Stickiness, Synchronization, and Passthrough in Intrafirm Trade Prices

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First Version: November 2007

This Version: December 2009

Abstract

About forty percent of all U.S. international trades occurs between related parties, or intrafirm, such as trades between a parent and subsidiary of the same multinational corporation. This paper uses a transaction-level dataset that distinguishes arm's length from intrafirm trades to demonstrate that for differentiated products, intrafirm prices are characterized by 1) less stickiness, 2) less synchronization, and 3) greater exchange rate passthrough.

1 Introduction

Between a third and a half of all U.S. trade occurs between related parties, or intrafirm, such as trades between a parent and subsidiary of the same multinational corporation. Despite the striking quantitative significance of global vertical integration and the many studies focusing on its determination, remarkably few papers have measured the implications for a firm's pricing decisions and for the macroeconomic environment at the sector or country levels. Empirical work with international prices often ignores heterogeneity in vertical structure or simply excludes intrafirm trade prices on the presumption that they are not economically meaningful. This paper's contribution is to document important differences in the dynamic behavior of intrafirm prices.

Analysis of transaction-level import data collected by the U.S. Bureau of Labor Statistics (BLS) demonstrates that for the set of differentiated products, intrafirm prices are characterized by 1) shorter price spells, 2) a lower degree of synchronization, and 3) greater long-run exchange rate passthrough.[1] The first two observations on duration and synchronization are novel to the literature, and the good-level differences in duration aggregate up to macroeconomic levels such as sectors or countries. The result of larger intrafirm passthrough corroborates previous findings from indirect calculations or aggregated data, but uses micro-data that allows for direct passthrough regressions run separately on arm's length and intrafirm prices.

First, the typical price spell of a differentiated good lasts about 3 months, or between 20 and 30 percent, longer for arm's length trades than for related party trades. This difference is not concentrated in a few sectors or countries. For example, the median intrafirm duration is shorter in 20 of the 23 2-digit SITC sectors and in 17 of the 22 countries with duration es-

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[1]The term "differentiated products" refers to the 75-80 percent of all goods that Rauch (1999) identified as neither trading on an organized exchange nor with a reference price. Section 2 discusses the reasons for focusing on differentiated goods in this paper.

timates on at least 100 differentiated goods. Good-level price duration is important because, given an invoicing currency, it is the key determinant of short- and medium-run responsiveness to shocks. Hence, price stickiness impacts the relationship between trade balances and exchange rate movements, determines the impact on other economies of a country's monetary policy, and may explain short-term deviations from the law of one price. Even setting aside these cross-border effects, input price stickiness remains important as perhaps the singularly significant observable determinant of final good price stickiness.[2]

Second, the timing of changes in these prices is less synchronized among related party exporters than arm's length exporters. On average, a one standard deviation increase in the share of competitors with the same vertical structure that are increasing prices raises the probability of an arm's length price increase by 33 percent, compared to 23 percent for related parties. For price decreases, the impact on arm's length firms is 59 percent, compared to 49 percent for related parties. This implies that price changes by competitors elicit responses from arm's length firms, but are less likely to provoke related party price changes. The relative disconnect in the timing of intrafirm price changes may help explain the disconnect in the times that upstream integrated manufacturers hold and invest in capacity compared with arm's length manufacturers in the same industry, documented empirically in Mullainathan and Scharfstein (2001).

Third, more precise and direct tests corroborate and quantify previous findings of higher exchange rate passthrough in intrafirm prices.[3] In the baseline long-run specification, intrafirm passthrough in differentiated products is about 30 percent, 8 percentage points larger than for arm's length prices. Although both passthrough rates decline when restricting the

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[2]Goldberg and Hellerstein (2008), for instance, report that in all of their dataset on the beer market, there is never an instance when the wholesale price changes but the retail price does not.

[3]Rangan and Lawrence (1993) and Hellerstein and Villas-Boas (forthcoming) find larger passthrough in highly aggregated sectors that had a larger share of intrafirm trades. Bernard et al. (2006) find a relationship between the gap in intrafirm and arm's length prices and the exchange rate that implies larger intrafirm passthrough.

analysis to dollar-priced goods, the size and significance of this gap remains intact, confirming that greater intrafirm passthrough does not merely reflect currency composition. The scale of the difference varies, but the estimate of differential passthrough for intrafirm trades is always positive, and it is statistically significant across the large majority of regression specifications. Hence, these results contribute to an extensive literature on cross-country and cross-sector heterogeneity in exchange rate passthrough (e.g. Campa and Goldberg, 2005, or Yang, 1997).

Empirical work often excludes intrafirm transactions due to the concern that transfer prices are not allocative. This paper cannot directly prove otherwise because the BLS data do not include transaction quantities. The presumption that intrafirm prices are primarily accounting constructs, however, might carry with it the expectation that transfer prices change less frequently and are less tied to fundamentals such as the exchange rate. It is noteworthy that these two common priors about intrafirm transactions are at odds with the data.

Further, this paper offers direct evidence against the hypothesis that these differences in intrafirm price changes are primarily driven by the desire to shift a firm's taxable income to countries with the lowest tax rates. Patterns in duration and exchange rate passthrough do not meaningfully differ when imports are sourced from countries with tax rates similar to the U.S. compared with countries with highly dissimilar tax rates.

Finally, it is worth noting that the topic of intrafirm trade is just as salient when the trading relationship does not cross international borders. Many domestic trades are conducted within the boundary of the firm. The international context is only critical for data availability, including the ability to observe exchange rate driven cost shocks. Although the nature and correlation structure of cost shocks may differ in closed and open economies, the results in this paper suggest that differences in intrafirm good price dynamics exist in both international and domestic settings.

2 Bureau of Labor Statistics' (BLS) Trade Pricing Data

This section briefly describes the data, particularly as it relates to intrafirm trade. Appendix A discusses in far greater detail other characteristics relevant to this analysis. Gopinath and Rigobon (2008) contains an in-depth description of the collection process for this survey data and a detailed discussion of its features and limitations.

The dataset aggregates surveys administered by analysts in the BLS's International Price Program (IPP) from 1993 to 2005 and contains the underlying data used to construct their import price indices. Unlike some datasets used to study transfer prices, such as the U.S. Census Bureau and Customs data used by Bernard et al. (2006), the BLS dataset is not associated with any tax collection authority. BLS analysts explicitly tell companies that their pricing data will be kept confidential and used only for price index construction and research.[4] Accordingly, this data set is uniquely suitable for this analysis: compared to other sources of transfer pricing data, it is less likely to capture non-allocative prices. Section 6 elaborates this statement and offers quantitative evidence suggesting that the empirical differences identified in this paper do not simply reflect the use of transfer prices to minimize tax liability.

From 1993 to 2005, there are data on about 57,000 different imported goods, with over 1.1 million import prices considered "usable" by the BLS. A usable price is generally not imputed and reflects an actual survey response intended to capture the price of a real transaction. Approximately 22,000 of the goods, or about 40 percent, were classified as intrafirm, and more than 446,000 of the usable prices, also about 40 percent, were also classified as intrafirm. The median life of goods in the dataset is a bit less than 3 years and differs by only one month when comparing related party to arm's length trades.

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[4] According to the documentation provided by the BLS, "Using information in identifiable form for any other purpose such as an administrative, regulatory, or law enforcement purpose is considered a non-statistical purpose and is strictly prohibited by law and BLS policy."

I match about 70 percent of these goods at the 4-digit SITC level with Rauch's (1999) classification of traded goods as either differentiated or as having "organized exchanges" or "reference prices." The former group contains, for example, specialized or branded goods, while oil or other commodities fall into the latter. Although the stickiness of non-differentiated prices for related parties decreases relative to differentiated goods, there is a far more drastic decline for the arm's length case.[5] Given that non-differentiated goods are classified in part based on the medium used for trade and that related party transactions are highly unlikely to occur over organized exchanges, it is difficult within this set of goods to ensure appropriate comparisons of otherwise similar arm's length and related party pricing decisions. Consequently, I restrict most of the analyses below to the set of differentiated goods. About 80 percent of both intrafirm and arm's length goods fall into this category, so the overall shares do not change much in the remaining dataset.

