I. Introduction

A large literature has recently emerged that documents patterns of nominal price stickiness at the very micro level—the good level. The documented nominal durations are significantly shorter than the estimated real effects of money on output. The long-lasting real effects of monetary shocks can be reconciled with moderate price stickiness if real rigidities are an important phenomenon. Real rigidities are mechanisms that dampen price responses of firms because of factors such as strategic complementarities in price setting, real wage rigidity, and the dependence of costs on input prices that have yet to adjust, among other causes. An important empirical literature has emerged recently that evaluates the question, are quantitatively important real rigidities present in the data? The answer appears to depend on which data one looks at.

In international economics there is a large and growing literature that estimates exchange rate pass-through from exchange rate shocks into prices. The estimated exchange rate pass-through is found to be incomplete; that is, if the dollar depreciates by 10% relative to the euro, dollar prices of goods imported from the euro area increase by less than 10% even in the long run. This incompleteness in pass-through is argued to be consistent with the presence of important real rigidities. Exchange rate changes generate relative price movements for the same good across markets despite costs being the same. This destination-specific markup is argued to be consistent with the presence of significant strategic complementarities in price setting.

The closed economy literature, however, uses indirect tests of real rigidities in the absence of well-identified and sizable shocks such as exchange rate shocks. The recent work based on micro evidence for
retail prices argues that real rigidities are not an empirically important phenomenon.

There are many developments in the measurement of real rigidities in the closed and open economy literatures, but these developments have taken place in parallel. In this paper we bring together the closed economy macro literature that focuses mainly on indirect tests of real rigidities with the international pricing literature that uses an observable and sizable shock, namely, the exchange rate shock, to evaluate the behavior of prices, particularly strategic complementarities in pricing. We first review the recent evidence on real rigidities to evaluate if there is a consensus emerging on the importance of these rigidities in the data. Second, since the two literatures use different metrics to evaluate the importance of real rigidities, we use unpublished international price data collected by the U.S. Bureau of Labor Statistics (BLS) to estimate both metrics using the same data. Third, we present new evidence on the dynamic response of international prices to exchange rate shocks and the response to competitor prices. Fourth, we calibrate sticky-price macro models (Calvo and menu cost) with a retail sector and a wholesale sector to the evidence on the variable markup channel of real rigidities. We evaluate the ability of our model to match the behavior of prices in the data and to measure the extent of monetary nonneutrality this channel generates.

In reviewing the literature, we group evidence based on whether the prices studied refer to retail (consumer) prices or wholesale prices. Wholesale prices can alternatively be viewed as intermediate-good prices in business-to-business transactions. The literature on exchange rate pass-through into at-the-dock prices of goods refers to wholesale prices. A review of the existing literature reveals one surprisingly consistent result across several studies, surprising since these studies use different methodologies and data sets. This result is that strategic complementarities, for example, operating through variable markups, play a small role for retail prices but appear to be quite important for wholesale prices.

We next use the BLS import price data to perform tests of real rigidity, in which we use measures employed in the closed economy literature, namely, the persistence of reset-price inflation (Bils, Klenow, and Malin 2009), and measures employed in the open economy literature, namely, the dynamic response of prices to exchange rate shocks. The actual import-price inflation series has a monthly persistence of 0.56, whereas the corresponding reset-price inflation series has a persistence of −0.04. In comparison, Bils et al. (2009) estimate for retail prices that the inflation
series has persistence of −0.05 and the reset-price inflation series has a persistence of −0.41. In comparison to retail prices, import prices have greater persistence, but the magnitude of this persistence suggests that there is very little sluggishness in price adjustment.

However, when we project the aggregate import reset-price inflation on lags of the trade-weighted nominal exchange rate changes, we find that the autocorrelation of the fitted series is substantially higher as compared to that of unconditional reset-price inflation (0.33 vs. −0.04). We find similar evidence using micro-level price adjustments. We show that individual import prices, conditional on changing, respond to exchange rate shocks prior to the last time the price was adjusted and that these lagged effects are large and statistically significant. The pass-through conditional on a price change to the cumulative exchange rate change since the last price adjustment is 0.11, and the response to the cumulative exchange rate over the previous price duration is 0.08. Both these pieces of evidence that evaluate the response to a specific shock suggest a more important role for real rigidities as compared to the point estimate of the autocorrelation of reset prices.

Next, we evaluate the importance of strategic complementarities in price setting for incomplete pass-through, using some measures that capture the pricing behavior of competitors and measures that capture the extent of competition in sectors. These measures are not perfect, but they do provide useful information about pricing behavior. We use the prices set by other firms in the same 10-digit or four-digit harmonized code in the import price sample to control for the behavior of competitor prices, and we find that they have an important positive effect on firms’ pricing, reducing the direct pass-through of exchange rate into prices. The point estimates are consistent with a markup elasticity of 1.5, which implies a 40% pass-through for purely idiosyncratic shocks. We also evaluate the sensitivity of firm pricing to shocks to competitors by measuring the response of prices to movements in the U.S. trade-weighted exchange rate that is orthogonal to the bilateral exchange rate for the country. We find the response to be sizable and significant. In a similar vein, we find that exchange rate pass-through is higher in response to a more aggregate shock as compared to more idiosyncratic shocks when comparing the response to bilateral exchange rate shocks versus trade-weighted exchange rate shocks.

We also relate the incompleteness in pass-through to certain sectoral features that proxy for the level of competition among importers. An important distinction between retail prices and wholesale prices is that the latter capture business-to-business transactions. Consequently, the
strength of the bargaining power of the buyer can have an impact on the extent of pass-through. We use unpublished measures of concentration in the import sector provided to us by the BLS—specifically, the Herfindahl index and the number of importers that make up the top 50% of trade—to evaluate this hypothesis. While the point estimates in many cases suggest that sectors dominated by a few large importers have lower pass-through from foreign firms, the estimated standard errors are large.

Finally, we use estimates from the data to calibrate a closed economy model with different degrees of variable markup elasticity at the wholesale and retail level. In the existing monetary literature, there is typically no interesting distinction made between the retail and wholesale sectors. We calibrate the parameters for the wholesale sector using the evidence from international prices. In the benchmark model, we use Calvo price setting and later evaluate the case of menu cost pricing. First, we show that sluggishness in response to monetary shocks in wholesale prices feeds into slow adjustment of retail prices. However, the aggregate inflation and reset-price inflation series exhibit little persistence since their movement is dominated by more transitory shocks. However, conditional on monetary shocks or exchange rate–like shocks, the inflation series exhibit considerable persistence. Similarly, output series can exhibit significant monetary nonneutralities. Second, while calibrated real rigidities in the form of variable markups increase the size of the contract multiplier, their effects are limited unless they are coupled with exogenous sources of persistence. However, the model fails to match the slow dynamics in price adjustment that was documented in the empirical data, which suggests that additional sources of persistence are missing from the model.

Why does one observe differences in markup variability at the wholesale and retail levels? We do not provide a definitive answer here, but we conjecture that this can be consistent with differences in the competitive environment at the two levels. That is, the retail sector can be described as monopolistically competitive, whereas the wholesale sector is better described as a bargaining environment between a final-good producer and a limited number of intermediate-good suppliers. This can be formalized using a bargaining model of wholesale pricing; this can be found in the working paper version of this paper (Gopinath and Itskhoki 2010b). In this model, each final-good producer bargains with its intermediate-good suppliers regarding the price of intermediate goods. Given these bargained prices, the final-good producer is free to choose quantities of the intermediate inputs, as well as to set the price of its final good in the
monopolistically competitive consumer market. This model results in constant markups at the retail stage but in variable markups at the wholesale level that depend, among other things, on the relative bargaining power of the final-good producer and on the market share of the intermediate-good supplier.

Important outstanding questions are whether wholesale prices are allocative and whether contracts specify fixed prices at fixed quantities. While there is no simple way to test this, Gopinath and Rigobon (2008) show that, in the case of contracts for international prices, contracts typically involve a fixed price with a quantity range specified as opposed to a fixed quantity. Moreover, firms export the same good at the same price to multiple destinations, and consequently prices behave in many cases like list prices. Further, the behavior of prices is consistent with models of monopolistic price setting, where prices are allocative, as is discussed in the papers of Neiman (2009), Gopinath and Itskhoki (2010a), and Gopinath, Itskhoki, and Rigobon (2010). Also, as we detail later, changes in intermediate-good prices effect final-good prices as they are fully passed through into retail consumer prices. These separate pieces of evidence are consistent with wholesale prices being allocative.

This paper is structured as follows. Section II provides a descriptive framework that spells out the sources of real rigidities that can result in sluggish price adjustment. Section III reviews the closed and open economy literatures on real rigidities that use micro price data. Section IV presents new empirical results on price adjustment, using international data. Section V presents the closed economy model with differential markup variability in the retail and wholesale sectors and sluggish price adjustment. All derivations and technical details are relegated to the appendices.

II. Real Rigidities: A Descriptive Framework

In this section, we set up some notation and spell out the sources of real rigidities that can result in sluggish price adjustment. This simple descriptive framework is used to organize the discussion of empirical evidence in the following two sections.

Define the desired price of a firm as the price it would set if it could adjust its prices flexibly in a given economic environment. With sticky prices, a forward-looking firm sets its price as a weighted average of future desired prices. The presence of real rigidities slows down the response of the desired price to a shock. Real rigidities, powered by nominal price stickiness and staggered price adjustment (Taylor 1980; Calvo
can be either at the aggregate/industry level or at the micro/firm level. Real rigidities at the aggregate level include round-about production structure, as in Basu (1995); real wage rigidity, as in Blanchard and Gali (2007); and segmented input markets, as in Woodford (2003). Real rigidities at the firm level, or strategic complementarities in pricing, arise either from nonconstant marginal cost (i.e., decreasing returns to scale), as in Burstein and Hellwig (2007), or from variable markups. In turn, variable markups can be due to nonconstant elasticity of substitution (non-CES) demand, as in Kimball (1995) and Klenow and Willis (2006), or to strategic complementarities in price setting between large firms, as in Atkeson and Burstein (2008).

Empirical studies attempt to determine which of these channels of real rigidities, if any, are present in the data, as well as the plausible magnitudes of their effects. Since the empirical literature has examined evidence at both the retail and wholesale levels, we will maintain this distinction in our descriptive framework.

Retail (final-good) pricing. The log desired price for a final good $i$ at time $t$ can be written as a log desired markup over the marginal cost of the firm:

$$\ln p_{it} = \mu_{it} + \alpha s_t + \frac{1}{C_0} \ln w_t + z_{it},$$

where $\mu_{it}$ is the log desired markup, $s_t$ is the log price of the intermediate input, $w_t$ is the log price of other inputs (e.g., labor), and $z_{it}$ represents the firm-specific marginal cost or markup shock, which may include an aggregate component common across a subset of firms; $\alpha \in [0, 1]$ is the share of intermediate inputs in the production cost of the final-good firms.

Throughout the paper we abstract from the nonconstant marginal cost channel of strategic complementarities in favor of the variable markup channel. For most purposes, the two mechanisms are largely substitutable; however, variable markups can additionally explain pricing to market, a phenomenon with strong empirical support, as discussed in the next section. We capture variable markups in the following reduced-form way:

$$\mu_{it} \approx \bar{\mu} - \Gamma_R (p_{it} - p_t),$$

where $p_t$ is the industry price index and $\Gamma_R \geq 0$ is the elasticity of the firm’s markup with respect to its relative price. That is, a higher relative
price of the firm reduces its desired markup. As we discuss in more detail in Section V, this simple description of the desired markup is consistent with various models of variable markups cited above.

Combining the above two equations, we arrive at

\[
\tilde{p}_{it} = \frac{1}{1 + \Gamma_R} \left[ \tilde{\mu}^R + \alpha s_t + (1 - \alpha) w_t - z_{it} \right] + \frac{\Gamma_R}{1 + \Gamma_R} p_t.
\]  

(1)

When \( \Gamma_R > 0 \), the desired price of the firm increases in the prices set by the competitors of the firm, which we refer to as strategic complementarities in price setting. Under the same circumstances, a response to an idiosyncratic marginal cost/markup shock (such as \( z_{it} \)) is incomplete: the pass-through of an idiosyncratic shock into the desired price equals \( 1/(1 + \Gamma_R) \in [0, 1] \).

To summarize, sluggishness in the response of desired final-good prices may arise either due to staggered price adjustment when \( \Gamma_R > 0 \) or due to sluggish adjustment in the marginal cost, that is, due to a slow response of \( s_t \) and/or \( w_t \) to shocks.

Wholesale (intermediate-good) pricing. Let the log desired price of an intermediate variety \( j \) be

\[
\tilde{s}_{jt} = \mu_{jt} + w_t + \phi_j e_t - a_{jt},
\]

where \( \mu_{jt} \) is the log desired markup, \( w_t \) is the price of inputs (e.g., labor), and \( a_{jt} \) is a firm-specific marginal cost or markup shock, which again may contain an aggregate component. Furthermore, we allow the cost of intermediate firms to respond directly to exchange rate fluctuations \( e_t \) with various elasticities \( \phi_j \). For example, some of the intermediate varieties may be produced abroad, so that their marginal cost in local currency fluctuates together with the exchange rate.\(^3\) Finally, the log price of the intermediate good, \( s_t \), is an aggregate of the prices of intermediate varieties, \( s_{jt} \), which can be approximated by a geometric average: \( s_t \approx \int_j s_{jt} d\tilde{j} \).

