LOST IN TRANSIT:
PRODUCT REPLACEMENT BIAS AND PRICING TO MARKET

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ABSTRACT

Product replacement is frequent in the micro-data that underlie U.S. import and export price indices, while price changes are infrequent. Consequently, over 40% of price series in the data have no price changes and roughly 70% have two price changes or less. In constructing price indices, price adjustments that occur at the time of product replacements tend to be dropped. If price adjustments disproportionately occur at the time of product replacements then price adjustments are disproportionately unobserved. We show that this \product replacement bias" may distort the measured long-run relationship between import and export prices and the exchange rate by a factor of between 1.7 and 2.2. Accounting for this bias, we find that the price of non-oil U.S. imports (relative to domestic consumption) responds by 0.6-0.7% for each 1% change in the U.S. real exchange rate, while the price of U.S. exports (relative to foreign consumption) responds by roughly 0.8%. This contrasts with conventional pass-through estimates of 0.2-0.4% for non-oil import prices and 0.9% for export prices. Thus, we find that the degree of pricing to market for U.S. imports and exports is more symmetric and the degree of pricing to market for U.S. imports more moderate than conventional measures suggest. Adjusting for product replacement bias also substantially raises the volatility of the terms of trade. These results improve the fit of the data to standard models.

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1 Introduction

The post-Bretton Woods period of flexible exchange rates has been characterized by large swings in the U.S. real exchange rate. Simple models suggest that these large swings in relative prices of U.S. products should lead to large amounts of expenditure switching between U.S. and foreign products. In practice, expenditure switching has been modest. This "disconnect" of exchange rates from the rest of the economy is one of the major puzzles of open economy macroeconomics.

A leading potential explanation for this puzzle has been that exporters "price to market". Exporters are said to "price to market" if they adjust their markups to stabilize the local currency prices of their products (Krugman, 1987). Pricing to market thus implies that import and export prices change less than 1% for every 1% change in the exchange rate. Krugman (1993) and Engel (2002), among others, emphasize that the extent of pricing to market has profound implications for the degree of expenditure switching associated with exchange rate changes. If firms stabilize their prices in local currency terms, expenditure switching associated with exchange rate changes can play little role in resolving external imbalances.

Conventional estimates suggest that the prices of non-oil imports into the U.S. relative to all domestic prices change by only 0.2-0.4% for each 1% change in the U.S. real exchange rate.[1] On the other hand, conventional measures suggest that the prices of U.S. exports relative to all foreign prices change by roughly 0.9% for each 1% change in the U.S. real exchange rate. In other words, these estimates suggest that pricing to market by foreign importers into the U.S. market is substantial, while pricing to market by U.S. exporters in foreign markets is much less substantial.

Simple models such as Backus, Kehoe and Kydland (1992), Stockman and Tesar (1995) and Obstfeld and Rogoff (1995) do not generate pricing to market. The large amount of measured pricing to market in U.S. imports has motivated a great deal of work on models designed to generate this kind of behavior by firms. One strand of the literature has focused on price rigidity as a source of pricing to market (Betts and Devereux, 1996, 2000; Chari, Kehoe and McGrattan, 2002). However, the apparent magnitude and persistence of pricing to market has shifted the focus of the literature towards mechanisms that can generate long run deviations in relative prices (Dornbusch,

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[1]Campa and Goldberg (2005) report an estimate of roughly 0.4 for long run exchange rate pass-through into U.S. import prices for the period 1975 to 2003. Marazzi and Sheets (2007) report that long-run pass-through excluding oil was around 0.5 for the period from the early 1970's to the late 1990's, but has fallen to around 0.2 over the period from 1997 to 2004. Similar estimates are obtained by Gopinath, Itskhoki, and Rigobon (2008). These papers run pass-through regressions in nominal terms. Knetter (1989) and Marston (1990) run regressions in real terms that are closer to the specifications we adopt in this paper. Burstein, Echenbaum, and Rebelo (2005) report much higher pass-through for developing countries during large devaluations.

1987; Goldberg and Verboven, 2001; Corsetti and Dedola, 2005; Atkeson and Burstein, 2008; Gust, Leduc and Vigfusson, 2006; Gopinath, Itskhoki and Rigobon, 2007).[2] Matching the measured amount of long run pricing to market for U.S. imports along the lines suggested in these papers requires a great deal of strategic complementarity in price setting. This leads to important changes in the dynamic implications of the model, potentially generating long delays in the responsiveness of prices to exchange rate fluctuations. Standard models of pricing to market also do not explain why pricing to market appears to be so much greater for U.S. imports than for U.S. exports.

We argue that conventional measures of pricing to market are seriously biased due to measurement issues associated with the frequent replacement of products in the modern economy. The U.S. import and export price indices are based on micro price-data collected from U.S. firms. In these data, roughly 5% of products are replaced each month, while the frequency of price change of the median product is only about 8% per month (Gopinath and Rigobon, 2008). This combination of rapid product turnover and infrequent price changes implies that roughly 45% of price series in these data have no price changes and more than 70% have 2 price changes or less. Gopinath, Itskhoki, and Rigobon (2008) discuss the idea that these features of the data might affect measured exchange rate pass-through.

A natural question is: What happens to prices when products are replaced? This is a difficult problem since the quality, size and specifications of the new product might not be the same as those of the old product. The ideal solution to this problem is to use hedonic methods to estimate the quality change associated with product replacements (Court, 1939; Griliches, 1961; Pakes, 2003). For most products, however, it is extremely costly and difficult to accurately measure quality change (Abraham et al., 1998). Hedonic adjustments are, therefore, used in only a tiny fraction of cases.[3] In practice, a large fraction of product replacements are "linked-into" the index; meaning that the price comparison between the first observation of the new product and the last observation of the old product is dropped when changes in the index are calculated.[4] Indices constructed by linking in new products are referred to as "matched model indices" since all price comparisons on which such indices are based are for identical items.

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[2]Other notable models of pricing to market include Alessandria (2004), Ravn, Schmitt-Grohe and Uribe (2007), Drozd and Nosal (2008), Hellerstein (2008) and Nakamura and Zerom (2009).

[3]Feenstra (1994) and Broda and Weinstein (2006) develop and apply an alternative to hedonic methods. These papers make strong enough functional form assumptions about demand that quality adjusted prices can be backed out from prices and quantities.

[4]In some cases, the IPP deems a change in a product to be sufficiently small that the concurrent change in price is used in the index with no adjustment for a change in quality. In these cases, the IPP does not record a product substitution.

The literature on the "new goods problem" in index number theory points out that if new products are systematically introduced at lower quality adjusted prices than the old products they replace, then a matched model price index will yield an upward biased estimate of inflation and a matched model quantity index will yield a downward biased estimate of output growth.[5] Our focus is *not* on this bias in the average level of measured inflation caused by product replacement. Rather, we argue that the combination of price rigidity and product replacement implies that matched model price indices yield biased estimates of the comovement between prices and other variables such as exchange rates. In other words, the new goods problem may cause average measured inflation to be 2.5% when true average inflation is 2%, while the bias we focus on may cause the measured long run response of prices to a 1% change in the real exchange rate to be 0.35% when the true response is 0.7%.

The implicit assumption embodied in the practice of linking new products into the price index is that the frequency and size of price changes at the time of product replacements are the same on average as the frequency and size of price changes for continuing products. If instead prices are disproportionately likely to change at the time when products enter the data set, the procedure of linking-in new products will disproportionately drop price change observations and overweight observations with no price change. This will lead the resulting measured price index to exhibit less responsiveness to aggregate variables such as the exchange rate than the true aggregate price index. We refer to this bias as the "product replacement bias."

It is helpful to consider an extreme example in which the price of each product remains fixed for the entire life of the product and all price adjustment occurs at the time of product replacements. Figure 1 depicts this type of setting. In the figure, the exchange rate is depreciating. Assume for simplicity that the product replacements involve no quality change. Agents living in this economy can observe the quality of each product. It is therefore obvious to them that prices are rising as the exchange rate rises. A price index based only on price comparisons for identical items will however remain constant throughout in this example since prices only change at the time of product replacements. Estimates of the comovement of prices and exchange rates using this price index will yield no comovement irrespective of what the true degree of comovement is.

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[5]The Boskin commission argued that the new goods bias led to an upward bias in the consumer price index of 0.6% per year (Boskin et al., 1996). Similarly, Bils and Klenow (2001), Hausman (2003), Nordhaus (1998), Pakes (2003) and Bils (2008), have emphasized the importance of the new goods bias in distorting measured inflation and economic growth. Goldberg et al. (2008) show that new imported varieties contributed substantially to effective price declines for Indian firms after a trade liberalization. In contrast, Moulton and Moses (1997) , Abraham et al. (1998), Triplett (1997) and Hobijn (2002) argue that the Boskin Commission overestimated the new goods bias.

Product replacement bias is a selection bias problem. It arises if prices are disproportionately flexible at the time of product replacements, since price changes at the time of product replacements are dropped in a matched model index. There is reason to believe that prices are in fact disproportionately flexible at the time they are introduced into the dataset. First, firms must set new prices for new products, while continuing products have a low probability of price change. Second, firms often negotiate new prices when they sign new contracts with their customers (Carlton, 1986). New products are often initiated into the dataset at these points, since this is when firms switch suppliers. Third, the sampling procedure used by the BLS implies that measured prices for products that have just been initiated into the import and export price datasets are likely to be "fresher" than prices for continuing products, regardless of the nature of the "true" underlying price rigidity. In section 5, we provide direct empirical evidence that prices are equally "fresh" when they are initiated into the import and export price datasets as they are following subsequent price changes.

The example depicted in figure 1 shows that product replacement bias causes a *downward* bias in the measured comovement between prices and the exchange rate when prices are measured in the currency in which they are rigid. Most U.S. imports display price rigidity in U.S. dollars—they are local currency priced (LCP)—leading to a downward bias in conventional measures of import price pass-through. However, product replacement bias causes an *upward* bias in the measured comovement between prices and the exchange rate when prices are rigid in another currency than the one they are being measured in. Consider a U.S. export into the Euro area that is priced in U.S. dollars—i.e. is "producer currency priced" (PCP)—and the price of which is only adjusted at the time of product replacements. A matched model Euro area price index based on a collection of such product lines would display one-for-one comovement with the exchange rate regardless of the true relationship between prices and exchange rates.

We use BLS micro data on import and export prices over the period 1994-2007 to estimate the quantitative importance of product replacement bias. We estimate product replacement bias separately for LCP imports, PCP imports, LCP exports and PCP exports. In each of these four cases, we find that conventional estimates of comovement are biased by a factor of between 1.7 and 2.2 toward zero pass-through in the currency in which prices are rigid. Adjusting for this bias, we find that the price of non-oil U.S. imports (relative to domestic consumption) responds by 0.6-0.7% for each 1% change in the U.S. real exchange rate, while the price of U.S. exports (relative to foreign consumption) responds by roughly 0.8%. This contrasts with conventional pass-through estimates

of 0.2-0.4% for non-oil import prices and 0.9% for export prices.[6]

Taken together, these estimates bring the data better in line with standard models. First, adjusting for product replacement bias yields estimates of pricing to market that are much more symmetric between U.S. imports and exports. Second, our adjusted estimates imply a more modest amount of pricing to market for U.S. imports. The large amount of pricing to market implied by conventional measures is difficult to match theoretically. It necessitates large markups and a degree of strategic complementarity in price setting that strains existing general equilibrium models along other observable dimensions. For instance, the preferred calibration of Corsetti and Dedola (2005) implies a roughly 0.9% change in import and export prices for each 1% change in exchange rates in the long run, while the corresponding number for the preferred calibration of Atkeson and Burstein (2008) is roughly 0.75%.

Product replacement bias helps to explain differences in measured pass-through across countries. Burstein, Echenbaum, and Rebelo (2005) report much higher pass-through for developing countries during large devaluations than for developed countries. Many developing countries and smaller developed countries have a large fraction of imports and exports priced in a foreign currency. The high fraction of PCP (foreign currency priced) imports implies that measures of import price pass-through are likely to be upward biased as a consequence of product replacement bias. In contrast, product replacement bias is likely to generate a downward bias in measures of import price pass-through for large developed countries in which most imports are priced in domestic currency.