One striking fact about intrafirm trade is how pervasive it is, even at very granular levels. One might guess that most industries or countries are dominated by either arm's length or intrafirm trade. In fact, the distribution of the share of intrafirm trade is not bimodal and instead has roughly equal mass across many intermediate percentage values. Figure 1 highlights this fact: it shows the share of all usable intrafirm prices (including non-differentiated) by 2-digit SITC industry and by exporting country over the period 1993-2005. The diameter of each circle is proportional to the share of total imports to the U.S.

The share of intrafirm prices in the dataset has slowly increased from 37 percent in 1997 to 42 percent in 2005. Goods such as wood pulp are mostly traded at arm's length and goods such as specialized semiconductors are mostly traded intrafirm. The bulk of trade, however, falls in the intermediate ranges, including high volume imports like metal manufactures and electrical parts. The distribution across countries also bears little resemblance to a

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[5]For example, the trade-weighted medians for arm's length differentiated and non-differentiated goods are 15 and 3 months, respectively, compared to 12 and 6 for the intrafirm case.

bimodal distribution, with major trading partners such as Canada exporting to the U.S. using a combination of both vertical structures. The propensity for both structures to exist in tandem suggests the importance of studying intrafirm prices for developing a complete understanding of international trade and macro dynamics. Intrafirm trades are pervasive and cannot be ignored by simply excluding a few countries or industries.

3   Intrafirm Trade and Price Stickiness

This section demonstrates empirically that the median differentiated good price spell is shorter for related party prices than for arm's length prices, that this difference holds across most sectors and countries, and that it remains after conditioning on relevant variables. Several methods can be used to measure the duration of prices. Parametric, semi-parametric, and non-parametric estimates of price duration all show that differentiated good prices are less sticky for related parties.

The key difficulty in estimating duration using the BLS data is the large number of censored entries. The baseline procedure for classifying price changes first examines whether the price differs before and after a censored, or missing, price. If the prices surrounding a gap are identical, the price is assumed to be sticky over that unobserved period. This is a reasonable assumption because price changes in these data typically represent permanent departures from the previous price. This is unlike retail price data, where sale prices and other temporary departures from a longer-term reference price are frequently observed (e.g. Eichenbaum et al., 2008, or Nakamura and Steinsson, 2008). When a good enters and exits the dataset, or the BLS documents a change in the good's quality, such breaks in the data series are coded as price changes. A price change is defined as a change in terms of the currency of denomination in order to ensure a closer mapping to standard "menu cost" models, where a fixed cost is involved in changing a price. 12 percent of intrafirm

differentiated prices are non-dollar denominated, compared with 9 percent for arm's length firms.

Following much of the literature, rather than taking an unweighted average of each spell length, the baseline duration estimates use the spell lengths as weights, although results without spell-weighting are also reported to demonstrate robustness.[6] There are myriad potential approaches to classifying price spells and estimating spell-weighted duration that could yield quantitatively meaningful differences in stickiness levels. This paper's focus, however, is on the comparative statics across vertical trade structures rather than the average level itself.

Table 1 shows in greater detail the facts about duration and the vertical structure of trade. Given significant heterogeneity in the data and the fact that many goods have a very short usable life, the baseline results presented here (and in the introduction) use a parametric approach to estimate duration on goods with more than 6 usable prices in the data. The key benefits of this approach include the ability to generate estimates at the good level and, by making assumptions about the likelihood of price changes during censored months, it allows for use of more of the data.

Following Gopinath and Rigobon (2008), it is assumed that each good exhibits a constant hazard rate for price changes, λ_j, which leads to a cumulative distribution function for price spell length of $1 - e^{-\lambda_j S_{i,j}}$, where $S_{i,j}$ is the length of a price spell i of good j. When price spells are right censored, the maximum length that the spell could have lasted is bounded by observing a different price $M_{i,j}$ months after the spell began. These factors are all taken

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[6]A two-good sector, with one good changing its price every two months and the other changing once a year, would have a spell-weighted duration of 7 months and an unweighted duration of about 3.5 months. The same would be true for a single good's price that changed every two months for the first year, but only after 12 months in the second. See Baharad and Eden (2004) and Gopinath and Rigobon (2008).

into account by maximizing the following likelihood function:

$$L(\lambda_j) = -\sum_i S_{i,j}^2 \lambda_j + \sum_{\text{Uncensored Spells}} S_{i,j} \ln(\lambda_j) + \sum_{\text{Right Censored Spells}} S_{i,j} \ln\left(1 - e^{-\lambda_j(M_i - S_i)}\right).$$

Although duration estimates for intrafirm prices are only slightly shorter in the overall dataset, the scale of this difference becomes economically meaningful when considering differentiated goods. Across the cases in Table 1 with and without trade-weighting, with and without spell-weighting, and with and without dollar prices, median arm's length price spells for differentiated goods last between 22 and 30 percent longer than intrafirm prices.[7]

Previous papers, such as Gopinath and Rigobon (2008) and Clausing (2001), have found that arm's length and related party prices have very similar average durations, which may seen inconsistent with these results. The difference is largely driven by this paper's focus on differentiated goods (and good-level estimation) and I am largely able to replicate their results to corroborate this fact. Indeed, the overall trade-weighted arm's length duration estimate of 11.6 falls squarely in the middle of the estimates, from 10.6 to 13.8, listed in Table 1 of Gopinath and Rigobon (2008) for comparable goods. This paper's estimate of 11.0 months for related parties matches closely with their frequency estimate of 9 percent. As such, the discussion in Gopinath and Rigobon of less dispersion in their intrafirm duration estimates implies that lower intrafirm duration for differentiated goods also exists in their data.[8,9] Organized exchange and reference priced goods, though a small share of the data, are

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[7]The dataset does not contain good-level weights throughout the life of the data, though it does contain total weights for each year of each "classification group," which may contain several goods. I generate trade weights by apportioning the classification group weights equally among all goods in a classification group each year and averaging this weight over the life of a good.

[8]Gopinath and Rigobon do list that related party prices are stickier in two of the five primary end-use categories. I replicate this in my overall dataset and find that after restricting to differentiated goods, only one (none) of the five categories demonstrates stickier related party prices with (without) trade-weighting.

[9]Clausing's estimates for 1997-1999 show that related parties are very slightly stickier. These results are less comparable as they are not done good-by-good and hence are highly sensitive to heterogeneity (the implied average durations are about 2 months for both arm's length and intrafirm prices). Nonetheless, I replicated these result using the same methodology and confirmed that after throwing out "non-usable" (such as imputed) prices and restricting the dataset to differentiated goods, the results switch and related

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characterized by dramatically lower arm's length duration. When pooled with differentiated goods, they roughly balance the overall intrafirm and arm's length stickiness statistics.

Figure 2 aggregates these good-level differences and plots the median duration estimates for countries with at least 100 differentiated goods against the share of these trades conducted intrafirm. It demonstrates that low intrafirm share countries, such as India, exhibit far greater stickiness (>19 months) than those with large intrafirm shares, such as Mexico (<13 months). Differences in this figure, of course, reflect both intrafirm and country-level effects. For example, one may posit that the differences in median duration shown in Table 1 and Figure 2 simply reflect country-level differences like the exchange rate. If intrafirm trading of differentiated goods happens to occur with countries where the exchange rate is extremely volatile, one might expect shorter price durations, independent of mechanisms related to a company's vertical structure. One chief benefit of the good-by-good parametric approach is that durations can be covaried with multiple good-level variables to address this concern.

Table 2 shows the estimates from weighted and unweighted regressions of both the log and level of duration (in months) of differentiated good prices on the vertical structure of the trade, the standard deviation of monthly percentage changes in the nominal exchange rate over the life of each good in the dataset, the currency of denomination, and a proxy for the good's elasticity of substitution. As indicated in the bottom rows, the regressions include combinations of country and 2-digit industry dummies. The top row in the trade-weighted and unweighted panels reinforce that for differentiated goods, related party duration is smaller. For instance, in the weighted log specification without dummies, the -0.2036 coefficient indicates that, on average, related party spells are 20 percent shorter, and the weighted level coefficient implies they last about 3.7 fewer months. Both results are largely consistent with the unconditional median durations presented earlier. Though the scale of the coefficient certainly fluctuates across specifications, it is always highly statistically

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parties become slightly less sticky.

significant and remains intact even in the full structure with country and industry dummies.