We similarly introduce the possibility of variable markups for the intermediate varieties:

\[
\mu_{jt} \approx \tilde{\mu} - \Gamma (s_{jt} - s_t),
\]

where \( \Gamma \) is again the price elasticity of markup that may vary across varieties \( j \). We can rewrite the desired price of the intermediate
variety as

$$\tilde{s}_t = \frac{1}{1 + \Gamma} \left( \tilde{\mu} + w_t + \phi_s e_t - a_t \right) + \frac{\Gamma}{1 + \Gamma} s_t. \quad (2)$$

Purely idiosyncratic shocks wash out in the aggregate and do not affect $s_t$. Therefore, they get passed through into desired prices immediately with a pass-through coefficient of $1/(1 + \Gamma)$. More aggregate shocks have two channels through which they affect the desired price—directly ($w_t$ or $e_t$) and via the prices of competitors ($s_t$). Coupled with nominal price stickiness and staggered price adjustment, variable mark-ups ($\Gamma > 0$) may generate sluggish adjustment that lasts past the periods of nominal stickiness.

**Aggregate real rigidities.** Various aggregate sources of real rigidities enter through a sluggish adjustment to shocks of the marginal cost component, denoted by $w_t$. Specifically, real wage rigidities, Basu (1995) round-about production, and segmented input markets can all be captured by a slow adjustment of $w_t$ to shocks. For concreteness, one can use the following structure to think about aggregate real rigidities (see Sec. V):

$$w_t = \gamma m_t + (1 - \gamma) p_t, \quad (3)$$

where $m_t$ is aggregate nominal spending, $p_t$ is the consumer price level, and $\gamma > 0$ is the elasticity of the cost $w_t$ with respect to monetary (nominal spending) shocks. The smaller is $\gamma$, the more sluggish is the response of aggregate costs to monetary shocks, that is, the stronger are aggregate real rigidities.

Equations (1)–(3) describe our simple framework to account for real rigidities.¹ Sluggish adjustment in the desired final-good prices $\tilde{p}_{it}$ can arise due to one of the three channels: (a) sluggish response of aggregate costs $w_t$ (small $\gamma$), (b) sluggish adjustment of intermediate prices $s_t$ ($\Gamma > 0$ provided $\alpha > 0$), and (c) gradual adjustment in final-good desired prices ($\Gamma_R > 0$). We now evaluate the evidence on these different sources of real rigidities.

### III. Evidence on Real Rigidities: A Review

Although appealing at the intuitive level, real rigidities are hard to identify and measure in the data. Aggregate real rigidities imply a sluggish

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¹ Basu (1995) round-about production, and segmented input markets can all be captured by a slow adjustment of $w_t$ to shocks. For concreteness, one can use the following structure to think about aggregate real rigidities (see Sec. V):
response of the marginal cost of firms to aggregate shocks. Firm-level real rigidities imply a muted response of the firms’ prices conditional on price adjustment to the marginal cost shocks. Since data on marginal costs are usually unavailable, it is hard to test these mechanisms directly. Therefore, the literature relies either on calibrations or indirect empirical tests.

In parallel, the international literature has a long tradition of estimating exchange rate pass-through, namely, the response of international prices to changes in exchange rates. A standard assumption in the empirical pass-through literature is that changes in the exchange rate represent a shock to the firm’s costs that is orthogonal to other shocks that affect the firm’s pricing decision and, in reverse, they are not affected by firm pricing. This assumption is motivated by the empirical finding that exchange rate movements are disconnected from most macro variables at the frequencies studied in the literature (1–2-year horizons). While this assumption might be more problematic for commodities such as oil or metals and for some commodity-exporting countries such as Canada, it is far less restrictive for most differentiated goods and most developed countries. Alternatively, the exchange rate change can be viewed not as a pure exogenous cost shock but rather as an observable signal about the underlying fundamental macro shocks that differentially affect the costs of domestic and foreign firms. In this case, the coefficients on the exchange rate have less of a structural interpretation, but they can still provide useful information about the nature of firm pricing.

The estimated exchange rate pass-through has typically been found to be incomplete. This incompleteness in pass-through can be consistent with various sources of firm-level and aggregate real rigidities, as well as with rational inattention and sticky information. The fact that firms export the same good to multiple destinations provides a mechanism to distinguish the source of the real rigidity. The pervasive evidence on pricing to market, the practice of charging different prices for the same good in different locations, provides support for the variable markup channel of incomplete pass-through.

In Sections III.A and III.B, we review the empirical evidence on real rigidities in the closed and open economy literatures, respectively. We restrict attention to the most recent literature that uses micro-level price data.

A. Closed Economy Literature

In the closed economy literature the evidence on real rigidities typically arises from indirect identification strategies. Klenow and Willis (2006)
evaluate the variable markup channel of real rigidities by calibrating a menu cost model with Kimball (1995) demand to match the BLS micro retail price evidence on the frequency and size of price adjustment. They conclude that the levels of real rigidity sufficient to generate significant monetary nonneutrality have implausible implications for the required size of menu costs and idiosyncratic productivity shocks. Intuitively, significant real rigidities compress price dispersion so that much larger idiosyncratic shocks are required to match the size of price adjustment. At the same time, large menu costs are required to match the average durations of price rigidity given large idiosyncratic shocks.

Burstein and Hellwig (2007) replace Kimball demand with increasing marginal costs at the firm level to generate strategic complementarity in price setting. They calibrate the extent of decreasing returns to scale (elasticity of the marginal cost with respect to output) in order to match scanner data from a large chain of supermarkets in the Chicago area (Dominick’s) on the comovement between prices and market shares. This calibration implies a moderate role for strategic complementarities generated via curvature in marginal costs.

Bils et al. (2009) develop a test to assess a broad class of models with real rigidities. They examine the persistence of the actual inflation and reset-price inflation series. Reset-price inflation is defined as an average reset price change within the group of goods that adjusts prices in a given period, while reset prices for all other goods are indexed by reset-price inflation (for a formal definition, see Sec. IV.A). If real rigidities are important, there should be significant persistence both in the actual inflation and the reset-price inflation series. The paper finds low persistence for actual price inflation and negative persistence for reset-price inflation (in both cases persistence is measured as the monthly autoregressive coefficient), suggesting that the selection mechanism present in menu costs models (emphasized by Caplin and Spulber [1987]) offsets the effects of real rigidities and that the real effects of money last less time than nominal price durations.

In a different paper, Klenow and Willis (2007) find that micro price changes respond to old macro information, known prior to their previous price adjustment. They interpret this as evidence of sticky information, in the spirit of Mankiw and Reis (2002); nevertheless, this evidence is also consistent with the presence of real rigidities and pricing complementarities that lead to incomplete price adjustment at the micro level.

As previously mentioned, these tests for real rigidities are indirect, relying mainly on calibrations or statistical properties of observable prices, given that marginal costs and markups are not directly
observable. The exception to this is a paper by Eichenbaum, Jaimovich, and Rebelo (2011), which uses scanner data for a large grocery store chain in the United States. Eichenbaum et al. observe both the wholesale price at which the store obtains the good (specific universal product code [UPC]) and the retail price at which the store sells it. At very short horizons the wholesale price can be viewed as the relevant marginal cost for the firm. They find that, conditional on changing reference prices, there is no evidence of variable markups; that is, all of the reference wholesale cost change is passed through into reference retail prices.7

B. Open Economy Literature

The open economy literature evaluates the response of retail prices and at-the-dock prices to exchange rate shocks. While pass-through is less than one in both cases, pass-through into retail prices is always much lower than into at-the-dock prices. This is not surprising given that distribution costs, which are mainly nontraded costs unaffected by the exchange rate, are an important component of retail prices (e.g., see Goldberg and Campa 2010). Goldberg and Knetter (1997) summarize a large body of the earlier empirical literature on pass-through.

A virtue of international price data is that one observes the prices at which the same firm sells its product in different destinations. There is considerable evidence of pricing to market; that is, firms sell the same product at different prices in different destinations. This concept, first proposed in Dornbusch (1987) and Krugman (1987), attributes an important role to strategic complementarities that generate variable markups. In a recent paper, Fitzgerald and Haller (2008) provide the most direct evidence of this phenomenon. They examine the pricing of Irish manufactures in domestic and export (UK) markets. They find a pronounced price differential response to exchange rate movements. Since the goods are manufactured in the same plant, the difference is attributable to variation in markups.8 Similarly, Burstein and Jaimovich (2008) use supermarket scanner data for stores in the United States and Canada and find that products with the same UPC produced in the same country sell at different wholesale prices in the United States and Canada. More specifically, relative wholesale prices in the two markets move closely with the exchange rate.

In a recent set of papers, Gopinath and Rigobon (2008), Gopinath and Itskhoki (2010a), and Gopinath et al. (2010) use micro import price data collected by the BLS for the period 1994–2005 to provide evidence of incomplete exchange rate pass-through at the dock even conditional
on prices being changed. While incomplete exchange rate pass-through per se is consistent with evidence on variable markups, it can arise from other sources of real rigidities, as previously discussed. Gopinath and Itskhoki (2010a) document a positive correlation between the frequency of price adjustment for a good and exchange rate pass-through conditional on price adjustment. They argue that this positive correlation is consistent with variation in markup elasticity across firms and cannot be consistent with other sources of firm heterogeneity. Neiman (2010) uses the same BLS import price data to document that, consistent with the greater importance of the strategic complementarity channel for arm’s-length transactions as compared to intrafirm transactions, for differentiated products intrafirm prices are characterized by more stickiness, less synchronization, and greater exchange rate pass-through.

Finally, there are papers that evaluate the response of retail prices and wholesale prices for the same good to an exchange rate shock. Gopinath et al. (2009) use scanner data on retail and wholesale prices for stores of the same supermarket chain in the United States and Canada and find that, in response to an exchange rate shock, movements in relative retail prices in a common currency are explained mainly by movements in relative costs and very little by relative markups. Goldberg and Hellerstein (2006) and Nakamura and Zerom (2010) find that, for beer and coffee sales in a supermarket store, the pass-through from exchange rate shocks to wholesale costs is incomplete but the pass-through from wholesale costs to retail prices is close to complete. In a recent study, Berger et al. (2009) match goods in the BLS import price index to those in the BLS consumer price index (CPI) for the period 1994–2007 and find that the overall distribution wedge, which is the percentage difference between retail and at-the-dock prices, does not vary systematically with the exchange rate, which implies a nearly complete pass-through from at-the-dock to retail prices.9

Summary. Overall, a consistent finding across studies in the closed and open economy literatures is that the variable markup channel of real rigidities is an important feature of the wholesale cost data but not of the retail price data. This is surprising consensus given the different approaches used in the closed and open economy literatures and the different data sets involved. In the terminology of Section II, the empirical evidence is consistent with $\Gamma_R \approx 0$ and $\Gamma > 0$. Given the feedback from wholesale prices $s_t$ to retail prices $p_t$ in equation (1), one could expect sluggish adjustment in wholesale reset prices to generate slow adjustment in retail reset prices. The fact that Bils et al. (2009) find a negative persistence for overall reset-price inflation could then be consistent with
either a small $\alpha$ or a large transitory variance for $z_{it}$, as long as these shocks are not purely idiosyncratic. In the next section we explore explicitly, using international price data, the dynamics of price adjustment both unconditionally and conditioning on exchange rate shocks.

IV. New Evidence on Real Rigidities

As discussed in Section III.A, Bils et al. (2009) find that regular inflation and reset-price inflation for retail price data are consistent with the absence of important real rigidities. In this section, we present evidence on the properties of regular inflation and reset-price inflation for import price data using the Bils et al. (2009) micro data on international prices. We then compare it to the properties of the inflation series, conditional on exchange rate shocks, which allows us to evaluate the conditional response of prices to a given aggregate cost shock. We then study the dynamic properties of price adjustment at the firm level.

First, we show that reset-price inflation for the import price data is less negatively autocorrelated as compared to that documented for retail prices. This is consistent with the conclusion that international prices depict higher real rigidities as compared to retail prices. Second, we present evidence regarding the sluggish response of price changes to exchange rate shocks. We do this in two ways. First, we project aggregate regular and reset-price inflation on lags of the exchange rate changes. We find that the autocorrelation of the fitted series is substantially higher as compared to that of the unconditional series. Second, using the micro data, we show that individual prices, conditional on changing, respond to exchange rate shocks that were realized prior to the previous price adjustment. Both pieces of evidence suggest a more important role for real rigidities as compared to the conclusions one can draw from the analysis of unconditional aggregate inflation series.

