \Gamma_i} &= -\frac{2\alpha_i^2}{(1+\Gamma_i)^3} var(\Delta e_i) + \frac{2\Gamma_i}{(1+\Gamma_i)^3} var(\Delta p) - \frac{2}{(1+\Gamma_i)^3} var(\eta_i) + \frac{\alpha_i(1-\Gamma_i)}{(1+\Gamma_i)^3} cov(\Delta e_i, \Delta p) \\
&\propto -2\alpha_i^2 var(\Delta e_i) + 2\Gamma_i var(\Delta p) - 2var(\eta_i) + \alpha_i(1-\Gamma_i) cov(\Delta e_i, \Delta p)
\end{aligned}
$$

Substituting the inequality (18) for $var(\eta_i)$ gives

$$
\begin{aligned}
\frac{\partial var(\Delta p_i)}{\partial \Gamma_i} &< -2\alpha_i^2 var(\Delta e_i) + 2\Gamma_i var(\Delta p) + \alpha_i(1-\Gamma_i) cov(\Delta e_i, \Delta p) \\
&\quad -2\left[(1+\Gamma_i)^2 - \Gamma_i^2\right] var(\Delta p) - 2\left[(1+\Gamma_i)^2 - \alpha_i^2\right] var(\Delta e_i) + 2\alpha_i \Gamma_i cov(\Delta e_i, \Delta p) \\
&= -2\left[(1+\Gamma_i)^2 - \Gamma_i^2 - \Gamma_i\right] var(\Delta p) - 2\left[(1+\Gamma_i)^2\right] var(\Delta e_i) + \alpha_i \left[\Gamma_i + 1\right] cov(\Delta e_i, \Delta p) \\
&< -2\left[(1+\Gamma_i)^2 - \Gamma_i^2 - \Gamma_i\right] var(\Delta p) - 2\left[(1+\Gamma_i)^2\right] var(\Delta e_i) + \alpha_i \left[\Gamma_i + 1\right] var(\Delta e_i) \\
&< -2\left[(1+\Gamma_i)^2 - \Gamma_i^2 - \Gamma_i\right] var(\Delta p) - 2\left[(1+\Gamma_i)^2\right] var(\Delta e_i) + (1+\Gamma_i)^2 var(\Delta e_i) \\
&= -2\left[(1+\Gamma_i)^2 - \Gamma_i^2 - \Gamma_i\right] var(\Delta p) - \left[(1+\Gamma_i)^2\right] var(\Delta e_i) \\
&< 0
\end{aligned}
$$

The second inequality uses the result that Δp moves less than one for one with the exchange rate.

In sum, even in the case when indirect GE effects are allowed, our central theoretical prediction still holds: changes in parameters that increase exchange rate pass-through ($\alpha_i \uparrow$, $\Gamma_i \downarrow$) also increase the variance of price changes.