The regression coefficient on the proxy for exchange rate volatility, surprisingly, is positive, but it is economically small and statistically insignificant across the majority of trade-weighted cases. Dollar denominated goods are significantly stickier than non-dollar goods, though as shown in Table 1, the number of non-dollar goods is so small that it generally matters little for this concept of stickiness. The final columns add the quintile of the sector's elasticity of substitution (ranging from 1 for the least substitutable to 5 for the most) using the Broda and Weinstein (2006) measure. As one would expect, this coefficient is negative – the more elastic goods are less sticky – but insignificant in the trade-weighted case. By conditioning on all these characteristics, as well as absorbing variation across countries and sectors, the specifications in Table 2 demonstrate that the finding of less sticky intrafirm prices applies throughout the set of differentiated goods.

Semi- and non-parametric approaches that are less restrictive about the shape of the hazard function over time confirm the robustness of this result. Each method carries its own costs and benefits but the finding of less sticky related party duration for differentiated goods holds in both. Though the magnitude of the difference shrinks somewhat, the spell-weighted Kaplan-Meier (non-parametric) estimator of duration verifies that related party spells are shorter not only at the median, but also at the 25th and 75th percentiles of the distribution. The Cox proportional hazards model, a semi-parametric estimator of duration, also puts no structure on the functional form of the underlying hazard rate, but it specifies that good characteristics shift this rate proportionately and hence allows for covariation with multiple variables. The magnitude of the intrafirm effect varies with the specification, but runs of the Cox model confirm a statistically significantly lower duration for intrafirm differentiated prices.

In summary, these good-level differences, which were shown in Table 2 to hold conditional on key variables and in Figure 2 to also hold at the exporting country level, are suggestive

that firm boundaries matter for price responsiveness and that vertical integration may be important for explaining aggregated patterns of stickiness.

4 Intrafirm Trade and Price Synchronization

This section demonstrates empirically that related party price changes are less synchronized than arm's length price changes. Fully synchronized price changes would imply that in a given month and sector, either all prices change or none do. A complete lack of synchronization, the other extreme, would imply that all firms' price change decisions are independent of their competitors', and the share of price changes in a sector remains roughly constant. The fact that related party prices are less synchronized is important because it suggests transfer pricing decisions are more inward focused and can produce different behavior at the sector level.

There are few established methods for measuring price synchronization, but this section follows Midrigan (2006) and considers results from an ordered probit of the form:

$$\Pr(Y = 1, 0, -1 | X = x) = \Phi\left(\beta X\right),$$

where Y indicates whether a price has increased, remained the same, or decreased. The key covariate in the vector X is a variable capturing the percentage of other prices within the same SITC 4-digit sector and the same vertical structure (intrafirm or arm's length) that shares the same value as Y. This estimate tries to measure the likelihood that competitors' price changes influence a firm to change its own price. The set of covariates additionally includes the percentage of other firms whose action is the opposite of Y, the cumulative change in the nominal exchange rate since the previous price change, 2-digit sector, country, and month dummies. Regressions only include sector-months with at least 5 observed prices

of each trade type, though little changes when this restriction is dropped.[10]

Table 3 gives the marginal effects for arm's length and related party price increases ($Y = 1$) of the percentage of other firms of the same type that also increase prices. Because each trade type has a different baseline unconditional probability of adjustment, the coefficients are reported as the impact of a one standard deviation change in the covariate relative to this baseline (standard errors scaled accordingly). The coefficient 0.3306 indicates that a one standard deviation increase in the share of other arm's length price increases in the 4-digit SITC sector raises by 33 percent the probability that a given arm's length firm also increases its price. The coefficient 0.2307 indicates that a one standard deviation increase in the share of other related party price increases in the sector raises by 23 percent the probability that a given related party firm also increases its price, a smaller effect. These patterns also hold for the case of price decreases ($Y = -1$), with the differential effect equalling 59 percent for the arm's length case, compared with 49 percent for related parties. In sum, these results suggest a lower degree of synchronization at the 4-digit sector level in related party price changes – increases and decreases – compared to arm's length price changes.

5   Intrafirm Trade and Long-Run Passthrough

Having demonstrated differences in the decision to change prices, I now consider how such changes relate to firms' marginal cost of production. While one cannot observe firms' true cost shocks, a key benefit of international data is that one can observe exchange rate shocks, surely a meaningful component of overall cost shocks. Previous studies using BLS price indices, including Rangan and Lawrence (1999) and Hellerstein and Villas-Boas (forthcoming), found that highly aggregated sectors with larger shares of intrafirm trade exhibit higher passthrough. Bernard et al. (2006) uses census data and considers changes in the gap be-

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[10] There is a "January effect" in the data where most price changes occur in the first month of the year, but it is not strikingly large for either type of trade. The results on synchronization are not materially changed whether including or omitting the month dummies.

tween intrafirm and arms length prices as the exchange rate varies. Though none of these is a direct measurement, the results in all three of these papers suggest intrafirm passthrough is higher. This section considers direct passthrough measurements for intrafirm prices and for arm's length prices over all differentiated import sectors and shows, consistent with those earlier results, that intrafirm passthrough is materially higher. Further, the micro data allows for a quantitative measurement of the difference in passthrough and makes it easy to verify that it is not simply a reflection of the differential composition of dollar and non-dollar prices.

This section presents results using two different methods for measuring passthrough in intrafirm and related party differentiated goods prices. First, estimates are generated from regressions that only include non-zero price changes and that match each price change to the accumulated exchange rate change associated with the preceding price spell. Second, a pooled regression is estimated of all monthly price changes, including zeros, on the concurrent and lagged bilateral exchange rate changes with the exporting country. In the context of censoring and measurement error, each method offers different costs and benefits, but both demonstrate higher passthrough for intrafirm goods. For each estimate, results are reported with and without dollar-denominated prices in order to clarify that currency composition is not driving the results.[11]

5.1 Conditional Passthrough Estimates

The first estimation method measures "conditional passthrough" as the β_1 coefficient from a regression of the form:

$$\Delta \ln p_j^{c,\ t_j - t_j^{-1}} = \alpha + \beta_1 \Delta \ln e_j^{c,\ t_j - t_j^{-1}} + \beta_2 \Delta \ln \pi_j^{c,\ t_j - t_j^{-1}} + \beta_3 \Delta \ln \pi_j^{U.S.,\ t_j - t_j^{-1}} + dummies + \varepsilon_j^{t_j} \quad (1)$$

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[11]Since currency is chosen by firms, however, there is no reason to believe that the dollar-only estimates are in any sense more meaningful for measuring a sector's responsiveness to exchange rate shocks.

where t_j and t_j^{-1} are good specific and respectively denote the times of the most recent and penultimate price changes. Hence, $\Delta \ln p_j^{t_j - t_j^{-1}} = \ln(p_j^{t_j}/p_j^{t_j^{-1}})$ denotes the size of the most recent price change and $\Delta \ln e_j^{t_j - t_j^{-1}} = \ln(e_j^{t_j}/e_j^{t_j^{-1}})$ denotes the accumulated change in the relevant bilateral exchange rate from the time of previous price change to the time of the most recent change, and the next two terms similarly capture changes in the foreign and U.S. consumer price levels. Only non-zero price changes are included in these regressions, which include both country and 2-digit industry dummies.

Table 4 reports the output of these regressions when all differentiated good prices are pooled and an interaction term is added to measure any differential passthrough of related parties. The regressions include all spells, and as above, prices are assumed to be constant during censored periods that are surrounded by the same price. The earliest exchange rate is taken as the base for the first spell, and the BLS methodology is used to measure the price change (theoretically holding quality fixed) during the very limited instances in which a quality change is identified. Even among the data included in BLS price indices, there are some very small and very large changes that often result from input or rounding errors that are difficult to manually remove. As such, I generate these conditional passthrough results when including all price changes, those with absolute sizes ranging from 1 to 50 percent, and those ranging from 1 to 25 percent.