Finally, our results imply that conventional measures of import and export prices understate the volatility of the U.S. terms of trade. Measured in U.S. dollar terms, conventional indices suggest that non-oil import prices rise by 0.2-0.4% and export prices rise by 0.1% when the U.S. exchange rate depreciates by 1%. Taken together this implies a 0.1-0.3% deterioration of the non-oil terms of trade. Adjusting for product replacement bias, we find that non-oil import prices rise by 0.6-0.7% and export prices rise by 0.2% when the U.S. exchange rate depreciates by 1%. These estimates imply a 0.4-0.5% deterioration in the non-oil terms of trade. If movements in the terms of trade are dominated by movements in the exchange rate, our results suggest that conventional measures of the volatility of the terms of trade may be downward biased by a factor of between 1.3 and 5. As Drozd and Nosal (2008) emphasize, standard models without pricing to market imply that the

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[6]One might ask why we do not calculate an adjusted index rather than calculating the bias in the estimate for long run pass-through. The reason is that under the minimal assumptions we make, we are unable to calculate an adjusted index in a period by period sense. An adjusted index would require more structural assumptions about the relative timing of price movements and exchange rate changes.

terms of trade should be more volatile than the real exchange rate. Conventional measures suggest that the terms of trade is much less volatile than the real exchange rate. Our results suggest that these conventional measures substantially overstate this discrepancy between standard models and the data.

Our work builds heavily on recent papers that use the BLS microdata on import and export prices to study price rigidity and pass-through. Gopinath and Rigobon (2008) document basic facts about the frequency of price changes and substitutions in these data. Gopinath, Itskhoki, and Rigobon (2008) argue that there is a substantial difference in conventionally measured long-run exchange rate pass-through for LCP versus PCP imports. They also discuss the idea that substitutions could cause problems for price index measurement. Gopinath and Itskhoki (2008) argue that the degree of exchange rate pass-through is related to products' frequency of price change. Neiman (2008) documents a number of related facts regarding price rigidity and exchange rate pass-through for intrafirm prices.[7] Berger et al. (2009) link IPP micro-data with CPI micro-data to estimate the size of U.S. distribution wedges. Fitzgerald and Haller (2008) study pricing-to-market using analogous data from Ireland.

The closest antecedent to the ideas presented in this paper is the literature on disproportionately large price changes at the time when new products are linked into price indices (Armknecht and Weyback, 1989; Liegey, 1993; Reinsdorf, Liegey and Stewart, 1996; Triplett, 1997). Our work is also related to the measurement problems discussed in Houseman (2007). Our work is motivated in part by Erickson and Pakes (2008), who develop an experimental hedonic price index for televisions that accounts, among other things, for price rigidity. Finally, our findings regarding the importance of product replacement in generating low measured exchange rate pass-through are consistent with ongoing research by Jon Faust and John Rogers that argues that there is a negative relationship across industries between the frequency of product substitutions and measures of exchange rate pass-through (Rogers, 2006).

The paper proceeds as follows. Section 2 describes the BLS micro data underlying the U.S. import and export price indices that we use in our empirical analysis. Section 3 presents measures of pricing to market for U.S. imports and exports for the period 1982-2007 based on conventional methods. Section 4 derives expressions for product replacement bias as a function of the frequency of price change and the frequency of product replacement. Section 5 presents estimates of the frequency of price change and the frequency of product replacement and our quantitative estimates

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[7]See also, Gopinath, Itskhoki, and Rigobon (2008) and Clausing (2001).

of product replacement bias. Section 6 discusses alternative measures of pricing to market and presents an example that allows for errors in the timing of price changes due to data gathering problems. Section 7 discusses several additional implications of product replacement bias. Section 8 concludes.

2   Data Description

We use three sets of data. First, we use the microdata underlying the U.S. import and export price indices. These data are collected by the International Prices Program (IPP) of the Bureau of Labor Statistics (BLS). Second, we use aggregate U.S. import and export price indices produced by the Bureau of Economic Analysis (BEA) as a part of the National Income and Product Accounts (NIPA). Third, we use exchange rate data from the Federal Reserve Board and the International Monetary Fund (IMF). We describe these data in turn.

The U.S. import and export price indices were introduced in the early 1980's to provide a more accurate alternative to unit value indices. The micro data we use cover the time period 1995-2007. We exclude intrafirm prices in our baseline analysis. For our sample of countries, excluding intrafirm prices, the total number of product-months for which IPP attempts to record a price is roughly 1.5 million or about 100,000 per year.

The IPP data are collected using voluntary surveys. To initiate a product into the dataset, IPP collects a detailed item description and a particular set of transaction terms. Transaction terms include the number or type of units priced, the country of destination or origin, the port of exit or entry, the discount structure, and in some cases the duty applied to the product. The price provided by the reporter during the initialization must be a transaction price.[8]

After a product is initiated, price information is collected using a repricing form that includes pre-filled information collected during the initialization process. One concern about this procedure is that it may result in differential reporting friction for reporting a "new" price for continuing products as opposed to newly initiated products. The repricing form first asks whether the price has changed relative to the previous month and then asks the respondent to report a new price if the price did change. The easiest response for the reporters is simply to report that no change has

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[8]At initiation, reporters may also sometimes provide future prices if these are considered to be known, e.g., because of an existing contract. In a small number of cases—e.g., crude petroleum, grain, traded services, and various automobile and automobile part prices—IPP collects price data from secondary sources. In general, these data are not included in the database we study.

occurred.[9]

Price imputation is a major phenomenon in the import and export price data. According to our estimates, of the total number of product-months in the database, reported prices are not available in about 40% of cases. For continuing products, the reporter is allowed to report an estimated or "list" price if there was no transaction or a transaction price is not available. An estimated price is the reporter's estimate of "the price that would have been charged during the reporting period." Also, since many goods are imported or exported in large shipments rather infrequently, many price quotes are collected "off-cycle" at a less than monthly frequency. This is particularly the case for seasonal items. The frequency with which data are collected may also be reduced on an ad hoc basis to limit the firm's reporting burden. In the intervening months, the IPP computer system either generates an imputed price or pulls forward the last month's price. The price imputations are calculated by taking the average price increase within a product category and applying this change to the previous month's price. When the price is again observed, the price reverts back to the actual observed price. This implies that long-run inflation for prices imputed in this way equals long-run inflation for a price series in which prices are simply pulled forward through missing spells (Feenstra and Diewert, 2000).

In our analysis, we start by dropping all imputed and estimated prices. We then use a simple approximation to the IPP procedure for imputing missing prices. We fill forward the last observed reported price through all missing periods. To avoid introducing spurious price changes associated with numerical issues in converting prices quoted in foreign currencies into dollars, we use the "reported price" rather than the net price in our baseline analysis.

The IPP accepts reported prices in any currency. In practice, about 93% of import prices and 98% of export prices are reported in U.S. dollars. In situations where the item is covered by a written contract, the IPP also records whether the contract is contingent on the exchange rate, and incorporates this information when imputing prices. To maximize the likelihood of repricability over time, IPP accepts prices according to any pricing basis (FOB, CIF, etc.).

In principle it may be desirable to define a product sufficiently narrowly that products with any observable differences in characteristics are viewed as entirely different products. For import

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[9]To evaluate sensitivity of the price data to the method used to collect prices, Gopinath and Rigobon (2008) compared the behavior of prices during the anthrax scare of 2001, when the IPP data were collected by phone survey rather than mail, to the behavior of prices during other time periods. They find no significant differences in the frequency of price change over the period when prices were collected by phone. However, even in the phone survey, reporters were provided with the price they had previously reported—perhaps inducing a similar bias as in the mail survey.

and export prices, this approach would lead to difficulties in repricing some types of products since the products imported by a firm are found to differ slightly from one shipment to the next. To deal with this problem, the IPP takes a more pragmatic approach. In cases where there has been a substantial change in quality, the IPP discontinues the former item and initiates a new item with a new description. The product is also replaced if it is not possible to quantify the magnitude of the quality change. To avoid sacrificing repricability, however, the IPP may deem some product characteristics to be non-price determining. This pragmatic approach is likely to lead to some spurious price changes that reflect changes in the product, rather than true price changes.[10]

Hedonic adjustments for quality are infrequent in the IPP data. However, in some cases, the IPP attempts to adjust observed price changes for changes in quality using estimates of the associated changes in costs. Our analysis of substitutions focuses only on substitutions for which it is not possible to do this type of quality adjustment.

The sampling weights used in the IPP indices were, until 1997, specified at the level of detailed product groupings, called "classification groups". Within a classification group, products were sampled with a probability equal to their trade weight, eliminating the need for more disaggregated weights. For the purpose of calculating the frequency of price change and substitutions and constructing aggregate indices, we construct item-level weights using a procedure analogous to the one that the IPP used until 1997: by dividing the classification group weight by the total number of "net price" observations used in the BLS index. In this regard, our analysis differs from earlier analyses by Gopinath and Rigobon (2008) and Gopinath, Itskhoki, and Rigobon (2008), which do not incorporate sampling weights.

Since 1997, the IPP has introduced product-level weights below the level of individual classification groups. The historical product-level weights are unfortunately not available. The most detailed weights available are classification group weights. However, an IPP staff study in 1997 found that the change in methodology to the new item weights did not have a large effect on the index. This suggests that our procedure for constructing weights based on the pre-1997 methodology is likely to be reasonably accurate

For the regression analysis in our robustness analysis in section 6, we found that using weights at this detailed a level introduced a substantial amount of sampling error in our results. We therefore opted to calculate item-level weights as the average weight within a given HS2 code over the entire

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[10]Substitutions in the IPP dataset are conceptually the same as non-comparable substitutions in the CPI data. The CPI Research Database records information about the frequency of comparable substitutions. In contrast, the IPP Research Database does not record the frequency of comparable substitutions.

sample period and use these weights in our regression analysis.[11] In section 6, we also carry out some regressions that condition on a price change having occurred. To avoid overweighting product categories with a high frequency of price change, we calculate average item-level weights at the HS2-code level for these regressions by dividing the classification group weight by the total number of price change observations in that HS2 code.

We carry out most of our analysis at an aggregate level. However, in some cases, we also report statistics for the following sectoral groupings based on HS2 codes: animal products (01-05); vegetable products (06-15), foodstuffs (16-24), mineral products (25-27), chemicals and allied industries (28-38), plastics and rubber products (39-40), raw hides, skins, leather and furs (41-43), wood and wood products (44-49), textiles (50-63), footwear and headgear (64-67), stone and glass (68-71), metals (72-83), machinery and electrical (84-85), transportation (86-89) and miscellaneous items (90-97).

The second set of data we use is from the U.S. NIPA. We use the import price deflator for imported goods excluding oil. We use the export price deflator for exported goods excluding agricultural products. We use data for the period 1982 through 2007 since the IPP was introduced in 1982 and the price deflators before that time are based on unit value data. Finally, we make use of trade-weighted monthly and daily exchange rates downloaded from the Federal Reserve Board's website at http://www.federalreserve.gov/releases/h10/Hist/ as well as monthly bilateral exchange rates from the International Financial Statistics (IFS) database of the IMF.

3 Prices and Exchange Rates: Evidence

Measuring the degree of pricing to market for U.S. imports involves comparing the price of imports into the U.S. relative to all domestic prices and the U.S. real exchange rate.[12] A simple way to measure the degree of pricing to market for U.S. imports is to run the following cointegrating regression:

$$p_t^m - p_t = \alpha + \gamma t + \beta q_t + \epsilon_t, \tag{1}$$

where p_t^m denotes the log of the dollar price of U.S. imports, p_t denote the log of the dollar price of U.S. production and q_t denotes the log of the trade weighted U.S. real exchange rate. Let e_t denote the log of the trade weighted U.S. nominal exchange rate (U.S. dollar price of foreign currency)

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[11]There are roughly 95 HS2 code groups, while there are roughly 6900 classification groups.

[12]Knetter (1989) and Marston (1990) pioneered empirical estimation of pricing to market.

and p_t^* denote the log of the foreign currency price of foreign production. The real exchange rate is defined as $q_t = e_t + p_t^* - p_t$. The estimated degree of "exchange rate pass-through" is given by β. We interpret "incomplete pass-through" as "pricing to market". The estimated degree of pricing-to-market is thus given by $1 - \beta$. Micro studies of pricing-to-market control for differences in costs in ways that we are not able to do with aggregate data. The incomplete pass-through we estimate may thus partly be due to differences in costs across locations. Goldberg and Knetter (1997) emphasize the distinction between pass-through and pricing-to-market.

To capture potential dynamics in the relationship between movements in import prices and the real exchange rate, we consider the following vector error correction model (VECM),

$$\Delta y_t = \Pi(A y_{t-1} + \alpha + \gamma t) + \sum_{k=1}^{n-1} \Gamma_k \Delta y_{t-k} + \delta + \epsilon_t, \tag{2}$$

where y_t is the vector $(p_t^m - p_t, \ q_t)$, and A is a vector of coefficients in the cointegrating relationship given by $[1 \ -\beta]$. The parameter β is therefore a measure of the long-run responsiveness of import prices to exchange rates. We find strong evidence of a cointegrating relationship between import prices and the real exchange rate. We reject the null hypothesis of no cointegrating equations using the Johansen trace statistic method (Johansen, 1995). The Schwarz Bayesian Information Criterion selects one lag in the vector error correction model, so we set $n = 2$.