While this evidence is supportive of the presence of sizable real rigidities, it does not discriminate among various sources of real rigidities. Therefore, we next try to assess specifically the importance of strategic complementarities in price setting. We do so by studying the response of the firm’s prices to the shocks to its competitors. Our central result here is that the firm’s prices respond strongly to the prices of its competitors and that this channel explains a significant fraction of exchange rate pass-through into prices.

The data used in this section are the micro import price data underlying the construction of the U.S. import price index. This covers the period 1994–2005. For details regarding the data, see Gopinath and Rigobon (2008).
A. Reset-Price Inflation

We follow Bils et al. (2009) in estimating the reset-price inflation series for U.S. imports. The log price of a good $i$ at time $t$ is denoted by $p_{it}$. The log of the reset price at time $t$ for good $i$ is denoted by $p_{it}^*$ and defined as

$$p_{it}^* = \begin{cases} \ p_{it} & \text{if } p_{it} \neq p_{i,t-1} \\ p_{i,t-1} + \pi_{i}^{*} & \text{if } p_{it} = p_{i,t-1} \end{cases}$$

where $\pi_{i}^{*}$ is the average reset-price inflation of those goods whose prices change at time $t$. The inflation of the actual price series is referred to as regular-price inflation. In a Calvo pricing environment, where firms only differ in the exogenous frequency of price adjustment, the behavior of reset-price inflation will capture the extent of real rigidities. As discussed in detail in Bils et al. (2009), in the presence of real rigidities reset prices will adjust sluggishly, as firms have multiple price adjustments before they fully respond to a shock. The regular-price inflation series displays even greater persistence, as each firm waits for the random arrival of the opportunity to change its price. In the case of state-dependent pricing or variations in desired price responses across sectors, the behavior of measured reset prices is not as direct because there is selection of which firms change prices. Gopinath and Itskikhoki (2010a) show that sectors that have a higher frequency of price adjustment also have higher long-run pass-through (less real rigidities). Consequently, the reset-price series is affected by those goods that change prices more frequently and pass-through eventually a lot more.

It is important to point out that the sample sizes in the import price data are not as large as in the CPI data and that given the low frequency of price adjustment the number of actual price changes used to impute the reset price series is low. This would suggest that the import reset-price inflation series is more subject to noise as compared to the construct of Bils et al. (2009) for the U.S. CPI.

The persistence and volatility of the regular and reset-price inflation series within various subsamples are reported in table 1. We measure persistence as the AR(1) coefficient for the series, and we report the standard deviation of the series as a measure of volatility. The first row of table 1 reports these moments for consumer-price inflation from Bils et al. (2009). The rest of the rows provide these statistics for the BLS import-price data.

Columns 1 and 2 of table 1 report statistics for the unconditional inflation series. For all subsamples considered, import-price inflation is
more persistent relative to consumer-price inflation as calculated by Bils et al. (2009).\(^{15}\) For example, for the dollar-priced imports, the persistence of inflation is 0.56, whereas for consumer prices it is close to zero (−0.05). The import-price inflation is also more volatile. A similar comparison holds for reset-price inflation. While Bils et al. find that reset-price inflation for consumer prices is negatively autocorrelated (−0.41), we find essentially zero autocorrelation for import prices (e.g., −0.04 for the dollar-priced imports). This difference in persistence of consumer-price and import-price inflation is consistent with the different nature of pricing at the consumer and intermediate-good levels (since most imports are intermediate goods).\(^{16}\)

An important feature of the international data is that we can examine the response of the inflation series to a specific shock, namely, the exchange rate shock. This has advantages over just looking at reset-price inflation that aggregates (imperfectly in small samples) across idiosyncratic, sectoral, and aggregate shocks. We accordingly project the regular and reset-price inflation series on current and 24 lags of the log changes of the U.S. trade-weighted exchange rate and compute the moments for the projected series.

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<tr>
<th>Table 1</th>
<th>Volatility and Persistence of Regular and Reset-Price Inflation</th>
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<tr>
<td></td>
<td>Unconditional</td>
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<td></td>
<td>AR(1)</td>
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<tr>
<td>Regular-price inflation:</td>
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<tr>
<td>Consumer prices</td>
<td>−.05</td>
</tr>
<tr>
<td>Import prices</td>
<td>.51</td>
</tr>
<tr>
<td>Dollar-priced goods</td>
<td>.56</td>
</tr>
<tr>
<td>Market transactions</td>
<td>.43</td>
</tr>
<tr>
<td>Reset-price inflation:</td>
<td></td>
</tr>
<tr>
<td>Consumer prices</td>
<td>−.41</td>
</tr>
<tr>
<td>Import prices</td>
<td>.02</td>
</tr>
<tr>
<td>Dollar-priced goods</td>
<td>−.04</td>
</tr>
<tr>
<td>Market transactions</td>
<td>−.03</td>
</tr>
</tbody>
</table>

Source: Statistics for consumer prices are from Bils, Klenow, and Malin (2009).
Note: Import prices exclude petrol classifications; the rows for market transactions include only dollar-prices goods. The last two columns project the inflation series on the current and 24 lags of the log changes of the U.S. trade-weighted exchange rate and compute the moments for the projected series.
For instance, in the case of dollar-priced imports, the conditional regular-price inflation series has an overall persistence of 0.79, while its unconditional persistence is 0.56. Similarly, for the reset-price inflation series the conditional persistence is 0.33, as opposed to the unconditional persistence of −0.04.

This evidence is consistent with the presence of multiple shocks of different degrees of persistence driving the inflation process. Under these circumstances, the unconditional persistence of the inflation series might not accurately reflect the underlying sluggishness in the micro-level price adjustment. In the next subsection, we present further evidence of the sluggish response of prices to exchange rate shocks by examining the behavior of individual prices and their response to lagged exchange rates changes.

B. Micro Dynamics of Price Adjustment

At the good level, we estimate pass-through into prices of exchange rate shocks realized during the most recent period of price nonadjustment and of those that were realized prior to the previous price adjustment. In the absence of real rigidities, all adjustment should take place at the first instance of price change, and hence the coefficient on the exchange rate change prior to the previous price adjustment should be zero.

Formally, we estimate the following regression:

\[
\Delta \tilde{p}_{i,t} = \beta_1 \Delta \tau_{1,t} + \beta_2 \Delta \tau_{2,t} + Z_{i,t}^\prime \gamma + \epsilon_{it}, \tag{5}
\]

where \( i \) indexes the good and \( \Delta \tilde{p}_{i,t} \) is the change in the log dollar price of the good, conditional on price adjustment in the currency of pricing;\(^{17} \) \( \Delta \tau_{1,t} \) is the cumulative change in the log of the bilateral nominal exchange rate over the duration when the previous price was in effect (which we denote \( \tau_1 \)). Similarly, \( \tau_2 \) denotes the duration of the previous price of the firm, so that \( \Delta \tau_{2,t} \) is the cumulative exchange rate change over the previous (the one prior to the previous price change) period of nonadjustment. Figure 1 illustrates a hypothetical price series: if \( \Delta \tilde{p}_{i,t} \) is the price change between \( t_3 \) and \( t_{LL} \), \( \Delta \tau_{1,t} \) is then the exchange rate change between \( t_3 \) and \( t_{LL} \) and \( \Delta \tau_{2,t} \) is the exchange rate change between \( t_2 \) and \( t_3 \). Finally, \( Z_{i,t} \) includes controls for the cumulative change in the foreign consumer price level, the U.S. consumer price level and fixed effects for every BLS-defined primary strata (mostly two- to four-digit harmonized codes) and country.
pair. We allow $Z_t$ to include lagged foreign and domestic inflation. The standard errors are clustered at the level of the fixed effects. We restrict the sample to nonpetrol, dollar-priced goods and market transactions. Note that this specification requires the goods to have at least two price adjustments during their life. Since there are several goods that have only one price change during their life, we lose about 30% of the goods.

By conditioning on a price change we get past the period of nominal rigidity, which is essential to understanding the role of real rigidities. In a Calvo pricing environment, since the decision to change prices is exogenous, there are no selection issues to be concerned with. However, in an environment with endogenous frequency of price adjustment, conditioning on a price change will induce a bias in the exchange rate pass-through estimates, as it generates a conditional correlation between the exchange rate and the residual even if the unconditional correlation is zero. This is problematic if one tries to provide a structural interpretation to the coefficient. This is not our purpose here. We use this specification to provide a relation in the data between the response of prices conditional on adjusting to lagged exchange rate changes. Later we will estimate these regressions in the model-simulated data, where frequency is chosen endogenously, and infer how well models with real rigidities perform in matching the facts in the data. This exercise is accordingly similar to that of trying to match the behavior of the reset-price inflation series, a series

Fig. 1. Hypothetical good-level price series and nominal exchange rate
that also is affected by selection issues. Furthermore, when the selection bias is strong, the pass-through coefficient on first adjustment ($\beta_1$) is biased upward while the pass-through coefficient on second adjustment ($\beta_2$) is biased downward, which makes it harder to identify the presence of real rigidities.

Table 2 reports the results from estimation of regression (5). We provide evidence for various subsamples of the data. Across all specifications we find that exchange rate shocks that took place prior to the current period of nonadjustment have a significant effect on current price adjustments. This is consistent with the existence of real rigidities in pricing. The strength of these lagged effects is much stronger than what would be suggested purely by the reset-price inflation series. The first row of table 2 points out that the elasticity of current price changes to lagged exchange rate shocks for dollar-priced goods is 0.08, which is only slightly smaller than the response to the contemporaneous exchange rate movement (equal to 0.11). The importance of these lagged effects is consistently present in all subsamples. For the

<table>
<thead>
<tr>
<th></th>
<th>$\beta_1$</th>
<th>$\beta_2$</th>
<th>$N$</th>
<th>$R^2$</th>
</tr>
</thead>
<tbody>
<tr>
<td>All countries</td>
<td>.11</td>
<td>.08</td>
<td>69,917</td>
<td>.01</td>
</tr>
<tr>
<td>(0.02)</td>
<td>(0.01)</td>
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<td></td>
<td></td>
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<tr>
<td>Non-OECD countries</td>
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<td>.04</td>
<td>37,108</td>
<td>.01</td>
</tr>
<tr>
<td>(0.02)</td>
<td>(0.01)</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>High-income OECD countries</td>
<td>.23</td>
<td>.18</td>
<td>32,809</td>
<td>.02</td>
</tr>
<tr>
<td>(0.02)</td>
<td>(0.02)</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Euro area</td>
<td>.22</td>
<td>.14</td>
<td>5,933</td>
<td>.02</td>
</tr>
<tr>
<td>(0.04)</td>
<td>(0.03)</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Japan</td>
<td>.26</td>
<td>.24</td>
<td>4,249</td>
<td>.06</td>
</tr>
<tr>
<td>(0.04)</td>
<td>(0.03)</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Canada</td>
<td>.28</td>
<td>.34</td>
<td>14,620</td>
<td>.01</td>
</tr>
<tr>
<td>(0.16)</td>
<td>(0.05)</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Differentiated goods</td>
<td>.14</td>
<td>.10</td>
<td>21,360</td>
<td>.02</td>
</tr>
<tr>
<td>(0.02)</td>
<td>(0.02)</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>No missing prices:</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>All countries</td>
<td>.10</td>
<td>.08</td>
<td>45,765</td>
<td>.01</td>
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<tr>
<td>(0.02)</td>
<td>(0.01)</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>High-income OECD countries</td>
<td>.19</td>
<td>.13</td>
<td>22,436</td>
<td>.01</td>
</tr>
<tr>
<td>(0.03)</td>
<td>(0.02)</td>
<td></td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

Note: Items $\beta_1$ and $\beta_2$ are the pass-through coefficients at the first and second rounds of price adjustment, respectively; these are estimated from regression (5). Standard errors in parentheses are clustered at the country × four-digit sector level. $N$ is the number of price changes in the sample. Results under “No missing prices” exclude from the sample all price changes that were followed or preceded by a missing price.
high-income OECD sample, the contemporaneous and lagged responses are 0.23 and 0.18, respectively. For the non-OECD sample, the pass-through rates are overall lower, but there are still important lagged effects. This is similarly documented for the euro-area countries, Japan, and Canada.

In the data there can be spells that have missing prices and where the new price follows or precedes a missing price. In this case the exact timing of the price change is not known, so lagged effects can arise from getting the timing wrong. The last two rows of table 2 check for the robustness of the results by including in the sample only those price changes that were not followed or preceded by a missing price. This changes the sample composition, but lagged responses are still strongly evident.

The results in this section are consistent with the evidence in Gopinath et al. (2010) that long-run pass-through is much higher than pass-through conditional on the first adjustment to the exchange rate shock. Here we present explicitly the dynamics and extend the sample to more countries.