Panel A reports the unweighted results for the real specification (where estimates include the constraint $\beta_1^n = \beta_2^n = -\beta_3^n$) and Panel B reports estimates that additionally use trade weights.[12] Columns (1) through (6) include all prices, while columns (7) through (12) are restricted to dollar-denominated goods. The overall conditional passthrough estimate in column (1) for the unweighted case is about 22 percent, largely in line with recent estimates

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[12]Each price spell is weighted using the trade weight of a good's category, divided by the "usable" prices in that category at the end of the price spell. Unlike the duration estimates, the weights for a given good can vary dramatically from month to month depending purely on the reporting of other goods. As such, I consider the unweighted passthrough case to be most reliable and do not discuss the weighted case below.

from aggregate data by Marazzi et al. (2005). The even columns report interaction terms for related parties, and the positive and generally statistically significant difference indicates higher intrafirm passthrough. For example, the unweighted case in the second column suggests that arm's length passthrough of differentiated goods is about 18.9 percent, compared with about 26.9 percent (26.9=18.9+8.0) for intrafirm goods. As expected, the estimates are smaller when restricted to dollar-priced goods, but the qualitative result remains, with arm's length passthrough of about 13.5 percent and intrafirm passthrough nearly 20 percent. Overall passthrough, and the intrafirm differential, in all these regressions may be somewhat smaller using this "conditional passthrough" specification as it does not allow for any lagged effect of exchange rate movements from preceding spells and there is measurement error in associating a given price change with a given exchange rate change.

5.2 Pooled Passthrough Estimates with Lags

To capture these lagged effects as well as the impact of shorter intrafirm duration on passthrough, pooled specifications are estimated with the form:

$$\Delta \ln p_j^{c,\, t} = \alpha + \sum_{n=0}^{N} \beta_1^n \Delta \ln e_j^{c,\, t-n} + \sum_{n=1}^{N} \beta_2^n \Delta \ln \pi_j^{c,\, t-n} + \sum_{n=1}^{N} \beta_3^n \Delta \ln \pi_j^{US,\, t-n} + \varepsilon_j^t, \quad (2)$$

where $\Delta \ln p_j^{c,\, t}$ is the monthly percentage change in the price (in dollars) of good j, imported from country c, from period $(t-1)$ to t. The term $\Delta \ln e_j^{c,\, t-n}$ is the monthly percentage change in the nominal exchange rate between the U.S. and country c over the period $(t-n-1)$ to $(t-n)$. The coefficients β_1^n for $n = 0..N$ are the concurrent and lagged exchange rate passthrough coefficients. For consistency, results are reported from the real exchange rate specification of (2), but the nominal specifications yields qualitatively similar estimates.

These regressions are run with two different methods for handling censoring in the BLS

prices. The first method uses all prices in the dataset, including imputed or other prices that would not be included in BLS price indices. The second method considers censored prices to be unchanged if they are surrounded by identical prices, the same assumption used earlier for the duration estimates. When censored months end with a new price, the price change is assumed to have occurred at that time, with zero price changes in the preceding censored months. These methods yield highly similar results, so only the results from the second method are presented below.

Following Gopinath et al. (forthcoming), Figure 3 shows a plot of the cumulative sum of exchange rate passthrough coefficients, $\sum_{n=0}^{N} \beta_1^n$, from regressions run separately for intrafirm and related party trades of dollar-priced differentiated goods. In order to eliminate any impact of non-dollar prices, which exhibit much higher passthrough, only dollar prices are shown in the plot. Price changes of magnitude larger than 50 percent are excluded. Intrafirm passthrough reaches 25 percent, compared with the arm's length rate of 15 percent, and after about 1 year, the 95 percent confidence intervals no longer overlap. Table 5 gives the point estimates and standard errors for passthrough after one-year and in the long-run from various estimates of specification (2). All four long-run estimates exhibit higher intrafirm passthrough, and in the cases that exclude outliers, the difference is statistically significant.

The magnitude of the difference between related party and arm's length passthrough fluctuates across these specifications as seen in Figure 3, Table 4, and Table 5, but even after ruling out compositional effects from different invoicing currencies, passthrough is often estimated to be about one-third larger. Combined with the results on duration, this implies that differentiated good prices will respond more to shocks in both the short- and long-run when traded intrafirm.

6 Other Potential Explanations

The three key patterns documented in the previous sections could be generated by the use of allocative prices which generate different profit maximizing quantities, by different accounting and tax concerns faced by integrated firms, or by some combination of the two. To directly test the hypothesis that these prices are allocative, one would need data on quantities, and the BLS data unfortunately do not include this information.[13] One can generate indirect evidence by ruling out common alternative hypotheses. Toward that end, this section considers qualitatively and quantitatively the possibility that these patterns are produced by inaccurate reporting to the BLS, price changes that reflect the annual tax calendar, or the incentive of multinationals to shift profits to countries with low tax rates.

First, one might reasonably worry that firms simply do not use transfer prices to allocate goods. Barro (1977) notes that long-term contracts might specify quantities along with a sticky price, though Gopinath and Rigobon (2008) show for the BLS data that only 7% of the importers report that their prices are specific to the quantity ordered. Multinational profits generally reflect numerous complicated joint internal and external sourcing decisions across potentially large numbers of products and geographies, precisely the context in which prices are most helpful for aggregating information and implementing efficient allocations.[14]

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[13]Ideally, one could test whether these prices are allocative by estimating related party and arm's length price elasticities for the same industry, using either an instrumental variables approach or identifying off cross-country "varieties" of the same good as in Feenstra (1994) and Broda and Weinstein (2006). Unfortunately, because the BLS data is highly sampled, there are rarely enough prices for any disaggregated category to match with total trade volumes. Nonetheless, I regressed import volumes for each trade type for each 3-digit SITC for 1997-2004 on price indices constructed as unweighted averages of price changes. Specifically, I estimated $\ln V_{z,t}^j = \beta_z^j \ln P_{z,t}^j + \varepsilon_{z,t}^j$, where V is the total trade volume in industry z in month t for trade type $j = AL, RP$ as recorded by the U.S. Census Bureau. If price innovations were exogenous and supply perfectly inelastic, the regression coefficient would capture the elasticity as $\epsilon_z^j = 1 - \beta_z^j$. Statistically significant estimates of β_z^j are recorded in 72 percent of the related party industries and 75 percent of the arm's length industries. There are 29 industries with enough data for estimates for both types. Among the 17 cases in which both estimates imply positive elasticities, there is a correlation above 50 percent, and a statistically significant regression coefficient between intrafirm and arm's length elasticities of 0.73. The results are suggestive, but given the weak power of this test, fall well short of proof that the prices are allocative.

[14]Firms are legally constrained in what they report to the IRS, but are unconstrained in the way they use internal prices. Managerial goals routinely require internal transfer prices that would themselves not comply

Instructional cases typically teach business students about allocative transfer pricing strategies rather than teaching them to directly set quantity transfers (e.g. Bailey and Collins, 2006). Even in the archetypal centrally planned production system, the Soviet Union, allocations were generally implemented using internal prices, some of which varied across plants for the same commodity in order to aggregate local information about production costs (see Bornstein, 1978, and the work cited therein).

It is certainly true, however, that firms can achieve any internal allocation without using prices. Management might simply instruct one plant to ship a given quantity of products to another. Relative to other datasets where every firm is required to provide pricing information, the IPP dataset is less likely to include such firms. For instance, customs authorities might legally require a price for each cross border transaction, even if internal prices are not used by the company for allocative purposes. The BLS, by contrast, has no authority to require any company to give it any particular price, and in cases where the price is not economically meaningful, the BLS would not have an incentive to do so. After all, its only mandate is to collect data for use in constructing price indices and conducting research.

Readers may suspect that unlike arm's length prices, intrafirm prices are set exactly once a year, in line with a company's accounting cycle. This would counterfactually increase related party price synchronization due to bunching in the timing of corporate tax payments. Further, 54 percent of related party goods had at least one price change in an 11 month period or less, compared to 49 percent of arm's length goods. 22 percent of related party goods had at least one price that remained unchanged 12 months later, compared to 24 percent of arm's length goods. Hence, one can rule out the scenario where prices change exactly every 12 months for essentially the same percentage of related party goods (76) and arm's length goods (73).

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with tax laws. Publicly listed firms, for instance, often maintain one set of books with internal prices, a second for tax purposes, and a third for generally accepted accounting principals (GAAP) reporting.