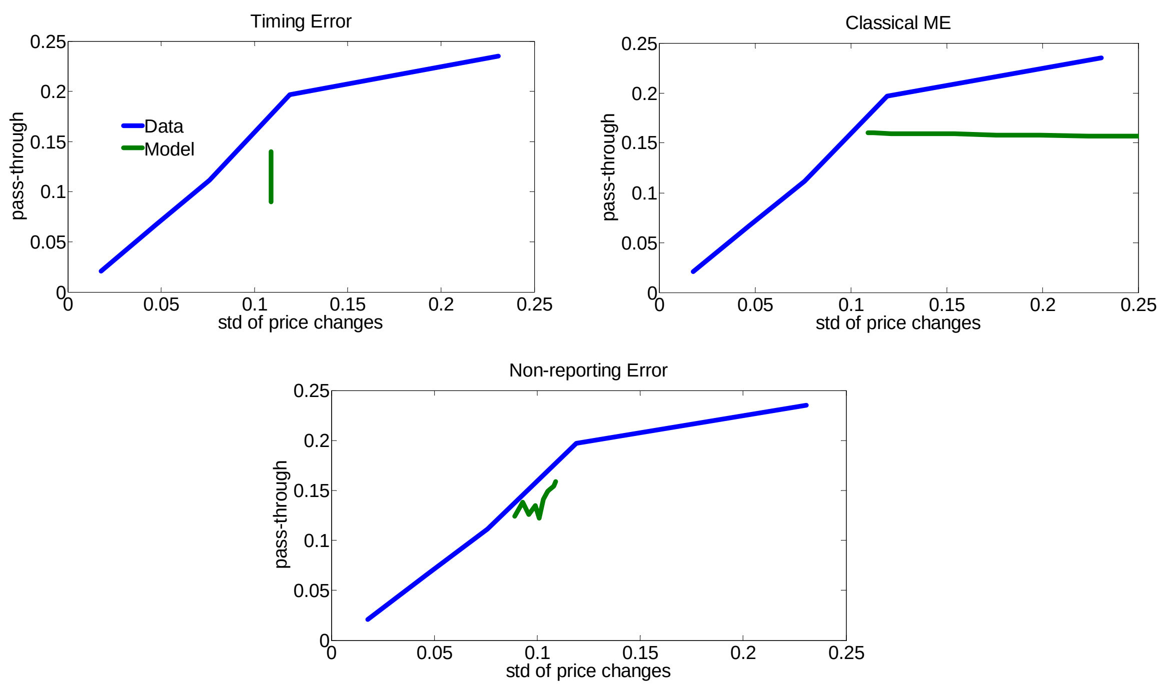
8.2 The Role of Measurement Error

As mentioned in the introduction as well as in Section 3.5, measurement error is a potential concern for our empirical estimates. We attempted to address this concern in our empirical results by estimating various alternative pass-through specifications. While time-series variation in these alternative pass-through specifications is less interpretable as time-variation in price flexibility, these specifications have the advantage of reducing measurement error. Since time-series variation in our benchmark MRPT is more easily interpretable, we now assess the extent to which measurement error is indeed a serious concern for this empirical specification. To do this, we use our model to simulate three sources of potential measurement error and show that such errors cannot explain our results.

We model three sources of measurement error that are likely to be important in the BLS data: 1) Errors in aligning the timing of measured price changes with the timing of exchange rates. 2) Mis-reported prices. 3) Failure to report actual price changes.

Prices are recorded in the BLS at the time they are received rather than at the time they are ordered. Production and delivery lags mean that this price may have been set several periods in the past, under a different prevailing exchange rate.[49] To model this timing error, we assume that while the price at time t is set using information on the exchange rate at time t, the price is reported at time $t+x$ where $x \sim U[0, X]$. That is, there is a potential mismatch between the exchange rate that is actually relevant for a firm's pricing decision and the exchange rate at the time a price is reported. The left hand column of Figure 12 shows the effects of timing errors on pass-through and price change dispersion as X is varied between 0 and 6 months. As X increases, measured pass-through falls as there is additional attenuation bias in the MRPT regression. However, measured price change dispersion is not affected. This is because mismeasuring the timing of price changes has no effect on their measured size.

Figure 12: Simulating Sources of Measurement Error



Thus, changes in timing errors could only explain the time-series relationship between price change dispersion and pass-through if there was some common factor that increased the dispersion of price changes at the same time that delivery lags fall. We can roughly assess this possibility by examining the composition of trade across time. Using data from the U.S. Census, we can compute the fraction of goods shipped by ocean vessels. These items are likely to have the longest delivery lags, so it would be concerning if the fraction of items shipped by vessel negatively comoved with the dispersion of price changes. However, we find that there is a positive correlation of 0.13 between the fraction of items shipped by ocean vessel and the month-level interquartile range of price changes. Thus, if anything, changes in the composition of trade across time would work against our empirical

[49] See Nakamura and Steinsson (2012) for additional discussion.

results. In the appendix we provide additional discussion of trade composition and evidence that this does not drive our results.