We also divide goods into four equal-sized bins based on their frequency of price adjustment and estimate equation (5) within each bin separately. The purpose of this exercise is to evaluate if the importance of lags varies across goods with different frequencies of price adjustment. One conjecture may be that it is only the very high frequency goods that have multiple price adjustments to respond to a shock. In fact this is not the case, as we find that lags are important even for goods that adjust prices very infrequently: for example, the first quartile contains goods that adjust prices less than once a year, and in the first round the pass-through is 0.12, whereas it is 0.08 in the second round. This finding further assuages concerns about measurement issues with the timing of price adjustment.18

We also break the sample down by the end use of the product. Again, we find the second rounds of price adjustment to be significant for most end use categories. In the case of “Food, feed, and beverages” exported from the non-OECD countries (which dominate the sample of all countries in panel A of table 3), the pass-through is generally very low and insignificant from zero. In the case of “Consumer goods (nonfood and excluding automotive),” the dynamics are less evident. However, one should be careful about interpreting the results for this subsample because these goods more often have a fixed price during their life and then get discontinued and replaced. Since we do not observe price changes across discontinuations, we might be excluding important adjustments that take place at the time of product replacement.
Overall, the micro evidence is consistent with the aggregate level evidence of sluggish adjustment to exchange rate changes. This sluggishness is consistent with many forms of real rigidities, including variable markups, the Basu (1995) intermediate-input channel wherein each firm’s output is used as an input in production, and sluggish response of other factor costs (like wages) to the underlying source of exchange rate shocks. Equivalently, it could arise from rational inattention or sticky information. The next subsection evaluates how the extent of product market competition affects the patterns of pass-through in order to identify the effects of strategic complementarities in price setting.

### C. Competition and Pass-Through

In this section we evaluate the importance of the strategic complementarity in price setting for incomplete pass-through, using some measures that capture the pricing behavior of competitors and measures that capture the extent of competition within sectors. Ideally, one would need to perform this analysis with detailed industry data for each product and information on prices and market shares of different firms. These data, however, do not exist for the large number of products included in our study. Consequently, we use some proxies here, and they are necessarily

<table>
<thead>
<tr>
<th>Table 3</th>
<th>Dynamic Response to Exchange Rate Shocks, by End-Use Sectors</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>$\beta_1$</td>
</tr>
<tr>
<td>A. All countries:</td>
<td></td>
</tr>
<tr>
<td>Food, feed, and beverages</td>
<td>.05 (.04)</td>
</tr>
<tr>
<td>Industrial supplies and materials</td>
<td>.13 (.03)</td>
</tr>
<tr>
<td>Capital goods, except automotive</td>
<td>.16 (.03)</td>
</tr>
<tr>
<td>Consumer goods (nonfood)</td>
<td>.05 (.04)</td>
</tr>
<tr>
<td>B. High-income OECD countries:</td>
<td></td>
</tr>
<tr>
<td>Food, feed, and beverages</td>
<td>.16 (.04)</td>
</tr>
<tr>
<td>Industrial supplies and materials</td>
<td>.21 (.05)</td>
</tr>
<tr>
<td>Capital goods, except automotive</td>
<td>.21 (.04)</td>
</tr>
<tr>
<td>Consumer goods (nonfood)</td>
<td>.17 (.06)</td>
</tr>
</tbody>
</table>

Note: See the note to table 2. Clustered standard errors are in parentheses.
imperfect. Nevertheless, we find evidence that is consistent with the presence of significant strategic complementarities at the firm level.

Trade-weighted versus bilateral exchange rate. First, we evaluate the response of each firm’s pricing to its own bilateral exchange rate as compared to its response to the trade-weighted exchange rate. Movements in the trade-weighted exchange rate can be viewed as a more aggregate shock that affects a larger fraction of a firm’s competitors, as compared to a shock that only affects the bilateral exchange rate. An alternative interpretation could be that prices of a firm that uses production inputs from the rest of the world are sensitive to the movements in the trade-weighted exchange rate because it affects the firm’s costs.

More specifically, we estimate the following standard pass-through regression:

$$\Delta p_{k,t} = \alpha_i + \sum_{j=0}^{n} \beta_j \Delta e_{k,t-j} + \sum_{j=0}^{n} \gamma_j \pi_{k,t-j} + \epsilon_{k,t},$$

where $k$ indexes the country, $\Delta p$ is the average monthly log price change in dollars, $\pi$ is the monthly foreign-country inflation using the consumer price index, $n$ is the number of monthly lags, which varies from 1 to 24; $\Delta e_{k,t-j}$ is either a bilateral nominal exchange rate or the U.S. trade-weighted nominal exchange rate. Figure 2 plots $\sum_{j=0}^{n} \beta_j$ as a function of $n$ for each case, where we estimate a pooled regression restricting the coefficients $\beta_j$ to be the same across countries. Panel a of the figure plots the results for the all-countries sample, while panel b does it for the high-income OECD subsample. In both figures it is evident that the pass-through from the trade-weighted exchange rate exceeds the bilateral exchange rate pass-through, which is consistent with the hypothesis that firm’s prices are responsive to cost shocks of firm’s competitors. In further analysis, we find that this pattern is evident for countries in the euro area as well as the non-OECD countries, whereas it is less evident for Japan, Canada, and the United Kingdom.

We also perform the analysis using the individual price data, conditional on a price change. We evaluate the response to the bilateral exchange rate change since the last time the price was adjusted and to movements in the U.S. trade-weighted exchange rate that is orthogonal to the bilateral exchange rate for the country. More specifically, we run a first-stage regression, where we regress the trade-weighted exchange rate on the bilateral exchange rate. We calculate the residual and then estimate a second-stage regression, where we regress the price change, conditional on adjustment,
on the cumulative change in the bilateral exchange rate and in the residual.\textsuperscript{19} We include a control for the cumulative change in foreign country CPI inflation since the last price change. The results are reported in table 4. Consistent with the evidence in figure 2 using aggregate price changes, the effect of the residual is almost as large as the direct effect of the bilateral exchange rate, and in some cases it is even larger.

We should note that this evidence admits a number of interpretations. First, the sensitivity of prices to the trade-weighted exchange rate over and above the bilateral exchange rate is consistent with strategic complementarities in price setting when the competitors of the firm are exporters from other countries. However, it is also consistent with the fact that firms use production inputs imported from the rest of the world,

---

Fig. 2. Impulse responses to bilateral and U.S. trade-weighted exchange rate
and a movement in the trade-weighted exchange rate could therefore represent a relevant cost shock for the firm. In what follows we test for strategic complementarities in price setting more directly.

**Competitor prices.** A more direct test of the presence of strategic complementarities is to evaluate whether changes in the competitor prices affect the pricing decisions of the firm. We do so by estimating the following regression:

$$\Delta p_{i,k,t} = \beta_e \Delta r_{1e,i,k,t} + \beta_1 \Delta r_{1P,i,k,t} + \gamma Z_{i,t} + \epsilon_{i,t},$$

(7)

where $\Delta p_{i,k,t}$ is the change in the log dollar price of good $i$ in sector $k$, conditional on price adjustment, and $\Delta r_{1e,i,k,t}$ is the cumulative change in the log of the bilateral nominal exchange rate over the duration for which the previous price was in effect. Now $\Delta r_{1P,i,k,t}$ is a measure of the cumulative price change by firms other than firm $i$ in sector $k$.\textsuperscript{20} We also estimate the same regression for the life-long change in the price of the good and refer to the coefficients in this case as long-run pass-through. In terms of figure 1, this corresponds to having $p_{i,t} - p_{i,0}$ on the left-hand side of (7) and corresponding cumulative changes in variables on the right-hand side.\textsuperscript{21} Finally, $Z_{i,t}$ represents the cumulative change in the

---

**Table 4**

Pass-Through of Bilateral and U.S. Trade-Weighted Exchange Rate

<table>
<thead>
<tr>
<th></th>
<th>Bilateral Exchange Rate</th>
<th>Trade-Weighted Exchange Rate</th>
<th>N</th>
<th>$R^2$</th>
</tr>
</thead>
<tbody>
<tr>
<td>All countries</td>
<td>.11 (.01)</td>
<td>.19 (.02)</td>
<td>83,064</td>
<td>.01</td>
</tr>
<tr>
<td>Non-OECD countries</td>
<td>.07 (.02)</td>
<td>.18 (.02)</td>
<td>46,420</td>
<td>.01</td>
</tr>
<tr>
<td>High-income OECD countries</td>
<td>.22 (.02)</td>
<td>.17 (.05)</td>
<td>32,809</td>
<td>.02</td>
</tr>
<tr>
<td>Euro area</td>
<td>.27 (.03)</td>
<td>.31 (.07)</td>
<td>7,856</td>
<td>.03</td>
</tr>
<tr>
<td>Japan</td>
<td>.21 (.04)</td>
<td>.17 (.06)</td>
<td>5,733</td>
<td>.02</td>
</tr>
<tr>
<td>Canada</td>
<td>.23 (.12)</td>
<td>.12 (.10)</td>
<td>16,221</td>
<td>.01</td>
</tr>
<tr>
<td>Differentiated</td>
<td>.12 (.01)</td>
<td>.17 (.03)</td>
<td>21,360</td>
<td>.02</td>
</tr>
</tbody>
</table>

Note: The first column reports pass-through conditional on price adjustment of the bilateral exchange rate shocks. The second column reports the pass-through of the component of the U.S. trade-weighted exchange rate orthogonal to the bilateral exchange rate (i.e., a residual from the projection of the trade-weighted exchange rate on the bilateral exchange rate). Clustered standard errors are in parentheses.
CPI in the foreign country. We again restrict the sample to nonpetrol, dollar-priced goods and market transactions. We include fixed effects for every BLS-defined primary strata (mostly two- to four-digit harmonized codes) and country pair; the standard errors are clustered at the level of the fixed effects.

The results are reported in table 5. The first row of each panel (labeled “No $\Delta_{t_1} P_{k,t}^I$”) presents the results where we exclude any industry competition effect. The next two rows include industry price effects aggregated at the BLS-defined primary strata level (mostly two- to four-digit harmonized codes) and at the 10-digit harmonized code level, respectively. As is evident in all specifications, the effect of competitor prices is large and highly significant. Moreover, it significantly reduces the direct response of prices to the exchange rate shock. If one does a back-of-the-envelope calculation of the extent of strategic complementarities using expression (2) of our accounting framework in Section II, one obtains a measure of markup elasticity of $\Gamma \approx 1.5$ (from $\Gamma/[1 + \Gamma] \approx 0.6$). This value is consistent with the required markup elasticity to match the evidence of incomplete long-run pass-through, as we discuss in Gopinath and Itskhoki (2010a) and Gopinath et al. (2010). Furthermore, note that the direct impact of exchange rate changes ($\beta_e$) still increases from the specification conditional on one price adjustment (panel A of table 5) to the life-long specification (panel B of table 5), even when we control for competitor prices. This suggests that, although

<table>
<thead>
<tr>
<th>Table 5</th>
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</thead>
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<tr>
<td><strong>Response to Competitor Prices</strong></td>
</tr>
</tbody>
</table>

<table>
<thead>
<tr>
<th></th>
<th>$\beta_e$</th>
<th>$\beta_I$</th>
<th>N</th>
<th>$R^2$</th>
</tr>
</thead>
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<tr>
<td><strong>A. Pass-through conditional on first price change:</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>No $\Delta_{t_1} P_{k,t}^I$</td>
<td>.13</td>
<td>...</td>
<td>83,056</td>
<td>.01</td>
</tr>
<tr>
<td>($\text{primary strata}$)</td>
<td>(.01)</td>
<td></td>
<td></td>
<td></td>
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<tr>
<td>$\Delta_{t_1} P_{k,t}^I$</td>
<td>.07</td>
<td>.61</td>
<td>78,942</td>
<td>.13</td>
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<tr>
<td>($\text{10-digit harmonized code level}$)</td>
<td>(.01)</td>
<td>(.02)</td>
<td></td>
<td></td>
</tr>
<tr>
<td><strong>B. Long-run pass-through:</strong></td>
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<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>No $\Delta_{t_1} P_{k,t}^I$</td>
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<td>...</td>
<td>16,145</td>
<td>.06</td>
</tr>
<tr>
<td>($\text{primary strata}$)</td>
<td>(.03)</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>$\Delta_{t_1} P_{k,t}^I$</td>
<td>.13</td>
<td>.66</td>
<td>15,273</td>
<td>.24</td>
</tr>
<tr>
<td>($\text{10-digit harmonized code level}$)</td>
<td>(.02)</td>
<td>(.03)</td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

Note: These are results from estimation of eq. (7). Standard errors are in parentheses. The coefficient $\gamma$ on the consumer-price inflation in the foreign country also shrinks along with $\beta_e$ when we include the control for competitor prices.
strategic complementarities are an important feature of price setting, they do not fully explain the delayed pass-through of exchange rate shocks and that there are other sources of real rigidity present in the data.

**Sector concentration.** Finally, we relate the incompleteness in exchange rate pass-through to certain sectoral features that proxy for the level of competition among importers. An important distinction between retail and wholesale prices is that the latter originate from business-to-business transactions. Consequently, the strength of bargaining power of the buyer can have an impact on the extent of pass-through. To evaluate this hypothesis, we use measures of concentration in the import sectors constructed by the Bureau of the Census. This includes a Herfindahl index for importers and a measure of the number of importers that make up the top 50% of trade using census data on all imports entering the United States. We estimate the following long-run pass-through regression in which we interact the exchange rate change with a measure of concentration (at the two- to four-digit level):

$$\Delta \bar{p}_{i,k,t} = \beta \Delta \tau_{e,i,k,t} + \psi (\Delta \tau_{e,i,k,t} \times C_k) + Z_{i,k,t}' \gamma + \epsilon_{i,t},$$

(8)

where the second regressor is the interaction of the exchange rate change with a given concentration measure. In $Z_{i,k,t}$ we include separate controls for the concentration measure, the change in CPI inflation (both stand alone and interacted with the concentration measure), and country fixed effects. All standard errors are clustered at the primary strata level.