The BLS data is explicitly and legally isolated from the taxing authority, but some readers may still worry that the incentive to shift profits to low tax countries will lead companies to report prices different from those that drive actual production decisions. Such income shifting need not be implemented by altering good level internal prices. Accountants make year end "adjustments" to operating statements. Such broad-brush adjustments to revenues and cost of goods sold are likely to overwhelm the import of item level adjustments. Further, investigations of illegal transfer pricing practices generally do not consider item level tangible good prices as evidence, but rather, evaluate the return on invested capital or the overall operating structure as reported in annual financial statements.[15]

Whether profit shifting incentives lead to the reporting of non-allocative prices or if they actually change allocative transfer pricing behavior, we can test this by conditioning the empirical results on the tax differential between the exporting country and the U.S. After all, the incentive to use transfer prices to shift income grows with the magnitude of this tax gap, which is calculated as the average of the period-by-period gap (a very stable number over time) over the dataset. Two tax rate measures are used: the top marginal corporate rate as reported in the University of Michigan's World Tax Database (WTD) and Bureau of Economic Analysis (BEA) estimates of the effective tax rates paid by U.S. multinationals operating in foreign countries. Of the goods for which data is available for the tax gap and for good-by-good duration estimation, there are comparably sized groups with an average tax gap (in either direction) no greater than 5 percent in the WTD dataset and 10 percent in the BEA dataset.

Panels A and B of Table 6 report separate duration estimates for these two subsets of

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[15]For instance, the Internal Revenue Service (IRS, 2005) occasionally reports details on its "advanced pricing agreements" (APAs), binding contracts in which firms commit in advance to an IRS approved transfer pricing regime. In 2005, only about 15 percent of compliance tested APAs involved transfer pricing regimes that are evaluated at the product level, such as the "comparable uncontrolled price" (CUP) or "resale price method" of transfer pricing, in which reported prices must match equivalent market-based prices. About 75 percent were linked to aggregated items on year-end balance sheets or income statements, such as the "comparable profits method."

data. If income shifting is driving the lower intrafirm price duration for differentiated goods, one would expect this pattern to be particularly strong in the large tax gap group and weak or nonexistent in the small tax gap group. Though there are indeed some differences between the two groups, arm's length prices are stickier, and to similar degrees, in each subsample. Panel C shows results when this same exercise is applied to conditional passthrough regressions.[16] Two of the eight regressions do not exhibit significantly higher intrafirm passthrough, but the results taken together demonstrate that the result holds up equally well in countries with very similar and highly dissimilar tax rates compared to the U.S.

In sum, the BLS dataset is the most likely of available datasets to capture allocative transfer prices, the incentive for manipulating transfer prices at the good level (versus at the year-end financial statement level) is generally small, and the duration and passthrough results hold up equally for large and small tax differentials. Surely some reported transfer prices are non-allocative or are allocative but simply reflect tax concerns as discussed in, for example, Clausing (2003) and Bernard et al. (2006). My results suggest, however, that much of this behavior is orthogonal to these new empirical results. After all, if prices are purely non–allocative or solely set to shift income, they would likely change less often and have essentially no high-frequency relationship to the exchange-rate. The main empirical findings refute both of these predictions.

7   Cause and Effect of Intrafirm Pricing Differences

This paper has demonstrated empirical differences in intrafirm price dynamics. Further, Section 6 gave evidence that these differences are not driven by tax-shifting considerations. So, what is driving these differences?

One possibility is that transfer prices are generally allocative and the different maximiza-

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[16] This exercise cannot be applied to the synchronization results because separating any subset of price changes within 4-digit sectors will change all the measurements.

tion problems of arm's length and integrated firms leads to these empirical patterns. This type of environment is considered in Neiman (2009), which generates these empirical differences in a model where both types of firms face price adjustment costs. Intrafirm prices, unlike arm's length prices, are chosen to maximize the sum of the manufacturer's and distributor's profits. Transfer prices roughly follow marginal cost to avoid double marginalization, while arm's length prices are set at a markup which fluctuates depending on the relative price of the competitor. Alternatively, some results, such as the lower intrafirm price duration, may simply be driven by lower intrafirm price adjustment costs.

These explanations rely on transfer prices being allocative, and this paper cannot directly give support for this assumption because the BLS dataset lacks information on quantities. Nonetheless, if one assumes allocative pricing, how would the above results impact international macroeconomic dynamics?

To answer the question definitively, one would need to know if the price elasticities of demand for identical arm's length and intrafirm goods are the same. Neiman (2009) builds a model in which this is the case, but one might answer the question empirically by comparing price and quantity data for a given intermediate input both before and after an importer purchases its foreign supplier. Alternatively, these estimates could be done cross-sectionally as was attempted in footnote 13 of this paper. A better match of highly disaggregated price and quantity data, however, are required to do this with any power. Assuming the elasticities are similar, the shorter intrafirm price durations and higher long-run passthrough rates found in this paper would imply that a given percentage exchange rate shock would differentially impact trade flows based on the share of vertical integration. Or equivalently, if one asked what exchange rate change would be consistent with the elimination of a particular bilateral or multilateral external imbalance, as in Obstfeld and Rogoff (2005) or Dekle, Eaton, and Kortum (2008), the answer would differ depending on the share of trade that is conducted intrafirm.

The implications for welfare depend somewhat on the cause of these differences. One might start with the fact that price rigidities cause fluctuations in relative prices between firms that adjust and those that do not. Since these relative price movements are not driven (or at least not entirely driven) by changes in relative costs, they may reduce productivity. The differential degrees of stickiness implied by these results suggest that vertically integrated sectors or countries suffer less efficiency losses from nominal volatility (whether it be from exchange rates or inflation). Models that try to quantify this loss, in the spirit of Burstein and Hellwig (2008), would explicitly wish to capture this heterogeneity and should be calibrated to match data that includes intrafirm prices.

Finally, if marginal cost transfer pricing and variable markup arm's length pricing are driving the empirical results, it would imply that there are new dynamic costs and benefits of vertically integrating above and beyond the simple static gain from eliminating double marginalization. Put differently, the surplus created by vertically integrating would differ depending on the volatility of exchange rate or inflation shocks in the environment. A dynamic general equilibrium model with adjustment costs that endogenously generated vertical integration would be required to quantitatively assess these new benefits.

8   Conclusion

Given the remarkable scale of intrafirm transactions as a share of global trade, as well as the obvious heterogeneity in the vertical configuration of domestic industries, it is crucial to understand how such transactions differ from trades made at arm's length. I document significant new empirical differences between arm's length and intrafirm trades of differentiated goods in key pricing dynamics – intrafirm trades are less sticky, less synchronized, and exhibit higher exchange rate passthrough. The incentive to shift income to low tax countries is unlikely to be causing these results since they hold equally well for exporting countries

with comparable tax rates to the U.S. as for those with dissimilar rates. Hence, these pricing differences are suggestive that firm boundaries matter for allocations and their dynamics at the firm, sector, and country levels.

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Tables

Maximum Likelihood Estimates of Duration (months)				
Spell-Weighted		Non Spell-Weighted		
Median	Implied Frequency	Median	Implied Frequency	Obs (goods)
Panel A: Trade-Weighted				
After matching with Rauch Classifications:				
Overall	11.1	0.09	9.0	0.11	23,794
Arm's Length	11.6	0.09	9.3	0.11	14,911
Intrafirm	11.0	0.09	9.0	0.11	8,883
Differentiated Goods Only:					
Overall	13.5	0.07	11.0	0.09	18,260
Arm's Length	15.4	0.06	12.6	0.08	11,161
Related Party	12.0	0.08	9.7	0.10	7,099
Differentiated and Dollar-Priced Only:					
Arm's Length	15.8	0.06	13.0	0.08	10,324
Related Party	12.9	0.08	10.0	0.10	6,343
Panel B: Unweighted					
After matching with Rauch Classifications:					
Overall	12.7	0.08	10.2	0.10	23,794
Arm's Length	13.2	0.08	10.9	0.09	14,911
Intrafirm	12.0	0.08	9.7	0.10	8,883
Differentiated Goods Only:					
Overall	14.0	0.07	11.5	0.09	18,260
Arm's Length	15.5	0.06	12.4	0.08	11,161
Related Party	12.3	0.08	10.0	0.10	7,099
Differentiated and Dollar-Priced Only:					
Arm's Length	15.8	0.06	13.0	0.08	10,324
Related Party	12.3	0.08	10.0	0.10	6,343

Table 1: Maximum Likelihood Median Duration Estimates

Notes: Duration in months from good-by-good maximum likelihood estimation (methodology detailed in Section 3).