In addition to timing error, we allow for reporting errors by assuming that recorded price changes are equal to the true price plus classical measurement error. The second column of Figure 12 shows results for measurement error standard deviations ranging from 0 to 0.18. Increases in measurement error can dramatically increase the dispersion of measured price changes. However, greater measurement error leads to a decline in pass-through due to standard attenuation bias. Thus, classical measurement error is unable to explain our results.

Finally, we assume that when a price change actually occurs it is only recorded with some probability < 1. The third column of Figure 12 simulates results for non-reporting probabilities ranging from 0 to 0.9. Even huge non-reporting errors barely affect either pass-through or price change dispersion. They do not affect pass-through because once a price change is actually measured, it will reflect all of the previous pass-through that was not recorded. Furthermore, non-reporting does not affect the dispersion of price changes as long as the probability of a price change not being reported is independent of the size of the price change. The one statistic that declines dramatically with non-reporting error is the frequency of adjustment. Thus, if non-reporting error were a cause of concern for our results, this explanation would need to contend with the much smaller frequency pass-through relationships observed empirically.

8.3 Quantitative Calvo Model

In the body of the text, we presented a menu cost version of our quantitative model. We focus on this model because it is a better fit to the micro data, but it is straightforward to solve a Calvo version of our benchmark model. This appendix shows that our main results are robust to the underlying price-setting mechanism. The basic setup is identical to the menu cost model in the body of the text except that firms face a Calvo (1983) style friction: the firm is allowed to adjust prices each period with an exogenous probability $(1-\lambda)$. Define the state vector of firm j by $S_{jt} = (P_{j,t-1}, A_{jt}; P_t, W_t, W_t^*)$ where $P_{j,t-1}$ and A_{jt} are the idiosyncratic state variables and P_t, W_t, and W_t^* are the aggregate state variables. The value of the firm selling variety j is characterized by the following Bellman equation:

$$\begin{aligned} V(S_{jt}) \;\; = \;\; & (1-\lambda)(\max_{P_{jt}} [\Pi_{jt} + E\{Q(S_{jt+1})V(S_{jt+1})\}]) \\ & + \lambda\beta(\Pi_{jt}(P_{jt-1}) + E\{Q(S_{jt+1})V(S_{jt+1}|P_{j,t-1})\}) \qquad (19) \end{aligned}$$

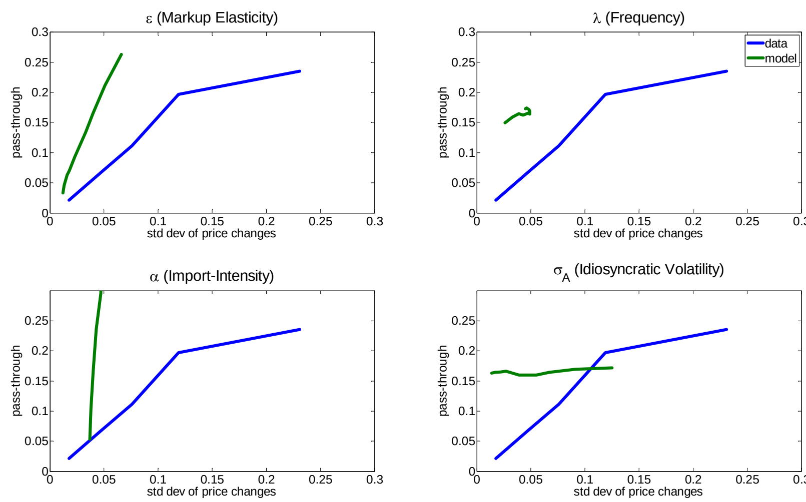
where $Q(S_{jt+1})$ is the subjective discount factor. The interpretation of equation (19) is intuitive. With probability $(1-\lambda)$ the firm changes its price (possibly keeping it the same) and moves onto the next period. With probability λ the firm is unable to change its price so it earns flow profits with its previous price and starts next period with the same price.