The results are reported in table 6 for the two measures of concentration. While the point estimates in both regressions suggest that sectors that are dominated by a few large importers (high Herfindahl index and small number of firms in the top 50%) have lower pass-through from foreign firms, the standard errors on these estimates are large. Overall, the evidence is inconclusive. We also performed this exercise for pass-through conditional on first price change, as well as restricted

<table>
<thead>
<tr>
<th></th>
<th>$\beta$</th>
<th>$\psi$</th>
<th>$N$</th>
<th>$R^2$</th>
</tr>
</thead>
<tbody>
<tr>
<td>Herfindahl index</td>
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<td>-.02</td>
<td>12,432</td>
<td>.06</td>
</tr>
<tr>
<td></td>
<td>(.03)</td>
<td>(.02)</td>
<td></td>
<td></td>
</tr>
<tr>
<td>No. importers in top 50%</td>
<td>.23</td>
<td>.02</td>
<td>12,435</td>
<td>.06</td>
</tr>
<tr>
<td></td>
<td>(.02)</td>
<td>(.02)</td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

Note: Clustered standard errors are in parentheses.
the sample to differentiated goods only; in all cases we found no clear relationship in the data.

Summary. The presented evidence is consistent with an important role of real rigidities in the data, reflected in a sluggish response of prices to exchange rate shocks even conditional on prices changing. The reset-price inflation series, when conditioned on exchange rate shocks, also exhibits persistence. We further find that the response of prices is greater to trade-weighted exchange rate shocks as compared to bilateral shocks and that the response to competitor prices is significant. These last two pieces of evidence are consistent with the importance of strategic complementarities in price setting; however, they are also possibly consistent with other explanations. For example, the trade-weighted exchange rate can reflect changes in the prices of inputs imported from the rest of the world, while competitor prices may proxy for sectoral shocks, which are not reflected in the exchange rate and which are common across firms from a number of countries. Presumably the most direct evidence of strategic complementarity arises from studies of pricing to market that control for costs, as discussed earlier in Section III. The best description of our evidence is as important facts regarding the dynamics of price adjustment that should be matched by models of price setting. We turn to one such model in the next section.

V. Model

In this section we quantitatively evaluate a reduced-form sticky-price model with a retail sector and a wholesale sector. Consistent with the data, we allow for variable markups at the wholesale level and constant markups at the retail level. In the existing monetary literature, there is typically no interesting distinction made between the retail and wholesale sectors. The goal is to evaluate the behavior of regular and reset-price inflation, both unconditional and conditional on aggregate shocks, as well as the dynamic response of good-level prices conditional on changing so as to compare it to the evidence in Section IV. We also evaluate the extent of monetary nonneutrality generated by this source of real rigidities.

The model generates sluggishness in response to monetary shocks in wholesale prices, and this feeds into the slow adjustment of retail prices. However, aggregate inflation and reset-price inflation exhibit little persistence, since their movements are dominated by more transitory shocks. However, conditional on monetary shocks or exchange rate–like shocks, inflation series exhibit considerable persistence. This is consistent with the data. However, the model fails to match the slow dynamics
in price adjustment documented in the empirical data. This suggests a need for an additional source of persistence in prices.

Similarly, the output series can exhibit significant monetary non-neutralities if the money growth process is sufficiently persistent. While calibrated real rigidities in the form of variable markups increase the size of the contract multiplier, their effects are modest unless they are coupled with exogenous sources of persistence.

We begin by laying out the familiar equilibrium conditions of the model; the details are relegated to appendices. We first consider the model with Calvo price setting (app. A) and later evaluate the robustness of the predictions in a menu cost model of price setting in the wholesale sector (app. C). We should clarify that this is only a numerical exercise where the empirical evidence disciplines some parameters of the model.

A. Setup of the Model

*Wholesale sector.* Wholesale firms use labor and a constant returns to scale production function to produce intermediate goods. Therefore, a wholesale firm $j$ faces a constant marginal cost $mc_{jt}$ (all variables in logs):

$$mc_{jt} = w_t + \phi_j e_t - a_{jt},$$

where $w_t$ is the nominal wage rate and $e_t$ captures an exogenous exchange rate–like shock that affects the wholesale firm with elasticity $\phi_j$ that varies across firms. Further, $a_{jt} = \bar{a}_t + \tilde{a}_{jt}$ is the sum of an aggregate (wholesale-sector-wide) and idiosyncratic (firm-specific) shock to the firm; $a_{jt}$ represents some combination of shocks to the marginal cost and the markup that affects the firm’s desired price.

The desired log price of a wholesale producer equals a log desired markup over the marginal cost:

$$\tilde{s}_{jt} = \mu_{jt} + mc_{jt}.$$

By desired prices we mean prices that a firm would set if it could adjust prices every period in a given general equilibrium environment; desired price is not the same as a reset price, which is set for a number of future periods.

We assume variable markups that depend on the firm’s relative price:

$$\mu_{jt} = \bar{\mu} - \Gamma(s_{jt} - s_t),$$
where \( \bar{\mu} \) is the steady state level of markup, \( \Gamma \) is the elasticity of markup with respect to price, and

\[
s_t = \int s_{tj} dj
\]

is (an approximation to) the price index in the wholesale sector. This specification is a first-order approximation to a more general model of variable markups. For example, it can be obtained from the Kimball demand with nonconstant elasticity (e.g., see Klenow and Willis 2006; Gopinath and Itskhoki 2010a) or a model of strategic interactions between large firms (e.g., see Atkeson and Burstein [2008] and the bargaining model in Gopinath and Itskhoki [2010b]).

**Retail sector.** In the retail sector, firms combine labor and intermediate goods supplied by the wholesale sector to produce a final good. Specifically, firm’s marginal cost is given by

\[
mc_{Rit} = \alpha s_t + \left(1 - \alpha\right) w_t - z_{it},
\]

where \( \alpha \) is the production cost share of intermediate goods and \( z_{it} = \bar{z}_i + \tilde{z}_{it} \) is the sum of aggregate (retail-sector-wide) and idiosyncratic (firm-specific) marginal cost and/or markup shocks that affect the firm’s desired price. Note that we assume that the exchange rate shock, \( e_t \), does not affect the retail sector directly and that each retail firm uses a full bundle of intermediate goods as input in production.

We assume constant-markup pricing in the retail sector (e.g., monopolistic competition and CES demand), so that the desired price of firm \( i \) is given by

\[
\hat{p}_{it} = \bar{\mu} + mc_{Rit}.
\]

In the notation of Section II, it is equivalent to assuming that \( \Gamma_R = 0 \). This assumption, along with \( \Gamma > 0 \), is consistent with the evidence discussed in Section III. In Gopinath and Itskhoki (2010b), we provide a bargaining-based model that can rationalize this difference.

**Wage rate and real output.** We assume that the nominal wage rate depends on the consumer (final-good) price level, \( p_t \), and aggregate nominal spending, \( m_t \):

\[
w_t = \gamma m_t + (1 - \gamma) p_t, \quad \gamma > 0.
\]
This reduced-form model of wages is common in macroeconomics (e.g., see Chari, Kehoe, and McGrattan 2000; Burstein and Hellwig 2007) and can be derived, for example, from a cash-in-advance model of money demand and the intratemporal optimality condition for consumption-leisure choice.24 Smaller values of γ imply a more gradual response of wages to aggregate nominal spending shocks and hence are a stand-in for various unmodeled aggregate real rigidities, such as real wage rigidity (Blanchard and Gali 2007), segmented labor markets (Woodford 2003), and round-about production structure (Basu 1995).

With our definition of $m_t$ as aggregate nominal spending, real output is given by

$$y_t = m_t - p_t.$$ 

Therefore, the extent of monetary nonneutrality can be measured as the persistence of $y_t$ in response to nominal spending shocks, since in a flexible-price world exogenous $m_t$ shocks have no effect on real output.

**Exogenous shock processes.** As commonly assumed in the literature (e.g., see Chari et al. 2000; Bils et al. 2009), nominal spending $m_t$ follows an exogenous AR(1) process in first differences:

$$\Delta m_t = \rho_m \Delta m_{t-1} + \sigma_m \epsilon^m_t,$$

where $\rho_m \geq 0$ is the measure of exogenous persistence in the model. All other exogenous shocks follow persistent but stationary AR(1) processes:

$$x_t = \rho_x x_{t-1} + \sigma_x \epsilon^x_t,$$

where $x_t \in \{e_t, \tilde{z}_t, \tilde{a}_t, \tilde{z}_{it}, \tilde{a}_{it}\}$. All innovations ($\epsilon^m_t$ and $\epsilon^x_t$) are mean-zero unit-variance i.i.d. (independent and identically distributed) random variables.

1. **Calvo Price Setting**

In the case of Calvo price setting, a given wholesale firm $j$ adjusts prices with probability $(1 - \theta)$ each period, while for any retail firm $i$ the adjustment probability equals $(1 - \theta_R)$. Up to a first-order approximation, at the instances of adjustment the firms set their prices to the discounted
expectation of their future desired prices (see app. A for a formal derivation):

\[
\tilde{s}_t = (1 - \beta \theta) \sum_{l=0}^{\infty} (\beta \theta)^l \mathbb{E}_t \tilde{s}_{t+l},
\]

\[
\tilde{p}_t = (1 - \beta \theta_R) \sum_{l=0}^{\infty} (\beta \theta_R)^l \mathbb{E}_t \tilde{p}_{t+l},
\]

where \( \beta \) is the discount factor, \( \tilde{s}_t \) and \( \tilde{p}_t \) are the (theoretical) reset prices, and \( \tilde{s}_{t+l} \) and \( \tilde{p}_{t+l} \) are the future desired prices (derived above) for the wholesale and retail firms, respectively.

Under Calvo pricing assumptions, the dynamics of aggregate wholesale and retail prices are given, respectively, by

\[
st = \theta s_{t-1} + (1 - \theta) \mathbb{E}_t \tilde{s}_t
\]

and

\[
pt = \theta_R p_{t-1} + (1 - \theta_R) \mathbb{E}_t \tilde{p}_t,
\]

where \( \mathbb{E}_t \) and \( \mathbb{E}_t \) denote the cross-sectional means. Combining these equations with the expressions for reset prices and substituting in the expressions for desired prices, we arrive at the familiar forward-looking Phillips curves—dynamic equations for aggregate wholesale and retail inflation (see app. B):

\[
\Delta s_t = \beta \mathbb{E}_t \Delta s_{t+1} + \frac{\lambda}{1 + \Gamma} \left[ \gamma (m_t - p_t) - (s_t - p_t) + \delta e_t - \tilde{a}_t \right],
\]

(10)

\[
\Delta p_t = \beta \mathbb{E}_t \Delta p_{t+1} + \lambda_R [\alpha (s_t - p_t) + (1 - \alpha) \gamma (m_t - p_t) - \tilde{a}_t],
\]

(11)

where the expressions in square brackets are the average marginal costs of retail and wholesale firms with \( w_t \) substituted in from (9). Note that all idiosyncratic shocks wash out from the aggregate price dynamic equations. The slopes of the Phillips curves equal \( \lambda \equiv (1 - \beta \theta)(1 - \theta) / \theta \) and analogously for \( \lambda_R \). Finally, \( \delta \equiv \int \phi_j d\tilde{a} \) is the sensitivity of the average wholesale marginal cost to exchange rate shock, \( e_t \).

Dynamic equations (10)–(11), together with the specifications for the exogenous shock processes, fully describe equilibrium dynamics in the case of Calvo pricing. The solution to this dynamic system can be obtained numerically using a conventional Blanchard and Kahn (1980) method.\(^{25}\)

Finally, in the Calvo case, reset-price inflation can be measured simply as

\[
\Delta s^*_t = (\Delta s_t - \theta \Delta s_{t-1}) / (1 - \theta)
\]

and

\[
\Delta p^*_t = (\Delta p_t - \theta_R \Delta p_{t-1}) / (1 - \theta_R),
\]

since the adjusting firms are selected randomly. Given the equilibrium dynamics of the aggregate variables, we can simulate firm-level prices by using the expressions for optimal reset prices provided above.
Aggregating firm-level prices, we arrive at the sample measures of regular and reset-price inflation, the counterparts to the empirical measures studied in Section IV.