The third specification we consider is the following "dynamic adjustment" specification:

$$\Delta(p_t^m - p_t) = \alpha + \sum_{k=0}^{6} \beta_k \Delta q_{t-k} + \epsilon_t. \tag{3}$$

This is the type of regression that is typically run in the exchange rate pass-through literature (e.g., Campa and Goldberg, 2005; Marazzi and Sheets, 2007). Long-run pass-through is then defined as the sum of the coefficients, $\hat{B} = \sum_{k=0}^{6} \beta_k$. This specification is misspecified if relative import prices and the real exchange rate are cointegrated.

For U.S. exports, measuring the degree of pricing to market involves comparing prices of U.S. exports relative to all foreign prices and the U.S. real exchange rate. The regression equations we use to make this comparison for U.S. exports are analogous to equations (1)-(3). First we consider:

$$p_t^{x*} - p_t^* = \alpha + \gamma t - \beta q_t + \epsilon_t, \tag{4}$$

where p_t^{x*} denotes the foreign currency price of U.S. exports. We construct $p_t^{x*} - p_t^*$ as $p_t^x + q_t - p_t$, where p_t^x is the dollar price of U.S. exports. Second, we consider a VECM analogous to equation (2) but with $y_t = (p_t^{x*} - p_t^*, \ q_t)$. Again, we reject the null hypothesis of no cointegrating equations

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using the Johansen trace statistic method. The Schwarz Bayesian Information Criterion selects two lags in the vector error correction model, so we set $n = 3$. We allow for a structural break in the cointegrating relationship for exports in 2004 by adding a dummy variable that is equal to one in the first quarter of 2004 and thereafter into the vector y_t. This dummy variable accounts for an apparent level shift in the cointegrating relationship between export prices and the real exchange rate after 2004, as we discuss below. Finally, we consider the following "dynamic adjustment" specification for exports,

$$\Delta(p_t^{x*} - p_t^*) = \alpha - \sum_{k=0}^{6} \beta_k \Delta q_{t-k} + \epsilon_t. \tag{5}$$

Results for these six regressions are presented in table 1. We use the NIPA price deflator for non-oil goods imports and non-agricultural goods exports. Our sample period is from 1982 through 2007. We begin our sample in 1982 because this is when the import and export price indices were introduced in the U.S. The exchange rate variable we use is the Federal Reserve's trade weighted real exchange rate index for major currencies.[13] We use consumer prices as our proxies for p_t and p_t^*.

We find that the dynamic adjustment regression and the VECM yield similar conclusions regarding the long-run relationship between trade prices and real exchange rates. For imports, the VECM yields an estimate of β of 0.41 while the dynamic adjustment equation yields an estimate of 0.43 for the sum of the coefficients $\hat{B}$. The levels regression yields a somewhat lower estimate of 0.36, presumably because a substantial fraction of the adjustment of trade prices is lagged by 2-3 quarters relative to the real exchange rate adjustment. These estimates are broadly in line with the existing literature on exchange rate pass-through. For example, Campa and Goldberg (2005) estimate long run exchange rate pass-through for U.S. imports to be 0.42 for the period 1975 to 2003. For export prices, all three models yield estimates within the tight range of 0.85 to 0.87.[14]

Figures 2 and 3 display the stability of the relationship documented above. Figure 2 plots the detrended relative dollar price of U.S. imports $p_t^m - p_t$ and the fitted values based on the cointegrating relationship, $\hat{\beta} q_t$. The two series are normalized to have the same means and detrended. Figure

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[13]These indices are similar to the index used in Campa and Goldberg (2005). The major currency exchange rate series seems more appropriate than the broader index for two reasons. First, the weights in the import and export price index are often 3-5 years out of date. This implies that the growing role of countries outside the group of major currencies is captured only with a substantial lag and is therefore small for the majority of our sample period. Second, the major exchange rate index is potentially more similar to an index of non-oil import prices. However, we also estimated these regressions with the Federal Reserve's broad exchange rate indices and we discuss the results for these alternative regressions below.

[14]If we instead use the Federal Reserve's broad exchange rate index, the VECM yields 0.50 for imports and 0.84 for exports, while the dynamic adjustment regression yields 0.52 for imports and 0.84 for exports.

3 plots analogous series for the case of exports. Over this time period, these relationships—both for imports and exports—have been quite stable aside from the upward level shift in the price of exports after 2003 noted above.

Several researchers have argued that exchange rate pass-through into U.S. imports has fallen in recent years (e.g., Olivei, 2002; Marazzi and Sheets, 2007). Table 2 reports pass-through estimates for U.S. imports from both the VECM and the dynamic adjustment regression for two subsamples: 1982-2008 and 1994-2008. For the dynamic adjustment regression, pass-through is indeed estimated to be quite a bit lower in the recent subsample—0.32 compared to 0.43 over the longer sample period. The results for the VECM, however, suggest that this apparent fall in pass-through might partly be due to model misspecification. For the VECM, the long-run pass-though estimate is slightly higher for the recent sample than it is for the longer sample—0.46 versus 0.41.

4   Prices and Exchange Rates: Theory

Consider an economy in which consumers purchase and consume products from a continuum of product lines. At each point in time, one product from each product line exists. Let C_{jit} denote the number of units of the current product from product line j produced in region i and consumed at time t and let γ_{jit} denote the quality of each of these units measured in terms of utility. Products from country i enter the consumer's utility function through the following consumption aggregator

$$C_{it} = \left[\int_{N_i} (\gamma_{jit} C_{jit})^{\frac{\theta-1}{\theta}} dj \right]^{\frac{\theta}{\theta-1}},$$

where N_i denotes the set of goods consumed from country i. Let P_{jit} denote the price per unit of the current product from product line j in region i at time t. The price index that gives the minimum cost of an additional unit of utility is then given by

$$P_{it} = \left[\int_{N_i} \left(\frac{P_{jit}}{\gamma_{jit}} \right)^{1-\theta} dj \right]^{\frac{1}{1-\theta}}. \tag{6}$$

On the firm side of the economy, firm j in region i produces goods according to the following production function

$$C_{jit} = \gamma_{jit}^{-1} F(K_{jit}, L_{jit}).$$

The function F is homogeneous of degree one in capital, K_{jit}, and labor, L_{jit}. Quality enters the production function multiplicatively. This implies that to raise the quality of its products by a factor ξ and produce the same number of units the firm must employ ξ times as much of each

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factor input. In other words, it costs the firm twice as much to produce goods that are twice as good in utility terms.

Now let $\hat{C}_{jit} = \gamma_{jit} C_{jit}$ denote effective consumption of product line j at time t and let $\hat{P}_{jit} = \gamma_{jit}^{-1} P_{jit}$ denote the corresponding effective price. Using these concepts, it is possible to rewrite both the consumer problem and the firm problem entirely without reference to γ_{jit}. In particular, the consumption aggregator and the price index become

$$C_{it} = \left[\int_{N_i} \hat{C}_{jit}^{\frac{\theta-1}{\theta}} dj \right]^{\frac{\theta}{\theta-1}} \quad \text{and} \quad P_{it} = \left[\int_{N_i} \hat{P}_{jit}^{1-\theta} dj \right]^{\frac{1}{1-\theta}}.$$

And the production function becomes

$$\hat{C}_{jit} = F(K_{jit}, L_{jit}).$$

This implies that the equilibrium allocations generated by models with this set of assumptions about product quality are the same as those of any number of standard models in the macroeconomics and international economics literature. The equilibrium allocations generated by standard models simply refer to effective consumption and effective prices in the corresponding model with changing product quality. Changing product quality, however, complicates the empirical evaluation of standard models if product quality is not observed. In this case, the models generate data on effective prices and quantities, while the real world data consists of raw prices and quantities. Below, we consider how the behavior of observed data on prices and quantities differs from the behavior of effective prices and quantities when product quality is unobserved.

Suppose that the prices of newly introduced products are fully flexible. For continuing products, however, prices are sticky in the local currency. Assume that firm j adjusts the price of continuing products with probability $f_j(s_t)$ in period t, where s_t denotes a state variable. It is not important for our purposes to describe what features of the economic environment govern the state-dependence of the frequency of price change. We therefore leave this unspecified. The frequency of price change across different product lines has some distribution $\Phi_t(s_t)$ that is also a function of the state of the economy.

Suppose that product replacement occurs in each product line with probability $z(s_t)$ in period t. For simplicity, assume that each time a product is replaced, a new γ_{jit} is drawn from a distribution $\Gamma_t(s_t)$ and this level of quality remains constant for the product until it undergoes another replacement. In other words,

$$\gamma_{jit} \begin{cases} \sim \quad \Gamma_t(s_t) & \text{with probability } z(s_t) \\ = \quad \gamma_{jit-1} & \text{otherwise} \end{cases}$$

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The distribution of product quality, $\Gamma_t(s_t)$, has no impact on our results since firm profits and consumer welfare depend only on quality adjusted prices and since inflation depends on price relatives $(\hat{P}_{jt}/\hat{P}_{j,t-1})$ for which quality drops out.[15]

Let lower case variables denote logarithms of upper case variables. A first order Taylor-series approximation of equation (6) yields

$$p_{it} = \int_{N_i} \hat{p}_{jit} dj. \tag{7}$$

First differencing this equation yields

$$\Delta p_{it} = \int_{N_i} \Delta \hat{p}_{jit} dj. \tag{8}$$

Assume for simplicity that the price index in equation (7) is the price index that the BLS seeks to construct when it calculates the U.S. import and export price indices.[16] A major complication in constructing this type of price index is the fact that product quality is unobserved. The ideal solution would be to use hedonic methods to estimate product quality. However, such methods are rarely used because they are extreme costly and difficult to apply in most cases. In practice, price comparisons that involve a change in quality are usually dropped from the index. Indices constructed in this way are referred to as "matched model indices" since all price comparisons on which such indices are based are for identical items.

4.1 Product Replacement Bias: A Factor Calculation

Consider the following regression for the change in log prices,

$$\Delta p_t = \alpha + B\Lambda_t + \epsilon_t, \tag{9}$$

where B is a vector of coefficients and Λ_t is a vector of aggregate variables. Given equation (8), it is straight-forward to show that the vector of regression coefficients for this regression, B, may be "decomposed" as follows

$$B = \int_S \int_{N_i} B_j(s) dj ds, \tag{10}$$

where $B_j(s)$ denotes the regression coefficient from estimating equation (9) for product lines with the same characteristics in terms of price adjustment behavior and product turnover and for time

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[15]Our model implies that desired markups are unrelated to movements in real exchange rates. If movements in the real exchange rate are systematically associated with movements in markups for new products—perhaps because of a combination of systematic movements in quality and non-CES demand—then the types of biases we consider associated with product replacement may be further exacerbated. See Auer and Chaney (2009) and Rodriguez-Lopez (2008).

[16]This price index is a special case of the Tornqvist index with fixed and equal weights.

periods for which the aggregate state is s. This equation allows us to analyze each set of product lines j and each state of the world s separately and then take an average over products and states.[17]

Suppose that data on γ_{jit} were available and that it were therefore possible to estimate equation (9) using *all* observations on changes in effective relative prices for product line j as the dependent variable. Our assumptions about price rigidity and product replacements imply that the effective price of product line j changes with probability $f_j(s) + z(s) - f_j(s)z(s)$. We can divide the observations on price changes and exchange rates for this regression into two groups based on whether $\Delta \hat{p}_{jit} = 0$ or $\Delta \hat{p}_{jit} \neq 0$. Let $B_j(s)$ denote the vector of regression coefficients for equation (9) using all available observations, while $B_j^{ch}(s)$ denotes the vector of coefficients based only on the observations for which $\Delta \hat{p}_{jit} \neq 0$. In this case, it is easy to show that $B_j(s) = (f_j(s) + z(s) - f_j(s)z(s))B_j^{ch}(s)$ asymptotically.

Now, consider the case where data on γ_{jit} are not available. In this case, the effective price change is not observed when a product replacement occurs. Consider estimating equation (9) using data from the remaining observations and let $B_j^{mm}(s)$ denote the vector of coefficients from this regression. Price changes occur in a fraction $f_j(s)$ of these observations implying $B_j^{mm}(s) = f_j(s)B_j^{ch}(s)$ asymptotically. We thus have

$$B_j^{mm}(s) = f_j(s)B_j^{ch}(s) = \frac{f_j(s)}{f_j(s) + z(s) - f_j(s)z(s)}B_j(s). \tag{11}$$

In other words, $B_j^{mm}(s)$ yields an estimate of the true regression coefficients (defined as $B_j(s)$) that is downward biased by a factor $f_j(s)/(f_j(s) + z(s) - f_j(s)z(s))$.