8.3.1 Calibration and Results

The overall calibration strategy is identical to the menu cost model. Most of the parameters are calibrated to the same value as in that model, but we find that the best fit α is somewhat larger at 0.25 vs 0.18 in the menu cost model. This is because without selection effects, there is no covariance term that drives up measured pass-through, so generating the same level of MRPT requires greater sensitivity to the exchange rate.

As in the comparative statics in the paper, each panel of Figure 13 shows what happens when we fix three of $\varepsilon, \lambda, \alpha$ and σ_A at their steady state values and vary the fourth parameter. For the sake of comparison, the empirical relationship between the standard deviation of price changes and MRPT that we documented in the IPP is shown by a blue line.

Figure 13: Calvo Comparative Statics



The top-left panel of Figure 13 shows the results from varying ε from 0 to 40. First, observe that variations in ε do indeed generate a positive relationship between the variance of price changes and MRPT. Qualitatively, a low ε implies high price change variance and high MRPT and increasing ε increases the curvature of firm's profit function, thus lowering both pass-through and the variance of price changes. Quantitatively, the slope of this relationship in the model is too high. Another way of putting this is that the variation in ε is unable to generate enough variation in the variance of price changes. This is not surprising because even with $\varepsilon = 0$ (CES case), the Calvo model has a very difficult time generating a large variance of price changes.

The bottom-left panel shows what happens when we vary α from 0 to 1. As in the menu cost model, this leads to large changes in MRPT but negligible movements in the variance of price changes.. Thus, variation in ε is better able to replicate the positive relationship between MRPT and the standard deviation of price changes; however, the fit is still not very good.

The top and both right side panels shows what happen when we vary the frequency of price change and the variance of idiosyncratic shocks, respectively. Variation in λ from 0 to 1 generates essentially no variation in either MRPT or the variance of price changes, whereas variation in σ_A from 0 to 0.2 generates some variation in the standard deviation of price changes but almost no variation in MRPT. Overall, as in the menu cost model, variation in ε provides the best quantitative match to the strong, positive relationship between the standard deviation of price changes and MRPT that we documented in U.S. import data. However, the Calvo model struggles to generate the wide variation in price change dispersion observed in the data. This is because firms don't want to make large price changes and get stuck with the wrong price. For this reason, we have focused on formal estimation only for the menu cost model.

Appendix Tables

Table A1:Interaction Specification - Cross-Item: Robustness

	Average pass-through		Volatility		Frequency		N_{obs}	R^2
	β^{avg}	$se(\beta^{avg})$	β^{Vol}	$se(\beta^{Vol})$	β^{freq}	$se(\beta^{freq})$		
At least 3 price changes	0.15	0.02	0.05	0.02			88964	0.06
	0.15	0.01	0.05	0.02	0.01	0.01	88964	0.06
Using trade-weighted broad xrate	0.41	0.03	0.26	0.04			87383	0.08
	0.44	0.03	0.21	0.04	0.27	0.03	87383	0.08
Using trade-weighted major country xrate	0.28	0.02	0.21	0.03			96512	0.07
	0.29	0.02	0.18	0.03	0.15	0.02	96512	0.07
Placebo num changes	0.15	0.01	0.00	0.01			100871	0.09
	0.15	0.02	-0.00	0.01	0.02	0.01	100871	0.09
Placebo num obs	0.15	0.02	-0.00	0.01			100871	0.09
	0.15	0.02	-0.00	0.01	0.02	0.01	100871	0.09
Median regression	0.16	0.00	0.07	0.00			95284	
	0.16	0.00	0.07	0.00	0.01	0.00	95284	

Note: Robust standard errors clustered by country*PSL pair. Primary strata lower (PSL), defined by the BLS, represents 2 to 4-digit sectoral harmonized codes