Table 2: Duration Patterns Conditional on Good Characteristics

	Duration (months)									
	Logs	Levels	Logs	Levels	Logs	Levels	Logs	Levels	Logs	Levels
Panel A: Trade-Weighted										
Arm's Length	Omitted	Omitted	Omitted	Omitted	Omitted	Omitted	Omitted	Omitted	Omitted	Omitted
Related Party	-0.2036 (0.0273)***	-3.7487 (0.4419)***	-0.1542 (0.0341)***	-2.9462 (0.5852)***	-0.1810 (0.0311)***	-3.0075 (0.5752)***	-0.1705 (0.0344)***	-2.7190 (0.6476)***	-0.1729 (0.0340)***	-2.7365 (0.6331)***
Std Dev of Exchange Rate Shock	0.0026 (0.0059)	0.1138 (0.0915)	0.0069 (0.0113)	0.1654 (0.1637)	0.0058 (0.0053)	0.1668 (0.0846342)**	0.0061 (0.0101)	0.1695 (0.1512)	0.0064 (0.0099)	0.1714 (0.1501)
Dollar Denominated	0.2328 (0.0515)***	3.8402 (0.5459)***	0.2253 (0.0484)***	3.6151 (0.5591)***	0.2517 (0.0484)***	4.1127 (0.4988)***	0.2728 (0.0466)***	4.1879 (0.5051)***	0.2706 (0.0473)***	4.1721 (0.5150)***
Elasticity Quintile									-0.0161 (0.0119)	-0.1162 (0.2115)
R-Squared	0.02	0.03	0.08	0.07	0.13	0.08	0.16	0.10	0.16	0.10
Observations	17787	17787	17787	17787	17787	17787	17787	17787	17787	17787
Panel B: Unweighted										
Arm's Length	Omitted	Omitted	Omitted	Omitted	Omitted	Omitted	Omitted	Omitted	Omitted	Omitted
Related Party	-0.1860 (0.0113)***	-3.0979 (0.1945)***	-0.1380 (0.0130)***	-2.3416 (0.2195)***	-0.1556 (0.0117)***	-2.4508 (0.2039)***	-0.1303 (0.0130)***	-2.0128 (0.2240)***	-0.1313 (0.0130)***	-2.0285 (0.2239)***
Std Dev of Exchange Rate Shock	0.0086 (0.0021)***	0.1976 (0.0402)***	0.0202 (0.0032)***	0.4044 (0.0564)***	0.0111 (0.0020)***	0.2370 (0.0387)***	0.0199 (0.0032)***	0.4097 (0.0565)***	0.0196 (0.0032)***	0.4060 (0.0566)***
Dollar Denominated	0.1315 (0.0170)***	3.2376 (0.2600)***	0.1341 (0.0195)***	3.2669 (0.3095)***	0.1530 (0.0175)***	3.5313 (0.2735)***	0.1580 (0.0197)***	3.5668 (0.3162)***	0.1601 (0.0197)***	3.5996 (0.3166)***
Elasticity Quintile									-0.0253 (0.0049)***	-0.4011 (0.0858)***
R-Squared	0.02	0.02	0.05	0.04	0.09	0.06	0.11	0.07	0.11	0.08
Observations	17787	17787	17787	17787	17787	17787	17787	17787	17787	17787
Country Dummies	N	N	Y	Y	N	N	Y	Y	Y	Y
2-Digit Industry Dummies	N	N	N	N	Y	Y	Y	Y	Y	Y

* Significant at 10%, ** Significant at 5%, *** Significant at 1%

Notes: Standard errors are reported in parentheses and are robust to heteroskedasticity. Duration in months from good-by-good spell-weighted maximum likelihood estimation (methodology detailed in Section 3).

Marginal Effect of Percentage of Other Price Changes of Same Direction and Type in Sector				
	Arm's Length		Related Party	
	Coefficient	Standard Error	Coefficient	Standard Error
Price Increases	0.3306	(0.0139)***	0.2307	(0.0134)***
Price Decreases	0.5928	(0.0158)***	0.4912	(0.0139)***

* Significant at 10%, ** Significant at 5%, *** Significant at 1%

Table 3: Price Synchronization and Vertical Structure of Trade

Notes: Standard errors are reported in parentheses and are robust to heteroskedasticity. Ordered probit estimated with 2-digit SITC and country dummies. Unreported covariates include cumulative changes in nominal exchange rate since last price change and other price changes in the sector in the opposite direction. Sector definition is 4-digit.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
Absolute Size of Changes	All	All	∈ (0.01,0.50)	∈ (0.01,0.50)	∈ (0.01,0.25)	∈ (0.01,0.25)	All	All	∈ (0.01,0.50)	∈ (0.01,0.50)	∈ (0.01,0.25)	∈ (0.01,0.25)
Panel A: Unweighted (Baseline)												
Exchange Rate	0.2234 (0.0110)***	0.1894 (0.0108)***	0.2039 (0.0073)***	0.1822 (0.0098)***	0.1709 (0.0057)***	0.1476 (0.0075)***	0.1605 (0.0110)***	0.1347 (0.0104)***	0.1401 (0.0068)***	0.1252 (0.0091)***	0.1284 (0.0054)***	0.1115 (0.0072)***
Intrafirm * Exchange Rate		0.0795 (0.0238)***		0.0520 (0.0142)***		0.0553 (0.0110)***		0.0612 (0.0243)**		0.0363 (0.0134)***		0.0407 (0.0106)***
Panel B: Trade Weighted												
Exchange Rate	0.1870 (0.0229)***	0.1684 (0.0174)***	0.1717 (0.0135)***	0.1586 (0.0152)***	0.1374 (0.0097)***	0.1307 (0.0113)***	0.1476 (0.0240)***	0.1286 (0.0176)***	0.1286 (0.0138)***	0.1166 (0.0153)***	0.1093 (0.0099)***	0.1044 (0.0115)***
Intrafirm * Exchange Rate		0.0407 (0.0510)		0.0294 (0.0288)		0.0149 (0.0210)		0.0417 (0.0535)		0.0271 (0.0296)		0.0110 (0.0213)
Includes Non-Dollar Prices?	Y	Y	Y	Y	Y	Y	N	N	N	N	N	N
Observations	45,385	45,385	37,055	37,055	35,227	35,227	43,045	43,045	34,972	34,972	33,292	33,292

* Significant at 10%, ** Significant at 5%, *** Significant at 1%

Table 4: Conditional Exchange Rate Passthrough

Notes: Estimates from passthrough specification (1) on differentiated good prices. All changes are included, only those of magnitude between 1 and 50 percent, or only those between 1 and 25 percent, depending on the column. Price changes are assumed to have occurred at the end of any censored period. Standard errors are reported in parentheses and are robust to heteroskedasticity. All regressions include 2-digit industry and country dummies.

	Long Run		After 1 Year		Long Run		After 1 Year	
	Estimate	Standard Error	Estimate	Standard Error	Estimate	Standard Error	Estimate	Standard Error
Dropping Large Changes								
Intrafirm	0.30	0.01	0.19	0.01	0.25	0.01	0.13	0.01
Arm's Length	0.22	0.01	0.16	0.01	0.16	0.01	0.10	0.01
Including All Changes								
Intrafirm	0.30	0.04	0.23	0.02	0.21	0.04	0.17	0.02
Arm's Length	0.21	0.02	0.16	0.01	0.16	0.02	0.11	0.01
Includes Non-Dollar Prices?	Y		Y		N		N	

Table 5: Long-Run Exchange Rate Passthrough

Notes: Estimates from passthrough specification (2) on differentiated good prices. Standard errors are reported in parentheses and are robust to heteroskedasticity. "Large Changes" include any of an absolute magnitude greater than 50 percent. "After 1 Year" gives the sum of the contemporaneous and first 11 lag coefficients, while "Long Run" gives the estimate when adding up all 3 years of coefficients.