Integrating over j and s yields

$$B^{mm} = \int \int \frac{f_j(s)}{f_j(s) + z(s) - f_j(s)z(s)}B_j(s)djds.$$

This implies that the matched model index results in an estimate of B that is biased by a factor

$$\frac{B}{B^{mm}} = \frac{\int \int B_j(s)djds}{\int \int \frac{f_j(s)}{f_j(s) + z(s) - f_j(s)z(s)}B_j(s)djds}. \tag{12}$$

It is instructive to consider a few special cases. First, if the frequency of price change, the frequency of product substitution and desired pass-through for each product type j are constant over time, equation (12) simplifies to

$$\frac{B}{B^{mm}} = \frac{\int B_j dj}{\int \frac{f_j}{f_j + z - f_j z}B_j dj}. \tag{13}$$

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[17]Here we must assume that for each state of the world we have data from enough time periods that $B_j(s)$ is identified. This implies that the state space for the frequency of price change and the frequency of substitutions must be somewhat "coarser" than the state space for the aggregate variable Λ_t.

If, in addition, desired pass-through, B_j, is constant across product types, the factor simplifies further to

$$\frac{B}{B^{mm}} = \left[\int \frac{f_j}{f_j + z - f_j z} dj \right]^{-1}. \tag{14}$$

In some cases we may be interested in the sum of the regression coefficients. One such instance is the dynamic adjustment regression—equation (3). Define $B(\text{sum})$ to be the sum of the elements of B. Since $B(\text{sum})$ is a linear function of the individual regression coefficients, we have,

$$\frac{B(\text{sum})}{B^{mm}(\text{sum})} = \left[\int \frac{f_j}{f_j + z - f_j z} dj \right]^{-1}. \tag{15}$$

The intuition for this expression is simple. The magnitude of product replacement bias depends on the relative frequency of price changes for identical items (numerator) and price changes including product replacements (denominator). Since the function $f_j/(f_j + z - f_j z)$ is concave, product replacement bias will be greater the greater is the amount of heterogeneity in the frequency of price change across products. This bias expression also shows that both price rigidity and product turnover must occur for product replacement bias to arise. If $f_j = 1$ or $z = 0$, there is no bias.

4.2   Product Replacement Bias and Pricing to Market

We now apply the product replacement bias factor adjustment derived above to pricing to market regressions. Consider first the dynamic adjustment regression—equation (3). The factor calculation does not in general apply directly to this regression since the dependent variable in this regression is the change in the *relative* price of imports, $\Delta(p_t^m - p_t)$. However, the factor calculation does provide a reasonable approximation when the covariance of p_t with the real exchange rate is small enough that equation (3) yields similar results with Δp_t^m as with $\Delta(p_t^m - p_t)$ as the dependent variable. This is true in practice for U.S. data.[18]

Deriving the extent of product replacement bias for the VECM is somewhat more involved. Under the assumptions that p_t is uncorrelated with the real exchange rate and that the real exchange rate and the relative price of imports are integrated of order one, we can show that the bias in β in the VECM—equation (2)—is given by equation (12). This derivation is presented in appendix A.

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[18]If p_t is positively correlated with the real exchange rate (prices rise when the real exchange rate depreciates), equation (12) understates product replacement bias, whereas with a negative correlation, equation (12) overstates product replacement bias. In practice, there is a slight positive correlation between prices and the real exchange rate.

4.3 Discussion

The assumption that prices are more responsive to the economic environment for newly introduced products than for continuing products is a crucial feature of our model. Matched model price indices drop the price changes that occur for newly introduced products. If prices are more likely to change at these points than at other points, dropping these observations leads to a selection bias in the measurement of the relationship between prices and aggregate variables. Matched model price indices will then underestimate the responsiveness of prices to aggregate variables since they disproportionately drop price change observations.

One way to view matched model price indices is as implicitly imputing a price change for the effective price of newly introduced products that is equal to the average change in the price of all continuing products. In our model this method of price imputation does not accurately capture the price dynamics of newly introduced products. The effective price change for newly introduced products has the same distribution as the effective price change of all continuing products *that change their price* in that period in our model. This implies that the change in the effective price of newly introduced products is much larger on average than the average price change of all continuing products.

In our baseline model, we assume that prices are equally flexible at the time of product replacements as at the time of price change. A more general model would allow for differences in the pricing behavior of firms when products are introduced versus when the prices of continuing products change. In such a model, the extent of product replacement bias would be given by a generalized factor:

$$\frac{f_j(s)}{f_j(s) + \alpha z(s) - \alpha z(s) f_j(s)}, \tag{16}$$

where α represents the difference in the pricing behavior of firms at the time of product introductions versus when they change the price of continuing products. One example of a more general model would be a model in which firms have both a menu cost and a quadratic cost of changing prices for continuing products but no such quadratic cost when setting the first price of new products. This type of pricing behavior could arise because of implicit contracts or consumer antagonism (Nakamura and Steinsson, 2008; Rotemberg, 2005). In this case, α is greater than one and the product replacement bias is larger than in our baseline model. On the other hand, if prices change less at the time of product replacements or if some products are introduced without the firm reoptimizing their prices, α is less than one and product replacement is smaller than in our baseline

model. We investigate the value of α empirically in section 5 and also discuss the quantitative sensitivity of our results to alternative values using equation (16).

One potential approach to solving the problem of product replacement bias in our model economy is to calculate the price index simply as the weighted average of all prices including both new and continuing products. Since this procedure makes use of all price comparisons, it avoids the selection bias problem associated with dropping price changes associated with new product introductions. There are two problems with this approach. First, it means we must compare the prices of entirely different products—say, last year's wool jacket versus this year's down coat. In a world of highly heterogeneous products, these price comparisons introduce a large amount of sampling error into the price index. In the IPP data, we find that an import price index calculated from average prices in this way is extremely noisy: the simple average of prices for imports and exports routinely fluctuates by 10-20% per month. The massive amount of sampling error in this type of index generates sufficiently large standard errors in the estimated relationship between an average price index and the exchange rate that almost nothing can be concluded about the nature of pricing-to-market. To match this feature of the data in our model, we assume that the distribution Γ is sufficiently dispersed that an index based on average prices yields an unacceptably large amount of sampling error. A second problem with analyzing average price indexes is that there may be a behavioral relationship between quality and the exchange rate. For example, if the exchange rate appreciates, consumers may switch toward higher quality products. This could bias upward the estimated relationship between prices (per unit quality) and the exchange rate.[19]

If movements in the exchange rate were exogenous, our estimate of exchange rate pass-through would represent a causal link from the exchange rate to prices. In this case, $B < 1$ could indicate that as the U.S. dollar appreciates versus foreign currencies, firms adjust their prices to stabilize it in local currency terms. In equilibrium, however, exchange rates vary as a consequence of demand and supply shocks that also affect firm's costs and market shares directly. For instance, if exchange rate movements arise largely from monetary shocks then the covariance between exchange rates and import prices may be low even if the price response to an exogenous exchange rate movement is much larger (Corsetti, Dedola and Leduc, 2008; Bouakez and Rebei, 2008). Our analysis of product replacement bias does not rely on a particular model of equilibrium exchange rate determination. Our objective is to improve on the measurement of the equilibrium relationship between prices and exchange rates.

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[19]See e.g., Ghironi and Melitz (2005) for a more detailed discussion of this issue.

Many of the products for which prices are sampled in the IPP data are intermediate products sold to other firms. An important question in interpreting the evidence on price rigidity for imported goods is whether the observed rigid prices are "allocative" (Barro, 1977). Since manufacturers and retailers interact repeatedly, the observed price in a particular month may not actually determine purchasing decisions. Rather, this price may be an "installment payment" on a "running tab" that adjusts continuously but is unobservable. We do not address this issue, since our focus is on documenting rather than interpreting the observed relationship between prices and exchange rates. However, it is worth noting that this phenomenon is less likely to influence the long run relationship between prices and exchange rates than it is to affect the short term dynamics of this relationship.

5   Product Replacement Bias: Measurement

Before the introduction of the IPP, import and export price indices were based on unit value data for highly disaggregated categories. This practice was criticized because it did not control for changes in quality and composition within these categories. An important reason for the introduction of the IPP at the BLS was to control more adequately for quality and composition and thereby measure pure price changes. The IPP has therefore taken great care in the way it defines a product. The definition of a product in the IPP data includes not only a unique product identifier such as a bar code, but also other "price determining characteristics" identified by the BLS such as the terms of the transaction, size of the shipment and in some cases even the identity of the seller. This is a much finer definition of a product than, e.g., 5 digit SIC codes or 10 digit HS codes, which are commonly used in the international trade literature (Bernard, Redding and Schott, 2008). We adopt the product definitions in the IPP. A product, as we use the term, is therefore often a contract between a particular buyer and seller. A new product is not necessarily totally new to the world but rather new to a particular buyer-seller interaction. Carlton (1986) provides evidence that defining a product in this way is important when studying the flexibility of prices.

In the model presented in section 4, product replacement bias arises because prices are disproportionately flexible at the time of product replacements. There are several reasons why we argue this assumption is reasonable in our data. First, a large fraction of product substitutions occur because one product is discontinued, to be replaced by a new product. Specifically, about 60% of the substitutions in the BLS data arise because a product a particular firm is importing

or exporting changes.[20] Second, firms tend to negotiate new prices when they sign new contracts with their customers and this is also the time when many products are initiated into the dataset.

Third, as we describe in section 2, the prices of products that are newly initiated into the BLS dataset are collected using a detailed personal interview, while the prices of continuing products are collected using a "repricing form" that first asks whether the price has changed relative to the previous month and then ask the respondent to report a new price if the price did change. This practice may generate a differential reporting friction for reporting a fresh price for newly initiated products as opposed to continuing products, since, for continuing products, the easiest response is to simply check the box indicating "no change" in price.[21] A differential reporting friction of this kind implies that newly initiated products have fresher prices than continuing products.

Table 3 presents direct evidence on the degree to which the prices of newly introduced products have been recently reset. This evidence is based on the idea that the size of a particular price change should be more related to exchange rate changes that have occurred during the price spell that ends with that price change than to earlier exchange rate changes. Panel A of table 3, reports results for the following regression:

$$\Delta p_{jk} = \alpha + \beta_S \Delta e_{jk,S} + \beta_{1Q} \Delta e_{jk,1Q} + ... + \beta_{6Q} \Delta e_{jk,6Q} + \epsilon_{jk}, \tag{17}$$

where Δp_{jk} denotes the log size of the kth price change for product j, Δe_{jkS} denotes the log change in the exchange rate over the course of the kth price spell for product j and $\Delta e_{jk\#Q}$ denotes the log change in the exchange rate over the course of the $\#$th quarter prior to the kth price spell for product j. We run this regression for the first and second price changes ($k = 1$ and 2) of all products that have exactly two price changes. We run these regressions separately for import and exports.

Figure 5 provides a graphical illustration of the "first" and "second" price change regressions

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[20]It is difficult to estimate the fraction of substitutions that involve a version change or upgrade. The dataset contains a flag indicating whether a product substitution is due to such a version change or upgrade. However, there are at least two reasons why this flag is unreliable. First, for most of the time period we study, to qualify as a version change or upgrade, the replacement product must fall into the same HS10 category. Since these categories are extremely disaggregated, it often happens that the replacement product falls in a different HS10 code. For example, male cows and female cows have different HS10 codes as do VHS players and DVD players. Second, BLS economists have indicated to us that many product discontinuations are followed by reinitiations of similar products by a BLS field agent. This may happen because firms find it easier to simply discontinue a product than to report the details of a replacement product to the BLS.

[21]One way of interpreting the timing of price changes in our model is as reflecting the timing of price reports as opposed to price adjustments. If newly initiated products have freshly reported prices while continuing products don't, then the assumptions of our model hold and product replacement bias arises. This is true even if all the observed price rigidity is due to infrequent price reporting rather than true price rigidity. We would like to thank Virgiliu Midrigan for suggesting this model of price rigidity in survey data.

presented in columns 1-2 of Table 3—i.e. equation (17) for $k = 1$ and 2. While we do not observe how much prices change when a new good is introduced into the dataset, we can observe how responsive subsequent price adjustments are to exchange rate movements that occurred before the good was introduced. If the first price change for each product is more strongly related to exchange rate movements that occur before that product's introduction into the data set than the second price change is to exchange rate changes that occurred before the first price change, this would suggest that the initial prices of products in our data were not newly reset. In fact, we find no evidence of this.

For imports, the pattern of coefficients is very similar for both the first and second price change. In both cases β_S is larger than 0.2. The coefficients then fall rapidly over the first two quarters before the price spell in question and are insignificant in most cases after that. In both cases, the difference between β_S and β_{1Q} are insignificant, while β_{2Q} is significantly different from β_S. There is no evidence that the first price change responds more to exchange rate changes before first price spell than the second price change responds to exchange rate changes before the second. The pattern is similar for exports. The main difference being that the coefficients are smaller—reflecting lower measured pass-through for exports—and that β_{2Q} is negative rather than positive.