**Nominal and real rigidities in the model.** We now discuss the sources of nominal and real rigidity in the model. First, nominal stickiness enters through the Calvo parameters $\theta$ and $\theta_R$ that reduce the slopes of the Phillips curves ($\lambda$ and $\lambda_R$) and increase the persistence of inflation. Real rigidities in the form of variable markups, as measured by $\Gamma$, further reduce the slope of the wholesale inflation Phillips curve and contribute to the sluggish adjustment of wholesale prices. Furthermore, aggregate real rigidities, measured inversely by $\gamma$, slow down the pass-through of monetary shocks into the marginal costs of both types of firms and reduce the slopes of the Phillips curves. Finally, the share of intermediate inputs in the final-good production costs $\alpha$ links retail marginal costs to wholesale prices and constitutes a channel through which sluggish adjustment in wholesale prices translates into persistence in retail prices.

2. **Calibration**

We calibrate the model to monthly data and summarize the benchmark parameters in Table 7. We set the discount rate to 4% annually, which implies a monthly discount factor $\beta = 0.961^{1/12}$. We calibrate the money growth process and exchange rate process to the data. Specifically, we use the monthly BEA (U.S. Bureau of Economic Analysis) data on M2 supply to calibrate $\rho_m = 0.5$ and $\sigma_m = 0.25\%$. Other papers in the literature use different numbers for the persistence of money growth. For example, Chari et al. (2000) use $\rho_m = 0.571^{1/3} \approx 0.83$, while Bils et al. (2009) use $\rho_m = 0$. Therefore, for robustness we also simulate the model for $\rho_m = 0$ and 0.8. Next, we let the exchange rate follow a very persistent AR(1) process, with the standard deviation of innovation equal to $\sigma_e = 2\%$ and autocorrelation parameter $\rho_e = 0.995$, consistent with the data on bilateral nominal exchange rates for developed countries.

We select the parameters for the idiosyncratic shock processes ($\sigma_a, \sigma_z, \rho_a$, and $\rho_z$) to match the micro data on price adjustment. Specifically, we set the persistence of idiosyncratic shocks to match the high autocorrelation of new prices in the BLS IPP (International Price Program, which reports U.S. import and export price indexes) data (0.77 for import prices), and we set the standard deviation of idiosyncratic shocks to match the absolute size of price adjustment (7.5% for import prices and 8.5% for consumer prices). This results in the standard
deviation of idiosyncratic shocks equal to 10% and 8% for wholesale and retail prices, respectively, while the persistence is set to 0.95 and 0.9, respectively.

Next, we set the parameters for the aggregate shock processes ($\hat{\alpha}_t$, $\hat{\alpha}_z$, $\hat{\rho}_{zt}$, and $\hat{\rho}_z$) to match the standard deviation and autocorrelation of the regular and reset-price inflation series, reported previously in table 1. This requires fairly large and transitory aggregate shocks at both the wholesale and retail levels (standard deviations of 4% and 5% and persistence of 0.75 and 0.5, respectively). These processes are a stand-in for all unmodeled shocks that hit the economy, including various economy-wide and industry-level marginal cost and markup shocks, as well as measurement error in prices.

We set the Calvo probabilities of nonadjusting prices ($\theta$ and $\theta_R$) to match the micro data on nominal price durations. Specifically, we choose

| Table 7 | Benchmark Parameters |
|-----------------|-------------|----------|
| Parameter       | Symbol     | Value    | Source |
| Discount factor | $\beta$    | $0.96^{1/12}$ | Monthly data |
| Money growth process, $\Delta m_t$: | $\sigma_m$ | 0.25% | BEA data on M2 |
| Volatility (%)  | $\rho_m$  | 0 or 0.5 | |
| Persistence     | $\rho_e$  | 0.95     | |
| Exchange rate process, $\Delta e_t$: | Volatility | $\sigma_e$ | 2% | OECD exchange rates |
| Persistence     | $\rho_e$  | 0.95     | |
| Retail idiosyncratic shocks, $\hat{z}_{it}$: | Volatility | $\hat{\alpha}_z$ | 8% | Size of price adjustment of 8.5% |
| Persistence     | $\hat{\rho}_z$ | 0.90 | |
| Wholesale idiosyncratic shocks, $\hat{a}_{jt}$: | Volatility | $\hat{\alpha}_a$ | 10% | Size of price adjustment of 7.5% |
| Persistence     | $\hat{\rho}_a$ | 0.95 | Persistence of new prices of 0.77 |
| Retail aggregate shocks, $\hat{z}_t$: | Volatility | $\hat{\alpha}_z$ | 5% | Volatility and persistence of CPI regular and reset-price inflation |
| Persistence     | $\hat{\rho}_z$ | 0.50 | (Bils et al. 2009) |
| Wholesale aggregate shocks, $\hat{a}_t$: | Volatility | $\hat{\alpha}_a$ | 4% | Volatility and persistence of IPP regular and reset-price inflation |
| Persistence     | $\hat{\rho}_a$ | 0.75 | |
| Calvo parameters: | Retail | $\theta_R$ | 0.75 | Duration of 4 months, CPI data |
| Wholesale      | $\theta$  | 0.90     | Duration of 10 months, IPP data |
| Share of intermediate inputs | $\alpha$ | 0.5 | Nakamura and Steinsson (2010) |
| Wholesale markup elasticity | $\Gamma$ | 1.5 | Evidence on pass-through |
| Aggregate real rigidities | $\gamma$ | 0.75 |
| Sensitivity to the exchange rate shock | $\phi$ | 0.225 | Gopinath and Itskhoki (2010a) |

parameters to produce 10-month durations in the wholesale sector (consistent with the evidence in Gopinath and Rigobon [2008] and Nakamura and Steinsson [2008]) and 4-month durations in the retail sector (consistent with Bils and Klenow [2004]). Next, we calibrate $\gamma$, the slope of the wage equation (9) and the aggregate real rigidity parameter of the model. The literature uses a wide variety of values for $\gamma$, ranging between 0.1 in models with segmented labor markets and round-about production and 4 in models with no real rigidities and strong concavity in the utility function. We set the benchmark value for $\gamma$ to be 0.75, and for robustness we also use a greater value of 1.5, so that these parameters lie on both sides of the strategic neutrality case of $\gamma = 1$ and depart only moderately from it. We view this as a conservative choice for an aggregate parameter for which we have little direct information.

We choose the benchmark value for markup elasticity to be $\Gamma = 1.5$. This number implies a 40% pass-through of idiosyncratic shocks and is consistent with the evidence in Gopinath et al. (2010) and Gopinath and Itskhhoki (2010a) on long-run exchange rate pass-through of about 50%. Moreover, it is consistent with the coefficients on the competitor prices reported in table 5 of Section IV. We additionally evaluate the robustness of our results using the values of $\Gamma$ of 0 and 4, the former being the case of constant markups and the latter being the case of strong strategic complementarities at the firm level.27

Finally, we set the share of intermediate inputs in the final-good production, $\alpha$, to equal 50%, at the conservative end of the spectrum of calibrations considered in Nakamura and Steinsson (2010). The sensitivity of the aggregate marginal cost to exchange rate shocks is set to $\phi = 0.225$, which is consistent with most domestic firms being unaffected by this shock directly while a small fraction of importers in the industry (e.g., 30%) are affected strongly by this shock (e.g., $\phi_j = 0.75$).28

B. Simulation Results

First, we study the persistence of regular and reset-price inflation generated by our model. We compute all inflation series as sample averages of the simulated firm prices using a procedure close to the one used on the BLS data in Section IV.29 The results are reported in table 8 for different values of parameters $\rho_m$, $\Gamma$, and $\gamma$. The first two columns report the results for the final-good (retail) inflation series, the next two columns provide the results for the unconditional wholesale inflation series, and the last two columns provide the results for the projection of wholesale inflation series onto lags of exchange rate changes. For wholesale prices, we
use only the subsample of foreign firms, that is, those that are affected directly by the exchange rate shock (with $\phi_j > 0$), to make this exercise as close as possible to our empirical evidence in Section IV, which uses data on import prices.

The first pattern that emerges from table 8 is that the aggregate consumer-price inflation series may not be very persistent even when wholesale inflation is significantly more persistent. Second, wholesale inflation is significantly less persistent than the wholesale inflation projected on the exchange rate. Next we examine reset-price inflation. Both for retail and wholesale prices, reset-price inflation is negatively auto-correlated while, when projected on the exchange rate, the autocorrelation becomes positive. These patterns are consistent with the empirical findings in Section IV.A. In our calibration, negative autocorrelation of the reset-price inflation series arises due to sampling error combined with transitory semi-aggregate shocks affecting wholesale and retail pricing.

The results in table 8 are largely similar across different parameter values considered. Higher values of markup elasticity $\Gamma$ result in a more persistent (wholesale) inflation series, particularly when projected on the exchange rate shock, while variation in $\gamma$ and $\rho_m$ has relatively little effect on persistence of the inflation series. There are no clear patterns in variation of retail price persistence for different parameter values. This is because in our calibration monetary shocks are not the key drivers of the

### Table 8

<table>
<thead>
<tr>
<th></th>
<th>Unconditional</th>
<th>Conditional on Exchange Rate</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>$\Delta p_t$</td>
<td>$\Delta p_t^*$</td>
</tr>
<tr>
<td>$\Gamma = 0$</td>
<td>.41</td>
<td>-.25</td>
</tr>
<tr>
<td>$\Gamma = 1.5$</td>
<td>.32</td>
<td>-.25</td>
</tr>
<tr>
<td>$\Gamma = 4$</td>
<td>.37</td>
<td>-.28</td>
</tr>
<tr>
<td>$\rho_m = 0$</td>
<td>.39</td>
<td>-.23</td>
</tr>
<tr>
<td>$\rho_m = 0.8$</td>
<td>.48</td>
<td>-.14</td>
</tr>
<tr>
<td>$\gamma = 1.5$</td>
<td>.37</td>
<td>-.27</td>
</tr>
<tr>
<td>Menu cost</td>
<td>...</td>
<td>...</td>
</tr>
</tbody>
</table>

Note: The entries are AR(1) coefficients for each series ($\Delta p$ refers to final-good inflation and $\Delta s$ refers to wholesale inflation for the subsample of foreign firms with $\phi_j > 0$; * indicates reset-price inflation). In the last two columns, the series are the projections on the current and 24 lags of the exchange rate changes. All inflation series are sample averages of the simulated firm prices, approximating the procedure used in the data. The default parameters are the benchmark parameters from table 7 (i.e., $\Gamma = 1.5, \gamma = 0.75$, and $\rho_m = 0.5$).
inflation series in the short run (at the monthly frequency); this appears to be a reasonable description of reality.30

We next evaluate the extent of monetary nonneutrality produced by the model. Table 9 reports the half-lives of output in response to a monetary shock for different values of the parameters. Specifically, we calculate the AR(1) coefficient of the real output series when all shocks other than monetary shocks are shut down; based on this AR (1) coefficient, we back out a measure of half-life that we report. From table 9 it is evident that the model can produce a wide range for the extent of nonneutrality, with half-lives of output ranging from about 1 quarter to over 20 quarters. However, this variation is largely driven by \( \rho_m \), the exogenous persistence introduced through the autocorrelation of the money growth rate. However, variation in the amount of real rigidity, \( \Gamma \) and \( \gamma \), has a relatively modest effect on the extent of non-neutrality: an increase in \( \Gamma \) from 0 to 4 nearly doubles the half-life, while a decrease in \( \gamma \) from 1.5 to 0.75 increases the half-life by around 50%.

When we fix parameters at their benchmark values, the model produces a fairly large half-life of slightly below 8 quarters, while shutting down the variable markup channel drops the half-life to less than 6 quarters. Without exogenous persistence (i.e., \( \rho_m = 0 \)), however, the model produces very little monetary nonneutrality (a half-life of around 1 quarter).31 We conclude that the empirically calibrated variable markup channel of real rigidities goes a fair way in amplifying the real effects on output; however, without exogenous persistence its absolute effect is modest.

Note that the variation in the persistence of output deviation in the model is not very tightly linked to the persistence of inflation, which does

### Table 9
Half-Life of Output in Response to a Monetary Shock (in Months)

<table>
<thead>
<tr>
<th></th>
<th>( \rho_m = 0 )</th>
<th>( \rho_m = .5 )</th>
<th>( \rho_m = .8 )</th>
</tr>
</thead>
<tbody>
<tr>
<td>A. ( \gamma = 0.75 ):</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>( \Gamma = 0 )</td>
<td>5.3</td>
<td>17.2</td>
<td>56.1</td>
</tr>
<tr>
<td>( \Gamma = 1.5 )</td>
<td>7.0</td>
<td>23.7</td>
<td>83.0</td>
</tr>
<tr>
<td>( \Gamma = 4 )</td>
<td>8.9</td>
<td>31.3</td>
<td>114.8</td>
</tr>
<tr>
<td>B. ( \gamma = 1.5 ):</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>( \Gamma = 0 )</td>
<td>3.6</td>
<td>11.8</td>
<td>40.0</td>
</tr>
<tr>
<td>( \Gamma = 1.5 )</td>
<td>4.4</td>
<td>15.5</td>
<td>58.1</td>
</tr>
<tr>
<td>( \Gamma = 4 )</td>
<td>5.4</td>
<td>19.8</td>
<td>80.0</td>
</tr>
</tbody>
</table>

Note: Half-life is defined as \( \log(0.5)/\log(\rho_{y/\varepsilon_m}) \), where \( \rho_{y/\varepsilon_m} \) equals the AR(1) coefficient of output \( y_t = m_t - p_t \) conditional on monetary shocks \( \varepsilon_m \) (i.e., when other sources of shocks are shut down).
not vary much with the amount of real rigidities or the extent of exogenous persistence. Therefore, fairly long periods of monetary non-neutrality can be consistent with transitory inflation dynamics and negatively autocorrelated reset-price inflation. Again, this is because monetary shocks are not the main drivers of inflation at very high frequencies.