Robustness of Duration and Passthrough Results to Different Tax Rates

| | World Tax Database (University of Michigan) | | | | Bureau of Economic Analysis | | | |
| | <=5 % Tax Differential | | >5 % Tax Differential | | <=10% Tax Differential | | >10% Tax Differential | |
	Median	Implied Frequency	Median	Implied Frequency	Median	Implied Frequency	Median	Implied Frequency
Panel A: Trade-Weighted Maximum Likelihood Estimates of Duration								
Differentiated Goods Only:								
Overall	14.5	0.07	16.1	0.06	15.8	0.06	16.7	0.06
Arm's Length	16.0	0.06	17.0	0.06	19.0	0.05	18.0	0.06
Related Party	12.9	0.08	14.5	0.07	13.0	0.08	13.9	0.07
Differentiated and Dollar-Priced Only:								
Arm's Length	16.3	0.06	17.2	0.06	18.0	0.06	18.2	0.05
Related Party	13.0	0.08	15.8	0.06	13.1	0.08	13.6	0.07
Panel B: Unweighted Maximum Likelihood Estimates of Duration								
Differentiated Goods Only:								
Overall	15.6	0.06	16.9	0.06	16.8	0.06	17.2	0.06
Arm's Length	16.6	0.06	17.7	0.06	18.0	0.06	18.5	0.05
Related Party	13.6	0.07	14.5	0.07	15.0	0.07	14.6	0.07
Differentiated and Dollar-Priced Only:								
Arm's Length	17.0	0.06	17.9	0.06	18.7	0.05	18.6	0.05
Related Party	13.8	0.07	15.0	0.07	15.5	0.06	14.9	0.07

	Coefficient	Coefficient	Coefficient	Coefficient	Coefficient	Coefficient	Coefficient	Coefficient
Panel C: Unweighted Conditional Passthrough (Absolute Value of Changes between 0.01 and 0.50)								
Exchange Rate	0.1786 (0.0115)***	0.1105 (0.0106)***	0.1552 (0.0189)***	0.1322 (0.0187)***	0.1805 (0.0166)***	0.1202 (0.0161)***	0.1750 (0.0175)***	0.0963 (0.0133)***
Intrafirm * Exchange Rate	0.0496 (0.0167)***	0.0428 (0.0158)***	0.0439 (0.0277)***	0.0063 (0.0264)	0.0788 (0.0248)***	0.0265 (0.0236)	0.0520 (0.0250)**	0.0743 (0.0197)***
Includes Dollar Prices?	Y	N	Y	N	Y	N	Y	N

Table 6: Duration and Passthrough by Vertical Structure and Tax Differential

Notes: Duration in months from good-by-good spell-weighted maximum likelihood estimation (methodology detailed in Section 3) and conditional passthrough estimates. The data are split using two measures of tax rates into a group with rates similar to the United States and a group with more dissimilar rates.
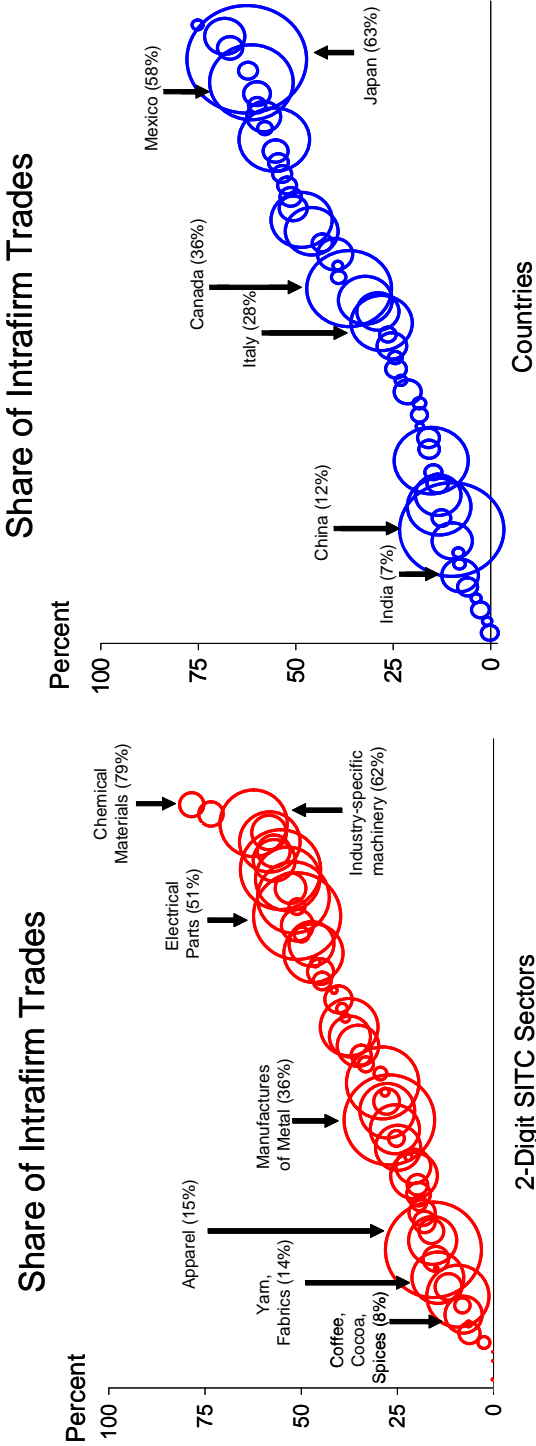
Figures

Figure 1: Intrafirm Trade Shares by Sector and Country

Notes: Intrafirm trades defined at the good level and pooled over the dataset. Circle diameters are in proportion to the share of imports to the U.S. Nearly all sectors and large countries have a non-zero share of intrafirm trading. The United States, not included in the above plot, exports about one third of its goods to related parties.
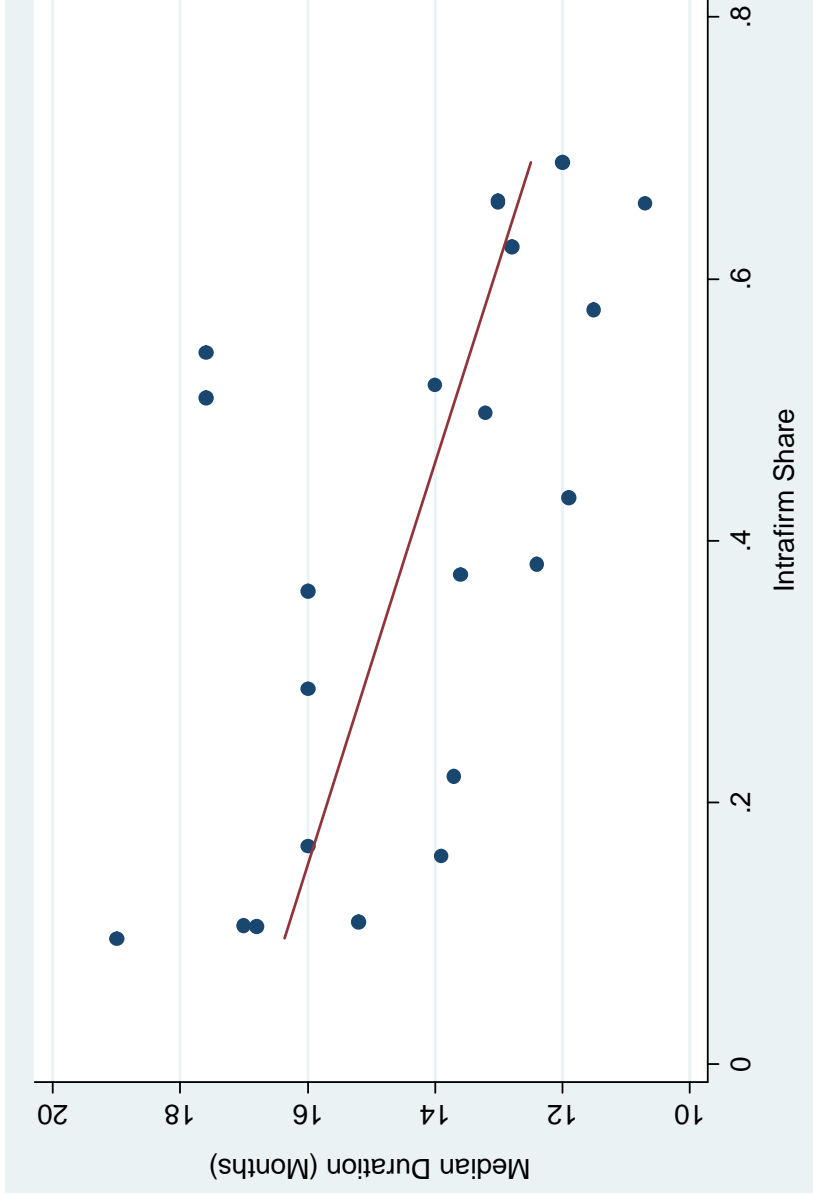
Figure 2: Duration and Vertical Structure by Exporting Country