Given the pattern of coefficients reported in panel A, we ran the following more parsimonious specification:

$$\Delta p_{jk} = \alpha + \beta_{S+1}\Delta e_{jk,S+1} + \beta_{2-4Q}\Delta e_{jk,2-4Q} + \epsilon_{jk}, \tag{18}$$

where $\Delta e_{jk,S+1}$ denotes the log change in the exchange rate over the course of the kth price spell and one quarter before this price spell for product j and $\Delta e_{jk,2-4Q}$ denotes the log change in the exchange rate over the course of the 2nd, 3rd and 4th quarter prior to the kth price spell for product j. This regression again yields very similar results for the first and second price changes for both imports and exports. For imports, β_{S+1} is larger than 0.2 while β_{2-4Q} is less than 0.1. In both regressions, the difference between the first and second coefficients is significantly different from zero. There is no evidence that the first price change responds more to the second regressor that the second price change. If anything, the opposite is true. For exports, the size of the coefficients on both regressors are again smaller than in the case of imports. The size of the second coefficient is again smaller than the first in both cases and the difference between them is significantly different from zero for the first price change regression.

These results can be thought of as providing evidence on the appropriate value of α in equation (16). The lack of evidence of any difference between the first price change and the second price

change suggests that $\alpha = 1$ is a reasonable assumption. To be conservative, we assume that the first price spell for each product in our data begins three months before the product enters the data set. This accounts for the fact that it is difficult to statistically distinguish the coefficients on the current and lagged quarters in the equation (17).

5.1 Results

We show in section 4 that product replacement bias is most severe when the frequency of product replacements is large relative to the frequency of price change. Table 4 reports our estimates of the weighted fraction of products that have less than or equal to 0, 1, 2, 3 and so on price changes. For LCP imports, 41% of products have no price changes, while 68% have two or fewer price changes. For PCP exports, 51% of products have no price changes, while, 78% have two or fewer price changes. These statistics motivate the idea that product replacement bias may be an important phenomenon in import and export price data.[22]

The small number of price changes per product reflects substantial price rigidity in the microdata on import and export prices collected by the BLS. Table 5 reports the weighted mean and median frequency of price change per month for imports and exports, separately for LCP and PCP goods. These statistics parallel those reported in Gopinath and Rigobon (2008), though our analysis differs somewhat from theirs in that we study a longer time period, and make use of product-level weights.[23]

Table 5 shows that both imports and exports exhibit substantial price rigidity. Most U.S. imports are local currency priced (93%). For these goods, the mean monthly frequency of price change is 14.1% and the median is 6.7%. Most U.S. exports are producer currency priced (98%). For these goods, the mean monthly frequency of price change is 11.7% while the median is 6.3%. Table 5 also reports the weighted mean frequency of product substitutions for imports and exports. This fraction varies from 4.7-5.7% across the four classes of goods considered.[24]

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[22]Our estimate of the fraction of price spells with no price change is somewhat higher than the estimate of Gopinath, Itskhoki, and Rigobon (2008). Their estimate for LCP imports is around 30%, while our estimate is 41%. Most of the difference arises because our estimate is for the entire dataset, while theirs is for a subset of OECD countries. Another difference is that our estimates incorporate product-level weights.

[23]We calculate the frequency of price change by constructing an indicator variable for whether a price change occurred and taking the mean of this variable. We calculate the frequency of product substitutions as the total number of product substitutions observed in the data, divided by the total number of periods that the price series are observed. The series we use for this are constructed by "filling in" the previously observed price through the large number of missing spells in the import price data as we discuss in section 2. This is appropriate given the imputation procedure used by the BLS and described in section 2. All of the statistics we report are calculated as weighted averages using the item-level weights described in section 2.

[24]One concern regarding our results might be that producers have a small number of list prices that do not change

Heterogeneity in the frequency of price change is an important determinant of the quantitative impact of product replacement bias as we discuss in section 4.1. A simple way to assess the degree of heterogeneity in the frequency of price change is to present results on this statistic for different industry groups. Table 6 reports the mean frequency of price change and the mean frequency of product replacement across different industry groups of imports and exports.[25] The frequency of price change varies widely—from 42% for Animals & Animal Products to less than 5% for several other categories. In contrast, the average frequency of product replacement varies much less across industry groups. Most industry groups have frequencies of product replacement of 3-6%. In what follows, we will assume for simplicity that the frequency of product replacement is the same for all goods. We will, however, take explicit account of the large amount of heterogeneity in the frequency of price change.[26]

While a great deal of heterogeneity in the frequency of price change is embodied in differences across sectors, there is also much heterogeneity within each sector. Ignoring intra-sector heterogeneity would seriously bias our estimate of product replacement bias. We therefore estimate a flexible distribution for the overall heterogeneity in the frequency of price change across products. Suppose that a product j has a constant hazard of adjusting its price, f_j, in each month. Suppose also that $f_j \sim \text{Beta}(a, b)$. We denote the product's lifetime by n_j. Given that the product has a constant hazard of adjusting its price, the total number of price changes x_j are distributed according to a binomial distribution, i.e., $x_j \sim \text{Bin}(n_j, f_j)$.[27]

In appendix B, we derive a simple expression for the log-likelihood function in this setting. We estimate this model by maximum likelihood for four groups of products: LCP imports, PCP

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when a new relationship is formed with a customer. However, this is at odds with the evidence presented in Carlton (1986) that prices appear to differ dramatically across customers. Furthermore, in our data, the frequency of product substitutions is slightly higher for exports than for imports. This is interesting because the export price data is gathered from sellers while the import price data is gathered from buyers. If sellers have list prices which apply to a large number of customers for each product, one might expect substitutions to be more frequent in data gathered from buyers than in data gathered from sellers. Each time the buyer switched products, a product substitution would occur in import price data. But in the export price data the BLS would continue to sample the product as long as there was another buyer buying at the same price. This suggests that there is a great deal of price dispersion across different buyers for identical products.

[25]In table 6, for imports, Mineral Products excludes petroleum productions. For exports, we report statistics for Animals & Animal Products, Vegetable Products and Foodstuffs even though these three categories are excluded from our weighted average statistics.

[26]Product replacement bias is especially important for a number of durable goods categories such as autos, furniture and computers. For autos, the frequency of price change in the IPP is 6.8%, while the frequency of substitution is 5.1%. For furniture, these frequencies are 8.2% and 4.4%, respectively, and for computers they are 13.7% and 5.8%, respectively.

[27]We have also considered the robustness of our results to using a sectoral model in which the frequency of substitutions is allowed to vary across sectors and the distribution of the frequency of price change across products within each sector is a different beta distribution for each sector. This model yields very similar results.

imports, LCP exports and PCP exports. Table 5 reports our estimates for the parameters of the beta distribution. For LCP imports, the estimated parameters are $a = 0.50$ and $b = 3.65$. These parameters imply a very large amount of heterogeneity in the frequency of price change across products. Figure 4 plots the cumulative distribution function of the distribution Beta$(0.50, 3.65)$. The mean of this distribution is 0.121 while the standard deviation is 0.144. For PCP exports, the estimated parameters are $a = 0.53$ and $b = 4.60$. The mean of this distribution is 0.103 while the standard deviation is 0.123.

Given the estimates in table 5, we can use equations (13) to produce estimates of the factor by which pricing to market is mismeasured because of product replacement bias. These estimates are reported in table 7. We assume for empirical tractability that the frequency of price change and the frequency of substitutions are constant over time for each product.[28] We carry out these calculations for two sets of assumptions regarding heterogeneity in true pass-through across goods with different frequencies of price change. First, we assume that true long-run pass-through is constant within each of the four groupings considered in table 5. For this case, the factor implied by equation (13) for LCP imports is 1.82, while the factor for PCP exports is 1.93. Second, we consider a case in which true pass-through is lower for products with a lower frequency of price change. Gopinath and Itskhoki (2008) argue that this pattern exists in the data. Their estimates suggest that true pass-through for LCP imports with a frequency of price change below about 25% per month may be only about 65% of true pass-through for LCP imports with a higher frequency of price change.[29] Under this assumption, the factors implied by equation (13) for LCP imports and PCP exports are 1.72 and 1.84, respectively.

Using these factors, we can recompute the comovement between prices and exchange rates adjusting for product replacement bias. The results of these calculations for the VECM are reported in the lower panel of table 1. Since product replacement bias affects LCP and PCP goods differently,

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[28]Empirical evidence suggests that these are reasonable assumptions for the particular application we study. We regressed the frequency of product replacements for dollar-priced imports and exports on the absolute magnitude of log movements in the trade-weighted exchange rate for the years 1995-2006 (we drop 2007 because only part of a year is available). The resulting coefficient is -0.023 (0.131) for imports and -0.078 (0.308) for exports, where we report standard errors in parentheses. For periods and countries for which exchange rate variation was more dramatic, there may be a stronger relationship between the frequency of product replacement and the real exchange rate. Burstein, Echenbaum, and Rebelo (2005) document clear evidence of a rise in the number of products that ceased to be imported into Argentina at the time of Argentina's 2000-2002 financial crisis and devaluation. In this case, it would be important to account for time variation in the frequency of price change and frequency of substitutions when calculating the adjustment factor.

[29]This difference in measured true pass-through could alternatively arise due to spurious price changes in the micro data. See section 2 for a discussion of reasons why spurious price changes may exist in the micro data on imports and exports.

we must calculate this effect separately for LCP and PCP goods and then take a weighted average. Gopinath, Itskhoki, and Rigobon (2008) argue that there is a large difference in pass-through between LCP and PCP imports. We allow for this difference in our calculations. We adopt their estimate of 0.94 for measured pass-through of PCP imports and use this in our calculation of aggregate pass-through adjusted for product replacement bias.[30] For exports, virtually all products are PCP. So, any reasonable heterogeneity across LCP and PCP products makes virtually no difference. As in table 7, we present results on pass-through adjusting for product replacement bias for both the case of no heterogeneity in true pass-through for goods with differing frequencies of price change and a case calibrated based on the results of Gopinath and Itskhoki (2008).

Adjusting for product replacement bias raises the comovement of U.S. import prices (relative to all U.S. prices) and the U.S. real exchange rate from 0.41 to between 0.66 and 0.69. And it lowers the comovement between U.S. export prices (relative to all foreign prices) and the U.S. real exchange rate from 0.87 to roughly 0.79. This improvement in measurement brings the data closer in line with standard models along two dimensions. First, our adjusted estimates imply a much more similar degree of pricing to market for U.S. imports and exports. Second, adjusting for product replacement bias implies more modest pricing to market for U.S. imports. This improves the fit of the data to models such as Corsetti and Dedola (2005) and Atkeson and Burstein (2008). These authors' preferred calibrations of the models in their papers imply exchange rate pass-through of roughly 0.9 and 0.75, respectively. Other models can generate more pricing to market (Gust, Leduc and Vigfusson, 2006, Gopinath, Itskhoki and Rigobon, 2007). But to do so they must introduce a great deal of strategic complementarity. This implies that pass-through becomes extremely slow in these model. In contrast, empirical evidence suggests that the bulk of pass-through occurs in less than one year (Campa and Goldberg, 2005; Marazzi and Sheets, 2007). Third, our adjusted estimates of pricing to market for U.S. imports and exports are also more in line with industry studies. These generally find substantially higher estimates of pass-through than recent estimates based on aggregate data.[31]

The calculations above assume that α in equation (16) is equal to one. This is consistent with the empirical evidence in table 3. For robustness, we can consider alternative values of α. The parameter α in equation (16) could be less than one if prices are sticky at the time of product

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[30]Given this assumption about the pass-through of PCP imports we assume a degree of measured pass-though for LCP imports that yields our estimate of measured pass-through of 0.41 for all imports. Reasonable variations in this set of assumptions have negligible effects on our results.

[31]See, e.g., Goldberg and Knetter (1997), Feenstra (1989), Goldberg and Verboven (2001), Hellerstein (2008), Nakamura and Zerom (2009), Goldberg and Hellerstein (2008).

introductions; or greater than one if firms use product introductions disproportionately to adjust prices. We have calculated the effect of product replacement bias for a range of different values of α. The effects of product replacement turn out to be non-linear in α. The marginal effect of increasing α is highest for low values of α and then falls. This implies that if $\alpha = 0.5$, product replacement bias is 2/3 as large as in the benchmark model that assumes $\alpha = 1$. If $\alpha = 0.75$, product replacement bias is 85% as large as if $\alpha = 1$. And if $\alpha = 1.5$, product replacement bias is 125% as large as if $\alpha = 1$.

Table 8, presents results for several alternative data and measurement assumptions. In our baseline results, we use the trade weighted "Major Currency" U.S. real exchange rate published by the Federal Reserve Board. Table 8 presents alternative results using the Fed's "Broad" real exchange rate series. Measured pass-through is somewhat higher for imports and slightly lower for exports using the Broad real exchange rate. Using this measure and adjusting for product replacement bias yields a pass-though estimate of roughly 0.81 for imports and 0.73 for exports. Also, in our baseline results, we count all substitutions. Table 8 presents alternative results where we count only forced substitutions. This alternative calibration yields a pass-through estimate of 0.6 for imports and has minor effects on our pass-through estimate for exports.