Our final results evaluate the success of the model at capturing the slow response of firm-level prices to exchange rate shocks at the micro level, conditional on price adjustment. This is done for wholesale foreign price with $\phi_j > 0$ to approximate the empirical analysis of Section IV.B. The results are reported in table 10. When there are no strategic complementarities, the second adjustment is negative (due to some mean reversion that we assumed in the exchange rate process). When strategic complementarities are present, pass-through at the second round of price adjustment is positive; however, it is much smaller than it is in the data (see table 2), where the second rounds of price adjustment are almost as large as the first. This failure of the model persists even when we assume strong strategic complementarities ($\Gamma = 4$) or shorter nominal price durations (not reported). Although the model with strategic complementarities captures incomplete pass-through in the long-run, it predicts very fast dynamics of pass-through relative to the data. This leads us to conclude that our model misses some important sources of persistence, which may further contribute to the extent of monetary nonneutrality produced by the model. Matching the very slow adjustment of prices to aggregate shocks at the micro level that we document in Section IV is an important challenge that we leave for future work.

C. A Menu Cost Model

In this section we briefly describe the setup and provide the simulation results of a menu cost model of price setting. The details are relegated
to appendix C, and further discussion of the estimation procedure can be found in Gopinath and Itskhoki (2010a). This exercise is important in order to evaluate the severity of the selection effects present in menu cost models and absent in the time-dependent pricing models.

We adopt a two-sector model (wholesale and retail, as above) with three types of shocks: idiosyncratic marginal cost shocks, semi-aggregate marginal cost shocks (a stand-in for exchange rate shocks) in the wholesale sector, and aggregate monetary (nominal spending) shocks. In order to maintain computational feasibility, we assume that retail prices are completely flexible. We introduce the variable markup channel, using the Klenow and Willis (2006) specification of the Kimball (1995) preferences. The rest of the setup is similar to the one discussed above in Section V.A.

With flexible prices, no strategic complementarities (e.g., CES [constant elasticity of substitution] demand), and no aggregate shocks in the retail sector, the final-good price level is given by

\[ p_t = \mu_R + \alpha s_t + \frac{1}{C_0} \alpha \omega_t. \]

The wage rate is still assumed to satisfy (9). These two equations allow one to solve for the final-good price level \( p_t \) and the nominal wage rate \( \omega_t \) as functions of aggregate nominal spending \( m_t \) and the wholesale price level \( s_t \).

A wholesale firm \( j \) faces a marginal cost

\[ mc_j = \omega_t + \phi_j \epsilon_t - a_{jt}, \]

where \( a_{jt} \) is an idiosyncratic shock and \( \epsilon_t \) is the semi-aggregate shock that affects the firm with elasticity \( \phi_j \) distributed on \([0, 1]\). The firm also faces a demand schedule with elasticity \( \sigma \) and the superelasticity (elasticity of elasticity) \( \epsilon \), evaluated when the firm’s relative price equals one (for details, see Gopinath and Itskhoki [2010a] and app. C). This implies a desired markup of \( \sigma/(\sigma - 1) \) with markup elasticity equal to \( \Gamma = \epsilon/(\sigma - 1) \). A firm maximizes its discounted present value by optimally choosing the instances of price adjustment at a menu cost \( \kappa \) and optimally resetting prices at these instances. This problem can be formalized with a standard Bellman equation (see app. C), which we solve numerically. We then use the derived policy functions to simulate a panel of prices on which we conduct similar empirical tests to those in Section IV.

For calibration we use the same benchmark parameters as in table 7. Additionally, we set \( \sigma = 5 \) and \( \epsilon = 6 \) so that the level of wholesale markup is 25% and the elasticity of markup is equal, as before, to \( \Gamma = 1.5 \). The menu cost is chosen to match the duration of 10 months of wholesale prices that implies a \( \kappa = 3.5\% \) of the steady state revenue conditional on adjustment (corresponding to total annual menu cost paid equal to 0.35% of annual revenues), a number consistent with the literature. Finally, in order to match the absolute size of price adjustment
of 7.5%, we set the standard deviation of the idiosyncratic shocks to $\sigma_a = 6\%$.

To keep this section brief, we report the results only for the benchmark values of parameters, $\Gamma = 1.5$, $\gamma = 0.75$, and $\rho_m = 0.5$. The unconditional autocorrelation of the wholesale-price inflation is 0.29, and it is 0.38 conditional on the exchange rate shock, both numbers being substantially smaller than in the case of the Calvo model (see table 8). The corresponding reset-price inflation series is strongly negatively autocorrelated—with autocorrelation of $-0.89$ unconditionally and $-0.66$ conditional on the exchange rate shock—emphasizing the powerful selection effects of the menu cost models. This negative autocorrelation of reset-price inflation conditional on the exchange rate shock goes against our empirical findings in Section IV.

Similarly, in the type of micro-level pass-through regressions run in Section IV.B, the menu cost model generates a pass-through coefficient conditional on the first round of price adjustment equal to 55%, while the coefficient for the second round of price adjustment is $-8\%$. This is also due to the strong selection effect of the menu cost models that dominated the persistence introduced through strategic complementarities in pricing. Recall that in the Calvo model these two pass-through coefficients were 44% and 4%, respectively (see table 10), and this is in contrast to our empirical findings that pass-through at the second adjustment is nearly as large as the first one. Finally, the half-life of output in response to monetary shocks is 15.1 months in the menu cost model, as opposed to 23.7 months in the Calvo model. This illustrates a well-known fact that the selection effect of the menu cost models reduces substantially the contract multiplier relative to the time-dependent models of price setting.

**Summary.** A number of insights come out of our simulation exercise. First, transitory aggregate inflation series are consistent with a persistent response of prices to certain aggregate shocks, including exchange rate and monetary shocks. Second, properties of the aggregate inflation series may be largely disconnected from the size of the contract multiplier for monetary shocks. Third, quantitatively exogenous persistence appears to be more important than strategic complementarities in generating long half-lives and large contract multipliers. Although strategic complementarities work to magnify the size of the contract multiplier, their absolute effects are modest unless coupled with exogenous sources of persistence. Fourth, the analyzed models (in particular the menu cost model but also the Calvo model) cannot match the very sluggish dynamics of prices in response to shocks at the micro level conditional on adjusting prices. This
suggests a need for additional sources of persistence that are lacking from the model.

Appendix A

Calvo Price Setting

In this appendix we derive a general log-linear approximation for the price setting equation and aggregate inflation dynamics (Phillips curve) in a Calvo model. Since these derivations are well known, we keep the exposition brief.

Consider a firm $j$ with a real profit function $\Pi_j(x_j|S)$, where $x_j$ is the firm’s log price and $S$ is the state of the economy. The desired price of the firm is $\hat{x}_j(S) \equiv \arg\max_{x_j} \Pi_j(x_j|S)$ with the necessary condition $\Pi_j'(\hat{x}_j(S)|S) = 0$, where the subscript denotes a partial derivative. We assume that the marginal cost of the firm does not depend on the price of the firm; that is, a firm faces a constant returns to scale in production where productivity depends on the state of the world. Then, we can decompose the desired price as

$$\hat{x}_j(S) = \mu_j(\hat{x}_j - x, S) + mc_j(S),$$

where $mc_j(S)$ is the log nominal marginal cost of the firm and $\mu_j(x_j - x, S)$ is the log (desired) markup, which we allow to depend on the relative price of the firm, with $x$ denoting the log of the relevant price index.

A general first-order approximation to the markup can be written as

$$\mu_j(x_j, S) \approx \bar{\mu} - \Gamma(x_j - x) + \epsilon_j(S),$$

where $\bar{\mu}$ and $\Gamma$ are some constants (assumed to be common across all firms at the point of approximation) and $\epsilon_j(S)$ is some linear function of the state $S$. It is natural to assume that $\epsilon_j(S)$ is stationary, while $mc_j(S)$ is cointegrated with the nominal variables of the model. With this approximation, we can solve explicitly for the desired price of the firm:

$$\hat{x}_j(S) = \frac{1}{1 + \Gamma} [\bar{\mu} + mc_j(S) + \epsilon_j(S)] + \frac{\Gamma}{1 + \Gamma} x.$$
with exogenous probability \((1 - \theta)\). Therefore, we can write the problem of the firm recursively as

\[
\bar{\pi}_t(S) = \max_{x_j} \mathbb{E} \left\{ \Pi^j(x_j|S_t) + \sum_{l=1}^{\infty} Q_{t,t+l}(S) \theta^{l-1} \theta \Pi^j(x_j|S_{t+l}) \right. \\
\left. + (1 - \theta) \Pi^j(\bar{x}_{j,t+l}(S)|S_{t+l})|S_t^l \right\},
\]

where \(Q_{t,t+l}(S)\) is the stochastic discount factor for real variables, \(S^l \equiv (S_0, \ldots, S_t)\) is the history of the states, and \(S \equiv (S^l, S_{t+1}, \ldots)\). The first-order condition for the price setting can be written as

\[
\sum_{l=0}^{\infty} \theta^l \mathbb{E} \left\{ Q_{t,t+l} \Pi^j(\bar{x}_j|S_{t+l})|S^l \right\} = 0,
\]

where we omit the explicit dependence on \(S\). Taking a first-order approximation of this optimality condition around a nonstochastic steady state with zero inflation, we obtain

\[
\sum_{l=0}^{\infty} (\beta \theta)^l \mathbb{E} \left\{ \bar{x}_{jt} - \tilde{x}_j(S_{t+l})|S^l \right\} = 0,
\]

where \(\beta^l\) is the nonstochastic steady state value of \(Q_{t,t+l}\). The price setting formulas in Section V.A are direct implications of this linearized optimality condition. Now, using the expression for the desired price, we have

\[
\bar{x}_{jt} = (1 - \beta \theta) \sum_{l=0}^{\infty} (\beta \theta)^l \mathbb{E}_t \left\{ \frac{1}{1 + \Gamma} (mc_{j,t+l} + e_{j,t+l}) + \frac{\Gamma}{1 + \Gamma} x_{t+l} \right\},
\]

where we switched notation for conditional expectation, suppressed the explicit dependence on the state of the economy, and omitted the constant by implicitly relabeling the variables to denote the deviations from the nonstochastic steady state.

Finally, since the nominal marginal cost is possibly integrated, we need to scale this expression by some monetary variable cointegrated with the marginal cost. A natural candidate is the competitor price
index \( x_t \) or a sector price level (in a number of models, including the Kimball demand model, the two variables coincide). We therefore define \( x_t = \int x_{jt} \, dj \). With some manipulation, we rewrite the deflated price setting equation as

\[
\bar{p}_{jt}/C_0 - \bar{p}_t/C_0 = \frac{1}{\lambda} \left( \frac{\beta \theta}{1 + \Gamma} \right) \sum_{l=0}^{\infty} \left( \frac{\beta \theta}{1 + \Gamma} \right)^l E_t \{mc_{jt+l} - x_{jt+l} + \epsilon_{jt+l} \} + \sum_{l=0}^{\infty} \left( \frac{\beta \theta}{1 + \Gamma} \right)^l E_t \Delta x_{jt+l},
\]

or in a recursive form: 34

\[
(\bar{p}_{jt} - x_{t-1}) - \beta \theta E_t (\bar{p}_{jt+1} - x_t) = \frac{1 - \beta \theta}{1 + \Gamma} (mc_{jt} - x_t + \epsilon_{jt}) + \Delta x_t.
\]

Next, with Calvo pricing, the dynamics of the aggregate price level can be written as

\[
x_t = \theta x_{t-1} + (1 - \theta) E_j \bar{p}_{jt} \Rightarrow \Delta x_t = (1 - \theta) E_j \{\bar{p}_{jt} - x_{t-1} \},
\]

where \( E_j \) is the cross-sectional average across all firms. Combining the above two equations and rearranging, we arrive at the traditional New-Keynesian Phillips curve:

\[
\Delta x_t - \beta E_t \Delta x_{t+1} = \frac{\lambda}{1 + \Gamma} E_j \{mc_{jt} - x_t + \epsilon_{jt} \}, \quad \lambda = \frac{(1 - \beta \theta)(1 - \theta)}{\theta}.
\]

Equations (10)–(11) in the text are special cases of this Phillips curve with the expressions for marginal costs substituted in (note that the cross-sectional expectation averages out all purely idiosyncratic shocks).