Notes: Median duration in months from good-by-good spell-weighted maximum likelihood estimation (methodology detailed in Section 3) for the largest 21 exporters to the U.S. of differentiated goods are plotted against the share of these trades that are intrafirm. For instance, one point represents the median duration of all differentiated goods imported from China and the percentage of Chinese exports of these goods that were intrafirm over the life of the dataset.
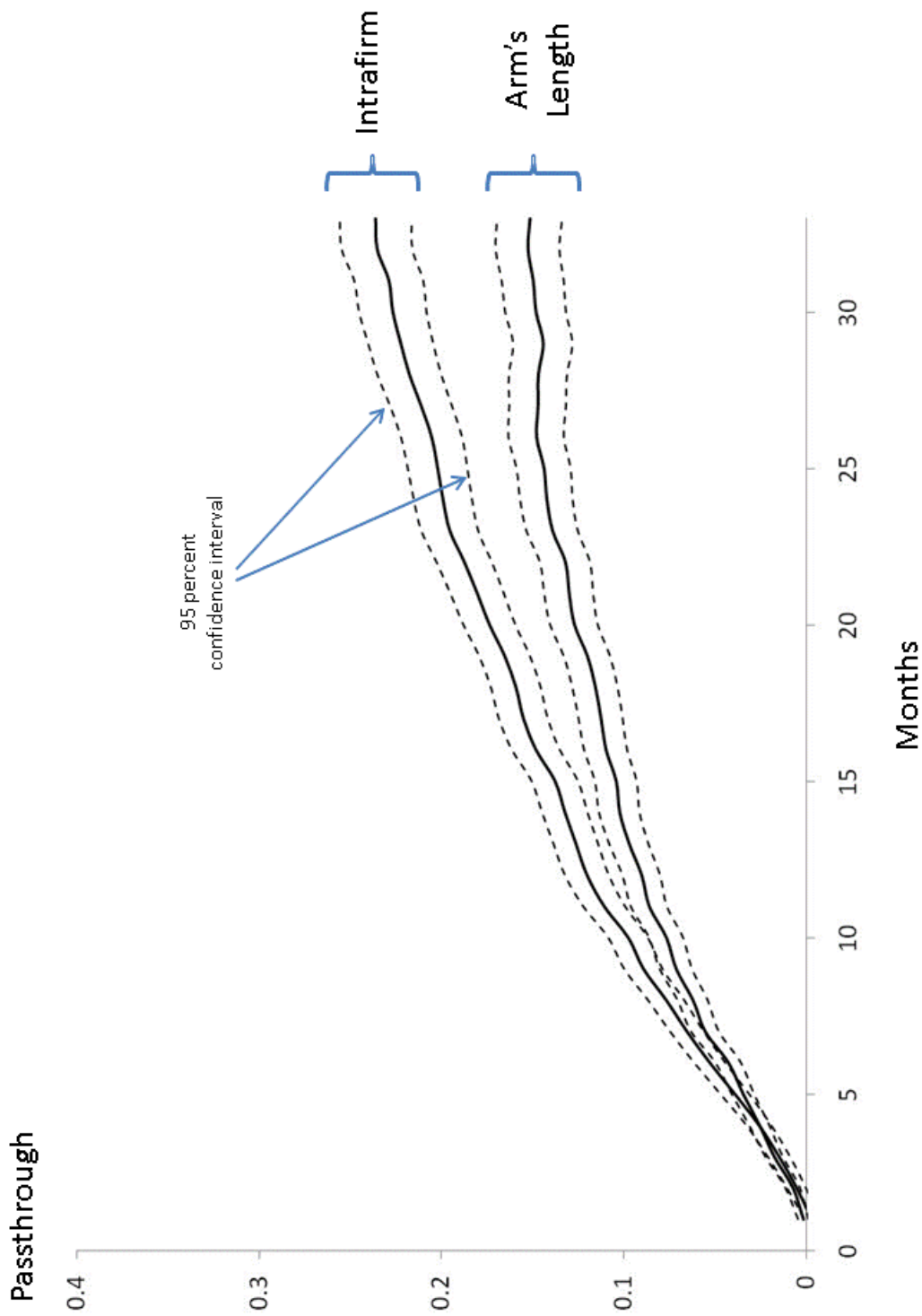
Figure 3: Passthrough and Vertical Structure (Dollar Prices Only)

Notes: Passthrough estimates from specification (2), where all monthly price changes are pooled and regressed against current and lagged monthly exchange rate changes. Moving from left to right along the x-axis, the plot shows the cumulative sum of these coefficients. The plot only includes dollar-priced differentiated goods, omits changes that exceed 0.5 in magnitude, and gives results from separate estimates of intrafirm and related party prices.

Appendix A: Details on the BLS Data

This appendix discusses the characteristics of the BLS data that are most relevant to the analysis. One key limitation in the data is that many months exist in which survey responses were not received and hence a price is either missing or is imputed. This paper only uses prices denoted "net usable" in the BLS database. This means that aside from rare exceptions, it only includes prices that reflect an actual survey response.

The BLS data is highly disaggregated and designed to track the price of an identical good over time, for example, a "Rug; 100% New Zealand wool; hand-tufted; hand-hooked; style name: XXX" or "Cello #XXX, XXX Brand, maple ribs and back, spruce top, ebony fittings, 4/4 size, not hand made" (where XXXs replace proper nouns used by the importer to identify the particular good). I use the term "good" in this paper to refer to what the BLS calls an "item_code". In collecting the data, analysts make every attempt to ensure the comparability over time of a good's quality, unit (i.e. "6 to a box"), and any other negotiated term of value related to the shipment. In addition to the price of the good, the BLS records the month the shipment is received in the U.S., the exporting/importing country, the currency of denomination, and a classification of the good by an internal code based on the 10-digit harmonized system.

The essential feature of the data for this paper's purposes is its delineation of intrafirm trades from arm's length trades. There is no technical definition used to classify trades as between related parties or otherwise. Rather, the BLS analysts ask the survey respondents (often shipping managers) whether there is an ownership relationship with the buyer. Though the cleanest example of intrafirm trade occurs between a multinational parent and its wholly-owned subsidiary (or vice-versa), not all intrafirm trades reflect this structure. Some analysts speculate that the 10 percent joint ownership threshold used by a different agency, the Bureau of Economic Analysis, in collecting other data is often implicitly used by respondents in characterizing the trade as intrafirm. Given the surveys are typically filled out by employees that deal with purchasing and shipping, and not necessarily by those familiar with corporate financial structure, it is unlikely that small holdings with no bearing

on prices would be reflected in this classification.

Within the grouping of intrafirm prices, the BLS also classifies trades as reflecting one of four categories: "cost-based pricing", "market-based pricing", "other non-market based pricing", and "unknown methods." The criteria used to classify a price into these sub-categories, however, have not been consistently applied over time, and the BLS staff feel far more confident in the distinction of arm's length versus intrafirm than the distinctions within the intrafirm category. As such, this paper focuses on the former classification and ignores the sub-categories.

Many of the price changes are very small, with about 11 and 16 percent of arm's length and intrafirm differentiated price changes less than 1 percent in absolute size. The majority of these changes appear to be correctly identified in the surveys, though some clearly reflect recording or rounding errors. It is not possible to disentangle these manually, but as a robustness check, all price changes smaller than 1 percent were excluded. The baseline duration, synchronization, and passthrough results did not meaningfully change.

Sampling procedures unfortunately do not yet take into account whether trades take place intrafirm or not. For instance, if 50 percent of an industry's trade happens to be intrafirm trade, the BLS will not necessarily set a sampling goal to have an equal mix of intrafirm and arm's length survey respondents. The correlations of the related party shares from each country and in each 6-digit NAICS sector in the BLS data with the shares in data collected by the Foreign Trade Division of the U.S. Census Bureau are 86 and 83 percent, respectively. These correlations are calculated using the 55 largest countries and 193 largest 6-digit NAICS sectors in the BLS data, or those countries and sectors with observations on at least 50 goods. Given the Census data is quantity data with a far broader coverage, these high correlations (even at highly disaggregated levels) indicate the suitability of using the share of sampled intrafirm prices as a good proxy for the share of intrafirm trade that actually occurs across various categories.