6   Alternative Measures of Comovement

Not all measures of exchange rate pass-through are affected by product replacement bias. Using the micro price data that underlie the U.S. import and export price indices, it is possible to avoid the problems associated with product replacement under certain assumptions. Consider regressing changes in relative prices on the change in the real exchange rate since the time of the previous price change:

$$\Delta(p_{jit} - p_t) = \alpha + \beta_\Delta \Delta^* q_t + \epsilon_{jit}. \tag{19}$$

where Δ^* is a difference operator representing the difference between the current real exchange rate and the real exchange rate at the time of the previous price change of product j (or in the case of the first price change of product j, the introduction of product j). In a simple model of price adjustment in which firms' optimal prices depend only on current exchange rates, the coefficient β_Δ in this regression provides a measure of the long-run relationship between prices and exchange rates that is unaffected by product replacement bias.

Another way to measure long-run pass-through is to run the following "levels" regression:

$$p_{jit} - p_t = \alpha_j + \gamma t + \beta q_t + \epsilon_{jit}, \tag{20}$$

on the subsample of observations for which a price change or product introduction occurs. Again, in a model in which firms' optimal prices depend only on current exchange rates, the regression coefficient β provides a measure of the long-run relationship between prices and exchange rates that is unaffected by product replacement bias.[32]

Table 9 presents estimates of these two regressions based on U.S. import price micro-data for the period 1995-2007, the years in which a full set of data are available in the IPP database. We estimate these regressions for dollar-priced non-oil imports for a subset of high income OECD countries: Canada, Sweden, U.K., Netherlands, Belgium, France, Germany, Switzerland, Spain, Italy, Japan. Observations from these countries account for the majority of observations in the IPP data set. For each country, we match the price data to exchange rate and consumer price data for that country from the IFS database of the IMF. The real exchange rate variable q_{it} is the end-of-month exchange rate for the month previous to the month identified for the reported price. We estimate both regressions using weighted OLS. The weights in the regressions are constructed at the HS2-code level to avoid overweighting HS2 codes with particularly large numbers of price changes. This procedure is described in more detail in section 2.[33]

For equation (19), we estimate $\beta_\Delta = 0.26$. For equation (20), we estimate $\beta = 0.54$. One important difference between these regressions is their differential reliance on the assumption that the price change at time t depends only on the exchange rate movement since the last price change. In particular, equation (19) relies much more heavily on this assumption than equation (20) (Griliches and Hausman, 1986). This pattern of regression coefficients, therefore, suggests that price adjustments may respond to exchange rate changes further in the past than the last price change.

There are a number of reasons why price adjustments in a particular time period may depend on exchange rate movements before the last price change, invalidating the use of β_Δ and β as measures of long-run comovement. One reason is the presence of strategic complementarities in price-setting (Gopinath, Itskhoki and Rigobon, 2007). Another reason is measurement error in the

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[32]This regression yields similar results as the "life-changes" regression run in Gopinath, Itskhoki, and Rigobon (2008).

[33]We have also run these regressions on only price change observations, i.e. excluding the first price change for each product because the previous price change is not observed for these observations. This yields similar estimates. The similarity of these estimates is consistent with our assumption that, for a given product, pure price changes and product replacements generate similar responses to exchange rate movements.

timing of price observations in the IPP data. One source of such timing error is that the prices requested by the IPP are the prices of products *received* by the firm as close as possible to the reference date. Production lags and delivery lags may therefore imply that this price will refer to a product ordered at a substantially earlier point in time. Furthermore, while the IPP requests that reporters provide a price for the transaction that occurs as close as possible to the first day of the month, in practice, importers and exporters often go for long periods of time without importing or exporting. As a consequence, reporters often provide prices for other days in the month. In some cases, average monthly prices are provided. Also, the large amount of price imputation done in the IPP and discussed in section 2 is an additional source of timing error.

In the remainder of this section, we analyze whether the results of the regressions described above, as well as standard measures of pricing to market, can be explained by a simple model with timing error and product replacement bias. The objective is to provide a concrete example of how product replacement bias affects various measures of the comovement between prices and exchange rates. However, it is important to note that the analysis in this section is much less general than in the previous sections. Models with strategic complementarities in price-setting, forecastable movements in exchange rates or other general equilibrium factors that cause prices to adjust slowly to exchange rates may imply a different relationship between prices and exchange rates than the simple data generating process considered here. We choose to study this simple case to focus attention on timing error and product replacement bias.

As in our previous analysis, firm j adjusts its price with probability $f_j \sim \text{Beta}(a, b)$. The timing of product replacement is governed by a Weibull distribution with shape parameter equal to five. The key new structural assumption is that conditional on adjusting its price, the firm sets its price as follows:

$$p_{jit} - p_t = \phi q_{it} + \eta_{jit}, \tag{21}$$

where η_{jit} is orthogonal to the exchange rate. We assume that η_{jit} follows the stochastic process,

$$\eta_{jit} = \mu + \rho \eta_{jit-1} + \epsilon_{jit}, \tag{22}$$

where $\epsilon_{jit} \sim \text{N}(0, \sigma_\epsilon^2)$.

We define true pass-through as in previous sections as the long-run adjustment of prices to an exchange rate shock. In this model, true pass-through is given by ϕ. As in our previous analysis, we allow for heterogeneity in ϕ across products that is correlated with the frequency of price change.

To capture the type of timing error discussed above, we assume that the price that the IPP

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records for a firm at time t is the price that firm charged at a different time $t + u_t$. Specifically, we assume that $p_t^r = p_{t+u_t}$, where p_t^r denotes the price recorded by the IPP for the firm at time t and u_t reflects random timing error. We assume that the timing error u_t has two components, $u_t = u_{1,t} + u_{2,t}$. The first component $u_{1,t}$ is distributed $u_{1,t} \sim \text{Unif}[-1, 0]$. This term reflects the simple fact that the reported price changes for a particular month occur randomly over the course of the preceding month, but are observed at discrete intervals.

The second term $u_{2,t}$ is distributed, $u_{2,t} = -kd_{jit}$, where d_{jit} is the time in months since the last price change, and $k \sim \text{Unif}[0, K]$. This term is motivated by the various forms of timing error discussed above, which lead recorded prices to be somewhat "stale". On average, we assume that the reported prices are "stale" by a fraction k of the duration since the last reported price.[34]

6.1 Parameters

We calibrate the frequency of price change and substitutions to match data from the subset of high income OECD countries discussed above. For this subset of countries, the weighted average frequency of product substitution is 4.67% and the estimated distribution of the frequency of price change is $\text{Beta}(0.52, 3.35)$. We use actual daily observations on the U.S.-German exchange rate over the time period 1995-2007. When we simulate the model, we use the fractional values generated by the timing error model described above to infer on which day of the month a price change occurs.[35] For simplicity, in our simulations we assume that the difference between home and foreign inflation is constant. As in our earlier analysis, we follow Gopinath and Itskhoki (2008) in assuming that products with a frequency of price change below 25% per month have a pass-through parameter of $\phi = 0.65\phi_{high}$. We set $\rho = 0.5$ based on previous estimates in Nakamura and Steinsson (2008a) for consumer price data, and we set σ_ϵ^2 to match the average size of price changes in the data (in practice, these assumptions have little impact).

The remaining parameters are the timing error parameter K and the parameter ϕ. We use a simulated method of moments procedure to estimate these parameters. The moments we use in this procedure are the coefficients of the regression equations (19)-(20), β_Δ and β, as well as the coefficient β_2 obtained from re-estimating equation (20) on only the subset of products that have exactly two price changes.

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[34]In the case of the first observed price, we set d_{jit} to the unconditional duration of prices for the product in question. We never allow the price for the current month to be older than the time of the last observed price change.

[35]For robustness, we have carried out analogous experiments using the Canadian, Japanese, and U.K. exchange rates. These experiments yield almost identical results.

Our estimation procedure is to select the values of K and ϕ that minimize the sum of the squared deviations between β_Δ, β and β_2 in the simulated and actual data. In our baseline estimation procedure, we weight the three moments equally. Since we are able to come very close to exactly matching the actual moments in the data, the choice of a weighting matrix makes little difference to our results.

This estimation procedure yields $K = 0.96$ and $\phi = 0.48$. The estimated value of K implies that, on average, delivery lags and other sources of timing error account for a delay in price reporting of about 48% of the average duration since the last price change or product replacement. This corresponds to an average reporting lag of about 4 months, which is consistent with existing estimates of delivery lags (e.g., Abel and Blanchard, 1988). The estimated value of ϕ implies that the low frequency of price change products in our model have a desired pass-through of 0.48, while the high frequency of price change products have a desired pass-through of 0.73, yielding aggregate pass-through of 0.52.

6.2   Results

Table 9 compares results from the data with results from the estimated timing error model discussed in this section. The timing error model fits the data well. The first three rows of the table present results for the three moments used to estimate the model: β_Δ, β and β_2. We match these statistics almost exactly. Notice that in both the model and the data, β_Δ yields a measure of pass-through that is much lower than β. The assumed amount of heterogeneity in desired pass-through is also quantitatively successful in matching the observed difference between β and β_2 in the data.

In addition to the moments used in the estimation, the model fits the fact that pass-through is much higher for the subset of goods that have eight or more price changes (row 4 in table 9). The high value of β for this subset of goods largely reflects a combination of higher desired pass-through and diminished importance of timing error for this subset of goods. The model also fits the data for the micro dynamic adjustment equation—equation (3) that has individual price observations as the dependent variable (row 5 of table 9) The covariates in this regression are the current change in exchange rates, one future change and 22 lagged changes. The discrepancy between the results from this regression and true long-run pass-through arises as a consequence of product replacement bias.[36]

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[36]The fact that measured pass-through from the dynamic adjustment equation for the LCP imports is 0.26—substantially lower than the aggregate estimate of 0.43 reported in section 3—reflects the fact that our analysis of the microdata covers a substantially shorter sample period and includes only LCP goods.

7 Other Implications of Product Replacement Bias

Our analysis has focused on the effect that product replacement bias has on the relationship between exchange rates and import and export price indices. However, product replacement bias may arise in other indices and in response to other shocks. In particular, product replacement may cause apparent long-run failures of purchasing power parity. Consider a situation where the nominal U.S. exchange rate versus Canada depreciates by 20%. Suppose that U.S. consumer prices rise by 20%, implying that the real exchange rate between the U.S. and Canada remains constant. Consumer prices may, nevertheless, appear to increase by less than 20% due to product replacement bias, leading the real exchange rate to appear to have permanently changed. Product replacement bias may also cause problems in estimating the response of import and export prices to increases in outsourcing (see, e.g., Houseman, 2007; and Mandel, 2007 and 2009).

While we do not study unit value indices in this paper, many of the measurement concerns we explore here also apply to unit value indices. In particular, for standard unit value indices, price comparisons are often dropped because of lack of availability of data for the previous period (for example, because the product was not traded in the previous period), potentially leading to product replacement bias. Also, large price changes are often excluded as outliers or trimmed. As a consequence, Alterman (1991) estimates that the U.S. unit value indices, produced in 1985, were calculated for only 56 percent of the value of imports and 46 percent of the value of exports.

Mismeasuring import and export price indices also affects measured trade volumes and trade price elasticities. Holding fixed nominal quantities, if the increase in import prices in response to an exchange rate depreciation is underestimated, then the corresponding decline in import quantities will be underestimated as well. If the corresponding price and quantity series are used to estimate trade price elasticities, then the resulting elasticity estimates will be biased away from one. If the true trade price elasticity is less than or equal to one, then the estimated price elasticity will be biased toward zero.

8 Conclusion

This paper argues that the simultaneous presence of import price rigidity and frequent product replacements may lead to serious problems in measuring the relationship between exchange rates and prices. Since many imports and exports exhibit only a small number of price changes over their lifetimes, we argue that the comovement of import and export prices and exchange rates may be

mismeasured by as much as a factor of two. We refer to the resulting bias as "product replacement bias". Adjusting for product replacement bias makes the behavior of import and export prices easier to reconcile with standard models. The degree of pricing to market for U.S. imports and exports is much more symmetric, and the degree of pricing to market for U.S. imports more moderate than conventional estimates suggest.

A   Product Replacement Bias in the VECM

Consider estimating equation (9) on a price index of all product lines with a frequency of price change f_j and with time dummies for each quarter as the aggregate variables z. We refer to all product lines with a frequency of price change f_j as sector j. The coefficient on each dummy variable gives the average change in prices in sector j for that quarter. Based on equation (8), it is straightforward to show that the coefficients B_j equal Δp_{it}^j, the measured change in the price index for sector j. Equation (11) implies that these measured changes, Δp_{it}^j, are all biased by the factor $f_j/(f_j + z - f_j z)$. Since the price index p_{it}^j is constructed by stringing together these changes, the true price index for sector j is equal to the measured price index for sector j multiplied by $f_j/(f_j + z - f_j z)$. This implies that for each sector of the economy, the estimated cointegrating vector in the VECM is biased by the factor $f_j/(f_j + z - f_j z)$.