Table A2: "Binned" time series results

	High volatility		Low volatility		Difference		N_{obs}	R^2
	β^{high}	$se(\beta^{high})$	β^{low}	$se(\beta^{low})$	$\beta^{high}-\beta^{low}$	t-stat		
All countries, all items ex petroleum								
- Interquartile range	0.21	0.03	0.08	0.02	0.12	4.35	62395	0.09
- Cross-sectional std	0.17	0.02	0.08	0.02	0.10	3.89	63095	0.09
- Bloom uncertainty	0.26	0.03	0.06	0.02	0.20	6.33	64204	0.08
OECD, all items ex petroleum								
- Interquartile range	0.22	0.03	0.12	0.03	0.10	2.88	34790	0.09
- Cross-sectional std	0.22	0.03	0.10	0.03	0.12	3.17	35288	0.09
- Bloom uncertainty	0.25	0.03	0.07	0.03	0.18	4.37	35398	0.09
All countries, all manufact. items								
- Interquartile range	0.18	0.03	0.08	0.02	0.10	3.41	51157	0.11
- Cross-sectional std	0.15	0.02	0.08	0.02	0.07	2.75	51943	0.11
- Bloom uncertainty	0.23	0.02	0.05	0.02	0.18	5.83	52661	0.10

Note: Robust standard errors clustered by country*PSL pair. Primary strata lower (PSL), defined by the BLS, represents 2 to 4-digit sectoral harmonized codes

Table A3: Aggregate pass-through

Month-Level Dispersion	Controls	β_0^{AVE}	t	$\sum \beta^{AVE}$	t	β_0^{Vol}	t	$\sum \beta^{Vol}$	t	n
IQR	None	0.14	3.43	0.31	4.59	0.07	1.19	0.19	2.20	71
	Freq	0.13	2.85	0.28	3.81	0.03	0.50	0.18	2.05	71
	Time/Qrt.	0.21	1.13	0.36	1.74	0.07	1.13	0.21	2.38	71
XSD	None	0.15	4.30	0.31	5.99	0.12	3.10	0.18	3.32	71
	Freq	0.15	2.37	0.31	5.31	0.10	2.02	0.17	2.56	71
	Time/Qrt.	0.21	1.33	0.32	1.84	0.13	2.28	0.21	3.50	71

Table A4: Month-Level IQR Regressions by Item-Level XSD Quintiles

Item-Level XSD Quintile	$\beta^{high\ IQR}$	$\beta^{low\ IQR}$	$\beta^{high\ IQR} - \beta^{low\ IQR}$	t-stat	n	R^2
1 (Lowest)	.035	.029	.005	0.29	6096	.64
2	.083	.052	.031	1.46	12522	.24
3	.133	.053	.079	2.24	16630	.15
4	.277	.127	.150	3.41	16470	.13
5 (Highest)	.417	.112	.304	2.92	10942	.12

Table A5: Interaction Specifications - LRPT

	Average pass-through		Volatility		Frequency		N_{obs}	R^2
	β^{avg}	$se(\beta^{avg})$	β^{Vol}	$se(\beta^{Vol})$	β^{freq}	$se(\beta^{freq})$		
All countries, all items ex petroleum								
- Dispersion	0.28	0.04	0.12	0.04			13962	0.49
- Frequency	0.27	0.04			0.12	0.05	13962	0.49
- Dispersion and Frequency	0.26	0.03	0.09	0.04	0.10	0.04	13962	0.50
OECD countries, all items ex petroleum								
- Dispersion	0.32	0.05	0.10	0.04			6549	0.52
- Frequency	0.33	0.05			0.12	0.06	6549	0.52
- Dispersion and Frequency	0.32	0.05	0.09	0.04	0.11	0.06	6549	0.52
All countries, all manufacturing items								
- Dispersion	0.24	0.04	0.07	0.04			12506	0.50
- Frequency	0.23	0.03			0.06	0.04	12506	0.50
- Dispersion and Frequency	0.24	0.04	0.06	0.04	0.06	0.04	12506	0.50

Note: Robust standard errors clustered by country*PSL pair. Primary strata lower (PSL), defined by the BLS, represents 2 to 4-digit sectoral harmonized codes