To calculate the overall bias in the estimated cointegrating vector, assumption that the cointegrating vectors across different sectors of the economy are the same. Without loss of generality, assume that the estimated cointegrating vector for sector j is $[1 \; \beta_j^{mm}]$. Define $\beta^{mm} = \int_{N_i} \beta_j^{mm} dj$ and notice that

$$p_{it} - \beta^{mm} q_t = \int_{N_i} (p_t^m - p_t) - \beta_j^{mm} q_t dj.$$

This implies that $p_{it} - \beta^{mm} q_t$ is stationary. The aggregate cointegrating vector is therefore $[1 \; \beta^{mm}]$. The ratio of the true long run relationship between relative import prices and the real exchange rate, β, and the measured value β^{mm} from the VECM is thus

$$\frac{\beta}{\beta^{mm}} = \left[\int \frac{f_j}{f_j + z - f_j z} dj \right]^{-1}. \tag{23}$$

B Log-Likelihood in the Presence of Unobserved Heterogeneity in the Frequency of Price Change

We assume that product i has a constant hazard of adjusting, f_i, in each month, where $f_i \sim$ Beta(a, b). Let us denote the product's lifetime by n_i. These assumptions imply that the total number of price changes in a product's lifetime is distributed according to the binomial distribution, $x_i \sim \text{Bin}(n_i, f_i)$. We assume, furthermore, that f_i is distributed according to the beta distribution, $f_i \sim \text{Beta}(a, b)$.

Given this model, we can write the likelihood of observing a product with length n_i and the total number of price changes x_i as,

$$L = \prod_{i=1}^{I} \frac{\Gamma(a+b)}{\Gamma(a)\Gamma(b)} f_i^{a-1}(1-f_i)^{b-1} \binom{n_i}{x_i} f_i^{x_i}(1-f_i)^{n_i-x_i} \tag{24}$$

$$= \prod_{i=1}^{I} \frac{\Gamma(a+b)}{\Gamma(a)\Gamma(b)} f_i^{x_i+a-1}(1-f_i)^{n_i-x_i+b-1} \binom{n_i}{x_i} \tag{25}$$

We can integrate out the $f_i's$ to get,

$$L = \prod_{i=1}^{I} \frac{\Gamma(a+b)}{\Gamma(a)\Gamma(b)} \binom{n_i}{x_i} \frac{\Gamma(a+x_i)\Gamma(b+n_i-x_i)}{\Gamma(a+b+n_i)}. \tag{26}$$

The log-likelihood function is, therefore,

$$\log L = n\log\Gamma(a+b) - n\log\Gamma(a) - n\log\Gamma(b)^n + \sum_{i=1}[\log n_i! - \log x_i! \tag{27}$$

$$- \log(n_i - x_i)! + \log\Gamma(a+x_i) + \log\Gamma(b+n_i-x_i) - \log\Gamma(a+b+n_i)]. \tag{28}$$

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TABLE I

Pricing to Market

	Imports	Exports
Measured:		
VECM	0.41	0.87
	(0.05)	(0.06)
Dynamic Adjustment	0.43	0.85
	(0.05)	(0.05)
Levels	0.36	0.86
	(0.02)	(0.03)
Adjusting for Product Replacement Bias:		
No Heterogeneity in Comovement	0.69	0.78
With Heterogeneity in Comovement	0.66	0.79

The top panel presents alternative measures of the long-run relationship between the trade-weighted real exchange rate and aggregate import or export price indices (standard errors in parentheses). "VECM" reports the estimated coefficient on exchange rates in the cointegrating relationship between prices and exchange rates (equation (2)). "Dynamic Adjustment" reports the sum of the coefficients on lagged exchange rate changes in a linear model for exchange rate changes with 6 quarterly lags (equation (3)). "Levels" reports the coefficient on the exchange rate in a simple OLS regression of the price index on the trade-weighted exchange rate (equation (1)). The bottom panel presents the estimated relationship based on the VECM model adjusted for "product replacement bias" according to the methods discussed in the paper, under alternative assumptions about heterogeneity in pricing behavior across products.

TABLE II

Pricing to Market over Subsamples

Period	VECM	Dynamic Adj.
1982-2008	0.41	0.43
	(0.05)	(0.05)
1994-2008	0.46	0.32
	(0.08)	(0.08)

The table presents alternative measures of the long-run relationship between trade-weighted exchange rate series and aggregate import or export price indices for the time periods 1982-2008 and 1994-2008 (standard errors in parentheses). "VECM" reports the estimated coefficient on exchange rates in the estimated cointegrating relationship between prices and exchange rates (equation (2)). "Dynamic Adjustment" reports the sum of the coefficients on lagged exchange rate changes in a linear model for exchange rate changes with 6 quarterly lags (equation (3)).

TABLE III
Price Change for First and Second Spell on Exchange Rate

	Imports		Exports	
	First Price Change	Second Price Change	First Price Change	Second Price Change

Panel A:

Exchange rate change:

During price spell that is ending	0.23	0.26	0.13	0.11
	(0.03)	(0.03)	(0.03)	(0.03)
1st quarter before price spell that is ending	0.19	0.18	0.03	0.14
	(0.05)	(0.04)	(0.07)	(0.06)
2nd quarter before price spell that is ending	0.13	0.11	-0.06	-0.02
	(0.04)	(0.04)	(0.06)	(0.06)
3rd quarter before price spell that is ending	-0.01	0.05	0.14	0.15
	(0.05)	(0.04)	(0.06)	(0.05)
4th quarter before price spell that is ending	0.07	0.09	-0.09	0.09
	(0.04)	(0.04)	(0.05)	(0.05)
5th quarter before price spell that is ending	0.05	0.04	0.19	0.08
	(0.04)	(0.04)	(0.06)	(0.06)
6th quarter before price spell that is ending	0.14	0.06	0.06	-0.01
	(0.05)	(0.04)	(0.07)	(0.06)
P-value for the null that coefficients for price spell and 2nd qrt before price spell are equal	0.052	0.002	0.008	0.059

Panel B:

Exchange rate change:

During the price spell that is ending and 1 quarter before this spell	0.21	0.24	0.09	0.12
	(0.02)	(0.02)	(0.03)	(0.03)
2nd to 4th quarter before the price spell that is ending	0.06	0.09	-0.02	0.08
	(0.02)	(0.02)	(0.03)	(0.03)
P-value for the null that the two coefficients are equal	0.000	0.000	0.021	0.340

Panel A presents coefficients from OLS regressions of price changes on current and lagged exchange rate changes, including the exchange rate change during the price spell that is ending, and exchange rate changes over the 1st to 6th quarters before this price spell (standard errors in parentheses). The table also reports the p-value testing the null that the coefficients on the current price spell and the 2nd quarter before the price spell are equal. Panel B reports the results of similar regressions where the regressors are 1) the exchange rate change during the current price spell and the quarter before the spell and 2) the exchange rate change over the 2nd-4th quarters before the spell started. The p-value for the null that these coefficients are equal is also reported.

TABLE IV
Number of Price Changes Per Product

Number of	Imports		Exports	
Price Changes	LCP	PCP	LCP	PCP
0	40.6	42.8	42.2	51.0
1 or less	57.2	65.1	61.6	68.8
2 or less	67.7	79.4	72.5	78.0
3 or less	74.6	86.0	79.3	84.6
4 or less	79.3	91.4	83.6	88.7
5 or less	82.3	94.4	86.3	94.2
10 or less	90.0	98.1	93.1	98.6
15 or less	93.1	99.4	95.4	98.9
20 or less	94.7	99.5	97.0	98.9

The table presents the fraction of products in the BLS microdata on import and export price data with less than or equal to a given number of price changes over the entire timespan for which they are in the data set. These statistics are reported for both local currency priced (LCP) and producer currency priced (PCP) products. The statistics are weighted percentiles, using as weights the cumulative product-level weights over each product's lifetime.

TABLE V
The Distribution of Price Changes and Substitutions

	Imports		Exports	
	LCP	PCP	LCP	PCP
Fraction of Imports/Exports	0.926	0.074	0.029	0.971
Mean Frequency of Price Change	0.141	0.074	0.087	0.117
Median Frequency of Price Change	0.067	0.036	0.035	0.063
Mean Frequency of Substitutions	0.048	0.047	0.057	0.052
Distribution of the Frequency of Price Change				
a	0.50	1.07	0.76	0.53
	(0.01)	(0.07)	(0.04)	(0.01)
b	3.65	19.12	7.76	4.60
	(0.06)	(1.32)	(0.83)	(0.11)

The top panel reports summary statistics for the mean and median frequency of price change and product substitution calculated using IPP microdata on import and export prices. Statistics are reported for both local currency priced (LCP) and producer currency priced (PCP) products. The weighted means are and medians are calculated using the item-level weights described in the paper. The lower panel reports our estimates of "a" and "b" which are the parameters in the estimated distribution of the frequency of price change, assumed to be Beta(a,b). This distribution is estimated using the BLS microdata on imports and exports.

TABLE VI

Sectoral Frequency of Price Change and Product Substitution

Industry Group	LCP Imports			PCP Exports		
	Freq. PC	Freq. Subs.	Weight	Freq. PC	Freq. Subs.	Weight
Animals & Animal Products	0.420	0.034	0.025	0.418	0.046	0.028
Vegetable Products	0.411	0.059	0.022	0.289	0.064	0.021
Foodstuffs	0.159	0.032	0.036	0.108	0.038	0.032
Mineral Products	0.120	0.091	0.007	0.190	0.047	0.033
Chemicals & AlliedIndustries	0.124	0.044	0.054	0.149	0.053	0.062
Plastics / Rubbers	0.186	0.045	0.038	0.162	0.038	0.046
Raw Hides, Skins, Leather, & Furs	0.060	0.043	0.019	0.279	0.064	0.003
Wood & Wood Products	0.338	0.038	0.080	0.210	0.040	0.054
Textiles	0.053	0.066	0.089	0.173	0.055	0.024
Footwear / Headgear	0.047	0.047	0.046	0.017	0.484	0.000
Stone / Glass	0.221	0.025	0.070	0.113	0.082	0.020
Metals	0.215	0.052	0.064	0.223	0.058	0.052
Machinery / Electrical	0.095	0.054	0.204	0.077	0.057	0.260
Transportation	0.087	0.052	0.143	0.067	0.042	0.274
Miscellaneous	0.046	0.044	0.103	0.053	0.079	0.092
Weighted Average	0.141	0.048	1.000	0.117	0.052	1.000
Weighted Median Across Industries	0.095	0.052	1.000	0.077	0.057	1.000

The table reports statistics on the frequency of price changes and product substitution based on BLS microdata on import and export prices. All statistics are weighted means using product level weights. "Freq. PC" is the frequency of price change, "Freq. Subs" is the frequency of substitutions and "Weight" is the total weight on the industry group. Statistics are presented for local currency priced (LCP) imports and producer currency priced (PCP) exports.

TABLE VII

Product Replacement Bias

	LCP	PCP
No Heterogeneity in Comovement		
Imports	1.82	2.11
Exports	1.92	1.93
With Heterogeneity in Comovement		
Imports	1.72	2.10
Exports	1.87	1.84

The table presents the estimated adjustment factor for product replacement bias resulting from the analysis in the paper. An adjustment factor of 2 implies that conventional measures of comovement are understated by a factor of 2. The results are presented for LCP (local currency priced) and PCP (producer currency priced) imports and exports under alternative assumptions regarding heterogeneity in pricing behavior across products.

TABLE VIII

Pricing to Market and Product Replacement Bias: Robustness

	Major Country RER		Broad RER	
	All Subs	Forced Subs	All Subs	Forced Subs
Measured Pass-Through				
Imports	0.41		0.50	
	(0.05)		(0.07)	
Exports	0.87		0.84	
	(0.06)		(0.06)	
A. No Heterogeneity in Comovement				
Adjusted Pass-Through				
Imports	0.69	0.62	0.85	0.77
Exports	0.78	0.82	0.72	0.76
Factors				
LCP Imports	1.82	1.61	1.82	1.61
PCP Imports	2.11	1.58	2.11	1.58
LCP Exports	1.92	1.59	1.92	1.59
PCP Exports	1.93	1.61	1.93	1.61
B. With Heterogeneity in Comovement				
Adjusted Pass-Through				
Imports	0.66	0.60	0.81	0.74
Exports	0.79	0.82	0.73	0.77
Factors				
LCP Imports	1.72	1.54	1.72	1.54
PCP Imports	2.10	1.57	2.10	1.57
LCP Exports	1.87	1.56	1.87	1.56
PCP Exports	1.84	1.55	1.84	1.55

The table presents the estimated adjustment factor for product replacement bias under various alternative assumptions. "All Subs" measures the frequency of product substitutions using all observed substitutions, while "Forced Subs" includes only forced substitutions. Panel A and B present results for different assumptions regarding heterogeneity in pricing behavior across products. The results are presented for LCP (local currency priced) and PCP (producer currency priced) imports and exports. The first two columns present results for the "Major Country" real exchange rate (RER) while the measure in the last two columns present results for the "Broad" real exchange rate measure.

TABLE IX
Results for Simulated Data with Timing Error

	LCP Imports	
	Real Data	Simulated Data
Change in Exchange Rate Since Last Price Change	0.26	0.25
	(0.02)	--
Levels Conditional on a Price Change	0.54	0.54
	(0.02)	--
Conditional on 2 Price Changes	0.24	0.25
	(0.05)	--
Micro Dynamic Adjustment Equation	0.27	0.31
	(0.02)	--
Conditional on 8 or more Price Changes	0.69	0.70
	(0.06)	--

The table reports regression coefficients from analogous regressions run using simulated and actual data. The "Real Data" statistics are based on calculations using BLS microdata on import prices denominated in dollars for high income OECD countries. The "Simulated Data" statistics are based on output from our simulation model. The first row reports the regression coefficient in a regression of price changes on exchange rate movements over the period when the price was fixed. The second and third row report the regression coefficient of prices on exchanges rates conditioning only on periods when a price change occurred. The fourth and fifth row report the sum of the regression coefficients from a regression of price changes on past exchange rate changes. A detailed discussion of these regressions is presented in section 6.
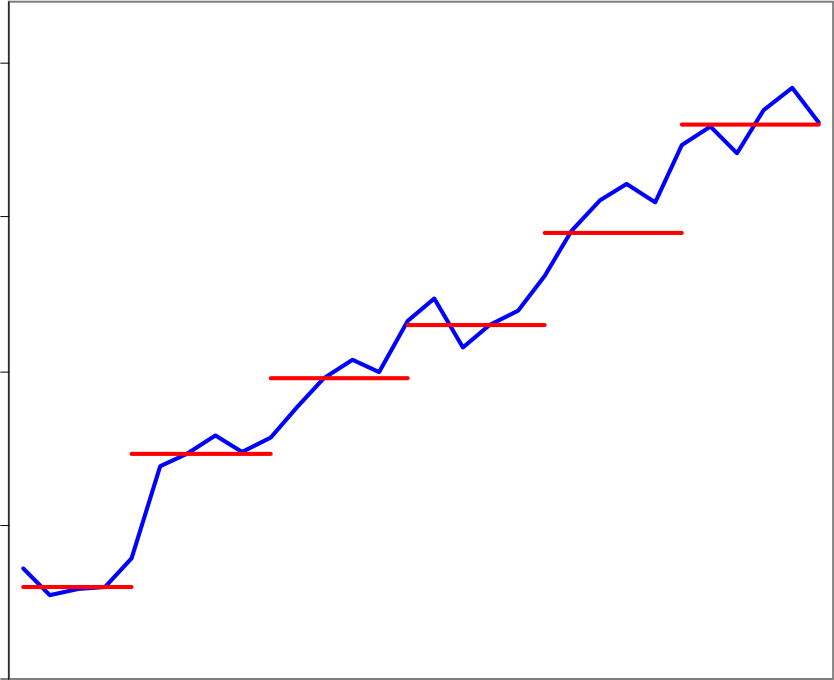
Figure I
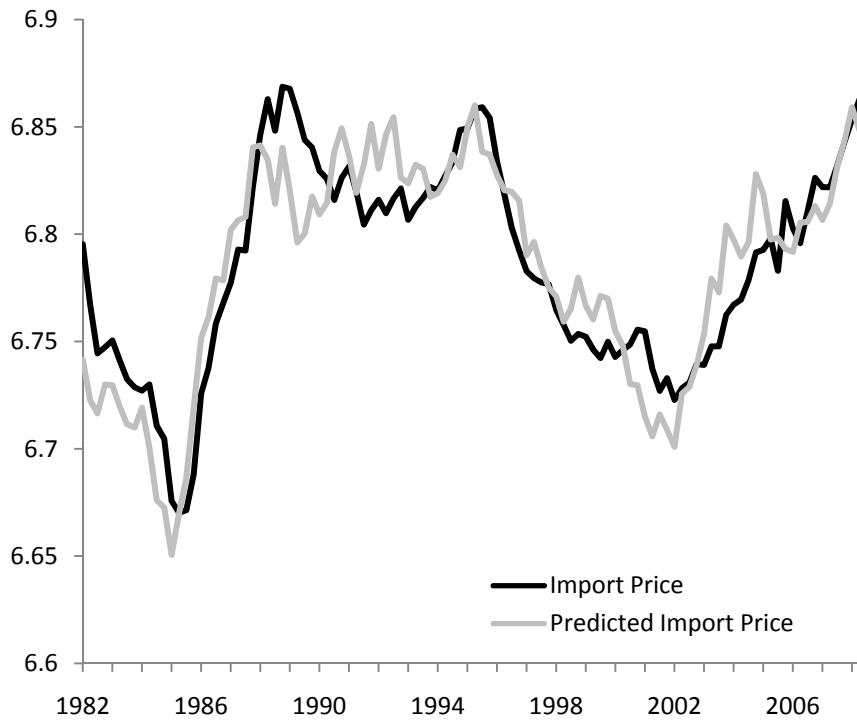Product Replacement and the Comovement of Prices and Exchange Rates

Figure II
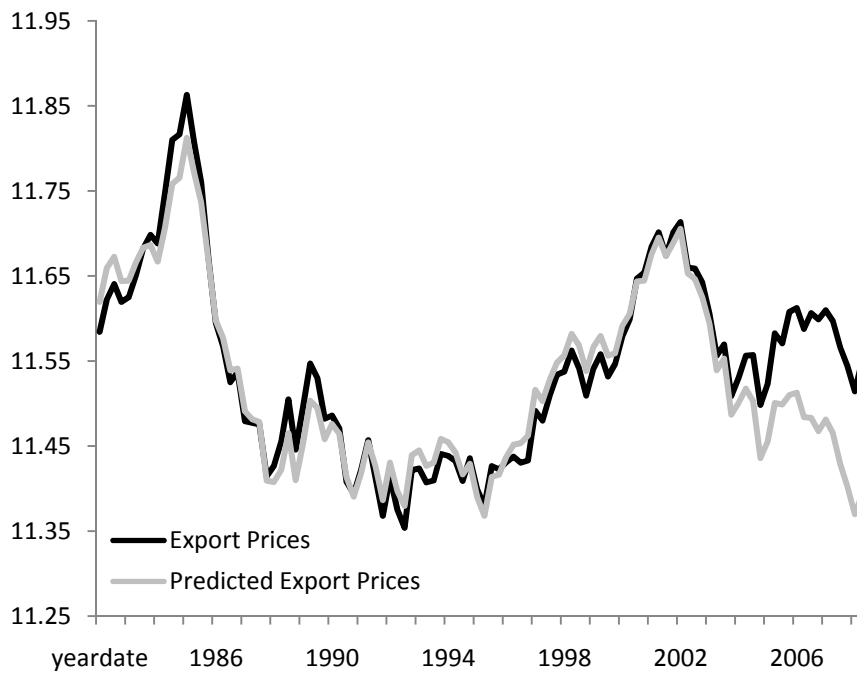U.S. Import Prices and the Real Exchange Rate

Figure III
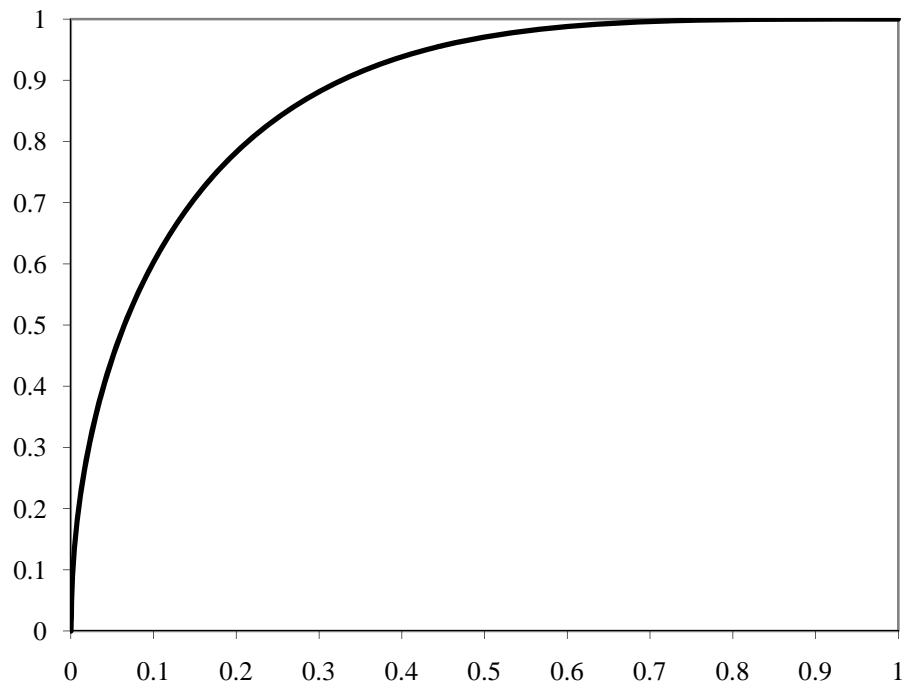U.S. Export Prices and the Real Exchange Rate

Figure IV
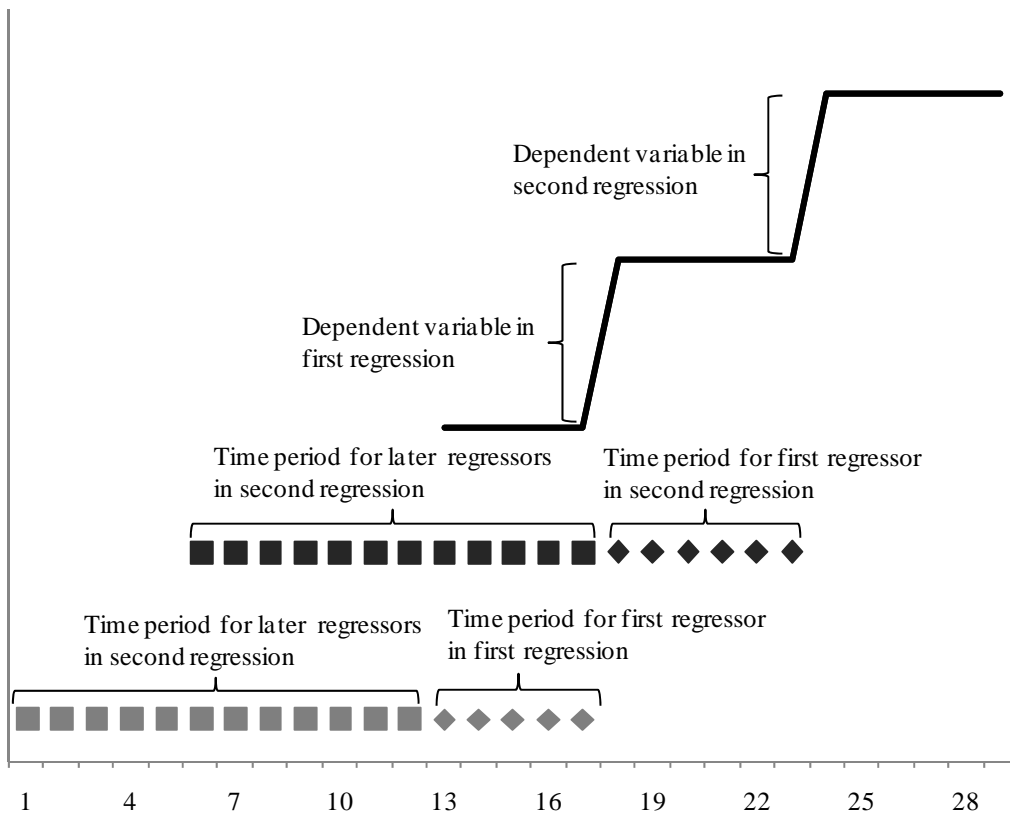Cumulative Probability Distribution of Beta(0.50,3.65)

Figure V

Graphical Depiction of Regressions in Table III

This figure provides a graphical depiction of the regression equation presented in equation (17). The solid line denotes the price of a product that has two price changes. The first regression has as its dependent variable the first price change and as explanatory variables exchange rate movements over the first price spell as well as the six quarters preceding the product's introduction. The second regression has as its dependent variable the second price change, and as explanatory variables exchange rate movements over the course of the second price spell as well as exchange rate movements in the preceding six quarters. If the first price is "stale" then the first price change should respond more to exchange rate movements preceding the first price spell, than the second price change does to exchange rate movements preceding the second price spell.