**Appendix B**

**Dynamics under Calvo Pricing**

The aggregate dynamic system contains three equations—the two Phillips curves for the wholesale and retail prices and the aggregate wage equation—for three variables \((s_t, p_t, w_t)\):

\[
\Delta s_t = \beta E_t \Delta s_{t+1} + \frac{\lambda}{1 + \Gamma} (w_t - s_t + \bar{\epsilon}_t - \bar{a}_t),
\]
\[ \Delta p_t = \beta \mathbb{E}_t \Delta p_{t+1} + \lambda_R [\alpha s_t + (1 - \alpha) w_t - p_t - \bar{z}_t], \]

\[ w_t = \gamma m_t + (1 - \gamma) p_t, \]

where \( \Delta m_t, e_t, \bar{z}_t, \) and \( \bar{a}_t \) follow exogenous stationary processes. This system can be solved using the conventional Blanchard and Kahn (1980) method, which results in the expressions for the endogenous variables \((s_t, p_t, w_t)\) as functions of the shocks to the exogenous variables. This solution allows us to study the statistical properties of endogenous variable time series, including their volatility and persistence.

When the retail prices are set flexibly (i.e., \( \theta_R = 0 \) or \( \lambda_R = \infty \)), there exists a tractable analytical solution for the aggregate dynamics. We discuss it briefly here. In this case, the expression for the consumer price level becomes static:

\[ p_t = \alpha s_t + (1 - \alpha) w_t - \bar{z}_t. \]

Together with the wage equation it allows us to solve for \( p_t \) and \( w_t \) as linear functions of \( m_t, s_t, \) and \( \bar{z}_t \). Substituting these expressions into the wholesale price Phillips curve, we obtain a second-order difference equation in \( s_t - m_t \):

\[ \Delta s_t - \beta \mathbb{E}_t \Delta s_{t+1} = \kappa (s_t - m_t) + \xi_t, \]

where

\[ \xi_t = \frac{\lambda}{1 + \Gamma} (\bar{a}_t - \bar{a}_t) - \frac{\lambda}{1 + \Gamma} \frac{1 - \gamma}{\alpha (1 - \gamma) + \gamma} \bar{z}_t \]

is the summary measure of all shocks other than \( m_t \) and \( \kappa = \frac{\lambda}{1 + \Gamma} \{\gamma / (\alpha (1 - \gamma) + \gamma}\} \) is the summary measure of nominal and real rigidities in the model. This difference equation can be solved forward. Assuming for simplicity that \( \xi_t \) follows an AR(1), the process for \( s_t - m_t \) is an ARMA(3, 1). Therefore, \( s_t \) is cointegrated with \( m_t \) and movements in \( m_t \) dominate the low-frequency movements in \( s_t \); however, the short-run dynamics of \( s_t \) (around slow-moving \( m_t \)) may be dominated by transitory shocks to \( \xi_t \). In particular, the MA component may reduce significantly the short-run persistence, while it does not affect the long-run persistence. This logic is consistent with the empirical
findings of Stock and Watson (2007). Finally, one can show that one of the AR roots is given by $\rho_{mtr}$, while the other root is decreasing in $\kappa$ and converging to one as $\kappa \to 0$. Furthermore, one can show that these roots also drive the persistence of the output response to monetary shocks.

Appendix C

Menu Cost Model

Kimball demand. To simulate the menu cost model, we first need to specify the explicit source of variable markups. We generate variable markups by introducing the Klenow and Willis (2006) specification of the Kimball (1995) demand. The demand function for firm $j$ in this case is given by

$$\psi(s_{jt} - s_t) = [1 - \varepsilon(s_{jt} - s_t)]^{\sigma}/\varepsilon, \quad \sigma \geq 1, \quad \varepsilon > 0,$$

where $s_{jt}$ is the log price of the firm and $s_t = \int_j s_{jt} dj$ is the sectoral log price index. In Gopinath and Itskhoki (2010a), we show that this price index is a valid second-order approximation to the ideal price index with this demand system. The price elasticity of this demand is given by

$$\sigma = \frac{\sigma}{1 - \varepsilon(s_{jt} - s_t)},$$

which equals $\sigma$ when the relative price of the firm is one. With this demand, the desired price is equal to a markup $\tilde{\sigma}/(\tilde{\sigma} - 1)$ over the marginal cost. The elasticity of the markup is given by

$$\Gamma = \frac{\tilde{\varepsilon}}{\tilde{\sigma} - 1},$$

where

$$\tilde{\varepsilon} \equiv \frac{\tilde{\varepsilon}}{1 - \varepsilon(s_{jt} - s_t)}$$

is the elasticity of the elasticity of demand. Therefore, the markup elasticity evaluated at the relative price of one is given by $\Gamma = \varepsilon/(\sigma - 1) > 0$. 

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Problem of the firm. The real profit of the firm is given by

$$\Pi(s_j|S_j) = \psi(s_j - s_t)[\exp\{s_j - p_t\} - \exp\{mc_j - p_t\}],$$

where $S_j$ is the state vector for the firm, $mc_j$ is the log nominal marginal cost of the firm, and $p_t$ is the final-good price level. As discussed in the text, the marginal cost of the firm equals

$$mc_j = w_t + \phi_j e_t - a_j.$$ 

Therefore, the state vector for the firm includes $(p_t, s_t, w_t, e_t, a_j)$. We can write the firm’s problem recursively as

$$V^N(S_j) = \Pi(s_{j,t-1}|S_j) + E\{Q(S_{j,t+1})V(S_{j,t+1})|S_j\},$$

$$V^A(S_j) = \max_{\bar{s}_j}\{\Pi(s_{j,t}|S_j) + E\{Q(S_{j,t+1})V(S_{j,t+1})|S_j\}\},$$

$$V(S_j) = \max\{V^N(S_j), V^A(S_j) - \kappa\},$$

where $V$ is the value of the firm, $V^N$ is the value of the firm if it does not adjust its price, and $V^A$ is the value of the firm if it adjusts its price; $\kappa$ is the menu cost, $Q$ is the stochastic discount factor for real variables (which we set to equal $\beta$ in the simulation), and $S_j$ includes in addition the previous price of the firm $s_{j,t-1}$.

General equilibrium. We assume flexible prices for the final good and no aggregate productivity shocks in the final-good sector. This implies that $p_t = \mu_R + \alpha s_t + (1 - \alpha)w_t$, where $\mu_R$ is the constant markup in the final-good sector. In turn, the wage is given by $w_t = \gamma m_t + (1 - \gamma)p_t$. This allows us to solve for $p_t$ and $w_t$ as a function of $m_t$ and $s_t$ and reduce the aggregate state space to $(s_t, m_t, e_t)$. The state vector for an individual wholesale firm additionally includes $(s_{j,t-1}, a_j)$. In the general equilibrium of the model, firms optimally decide to adjust prices given their current state vector and rational expectations about the evolution of the state vector, while aggregated individual firm pricing decisions are consistent with the aggregate dynamics of the wholesale price level $s_t$.

Simulation procedure. We iterate the Bellman equation for the firm pricing problem on the grid given a forecasting rule for the evolution of the state vector. This produces a policy function for firm pricing decisions, which allows us to simulate a panel of firm prices. In each period of the
simulation, we make sure that the wholesale price index is consistent with the firm pricing decisions (which constitutes a static fixed point problem). As a result, we obtain a time series for the equilibrium wholesale price level. Given this time series, we update the forecasting equation. We iterate this procedure until the forecasting equation converges. With the equilibrium forecasting rule we simulate a panel of firm prices, and we use it to estimate various statistical moments, as in Section IV. Additional details of the simulation procedure can be found in Gopinath and Itskhoki (2010a).

Endnotes

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2. For a precise definition, discussion, and examples of real rigidities, see the entry for that term by David Romer in the New Palgrave Dictionary of Economics (2008). The ability of real rigidities to magnify the effects of nominal price stickiness is often referred to as the contract multiplier, a term coined by John Taylor (1980) and commonly used in the macro literature (e.g., see Chari et al. 2000).

3. Of course, the exchange rate can have additional indirect (general equilibrium) effects via productivity or cost of inputs. We discuss this issue in more detail in Secs. III and IV.

4. In Sec. V, we combine this framework with specific models of nominal stickiness to study the quantitative predictions of dynamic price setting models.

5. Goldberg and Knetter (1997) provide a survey of the earlier empirical pass-through literature. The specification of pass-through regressions can either be in real or in nominal terms and accordingly include the real or the nominal exchange rate. Since the two move closely at most horizons, either specification gives similar results. Typical pass-through regressions also include some controls for the foreign cost level, such as the producer price index or the manufacturing wage rate.

6. See the literature following the seminal observation of Meese and Rogoff (1983).

7. Reference price (cost) is defined as the most often quoted price (cost) within a given time period.

8. Specifically, Fitzgerald and Haller (2008) find that, conditional on price adjustment in both markets, the common-currency price differential across the two markets moves nearly one-to-one with the nominal exchange rate. According to the descriptive model (2) of Sec. II, this is consistent with a \( \Gamma \gg 0 \) provided that the competitor prices \( s_i \) in the two markets move little with the exchange rate, as in the model of Atkeson and Burstein (2008).

9. Additionally, a recent study by Berman, Martin, and Mayer (2009) compares the extent of markup variability across French exporters of different size. It finds that larger exporters have lower exchange rate pass-through, which is consistent with a model in which both the level and the variability of the markup increases with the productivity
and hence size of the firm. Since exporting firms are typically larger, the implication of
this finding is that the extent of markup variability can be greater in the international data
than in the full sample of firms producing for the domestic market.

10. This feature is consistent with the evidence in Boivin, Giannoni, and Mihov (2009),
which finds that, while overall disaggregated prices are volatile, they are sluggish in re-
response to aggregate macro shocks.

11. In the sample that excludes petrol classifications, the median number of price ob-
servations per month is 6,335 and the median number of price observations whose price
is different from the previous month is 770.

12. The small number of price changes limits the analysis to large groups of goods.

13. We estimate an AR(1) coefficient so as to compare our results directly to Bils et al.
(2009). If, as argued by Stock and Watson (2007), CPI inflation is better modeled as an
ARMA(1, 1) or an IMA(1, 1), then the first-order autocorrelation understates the long-run
persistence of the series. We have estimated other measures of persistence, such as the
variance ratio to the long-run variance of the series. Although these measures suggest
greater persistence for the inflation series, our comparative results for the conditional
and unconditional inflation still hold.

14. In all specifications we exclude petrol classifications. For each series, we use 2002
weights at the four-digit level to aggregate across prices. More precisely, for actual infla-
tion, we estimate mean price change by four-digit harmonized code for each month, then
we average across the different harmonized codes using weights at the four-digit level.

fresh fruit and vegetables, and eggs. We report their results for the sample that excludes
sales price.

16. As documented by Stock and Watson (2007), among others, the short-run persist-
tence of consumer-price inflation decreased in the 1990s.

17. In the BLS database, the original reported price (in the currency of pricing) and the
dollar converted price are both provided. We use the latter, conditional on the original
reported price having changed. Since the first price adjustment is censored from the data,
we also perform the analysis excluding the first price change and find that the results are
not sensitive to this assumption.

18. Refer to Gopinath and Itskhoki (2010a) for a detailed analysis of pass-through con-
ditional on first adjustment and pass-through conditional on many rounds of adjustment
across goods with different frequencies.

19. The coefficient on the residual will be equivalent to the coefficient on the trade-
weighted exchange rate obtained from regressing the price change on the bilateral and
the trade-weighted exchange rate, but the coefficient on the bilateral exchange rate will be
different across the two specification.

20. For each good $i$ we calculate the average monthly import price change for all goods
in the same 10-digit or BLS-defined primary strata classification, excluding good $i$. We
then cumulate these price changes over time to arrive at an industry competitor price
index for each good. In our main specification, we include nonadjacent price changes;
that is, if prices are available for January and March but are missing for February, the
price change in March refers to the percentage difference between the March and January
price. We also perform an analysis in which we include only price changes across adjacent
months; we obtain qualitatively the same results.

21. Results are unaffected if we exclude the first price change that can be censored. For
more details on the comparison between life-long pass-through and pass-through condi-
tional on first price adjustment, see Gopinath and Itskhoki (2010a).

22. In Gopinath and Itskhoki (2010b), we discuss a bargaining-based micro foundation
for these reduced-form assumptions.

23. Note that our simple model does not rely on any mechanisms of differential response
to aggregate vs. idiosyncratic shocks as, e.g., in MacKowiak and Wiederholt (2009).

24. Specifically, a model with log utility of consumption and linear disutility of labor
results in $\gamma = 1$, provided that there is no additional source of aggregate real rigidities.
Golosov and Lucas (2007) derive the same specification in a money in the utility model.
Ball and Romer (1990) refer to this benchmark as the case of strategic neutrality. Aggregate real rigidities work to reduce the value of $\gamma$.

25. When final-good prices are flexible ($\theta_R = 0$), this dynamic system has a simple closed-form solution, which we discuss in app. B.

26. For details, see Gopinath and Itskhoki (2010a).

27. Our benchmark number of $\Gamma = 1.5$ is considerably smaller than the markup elasticity of 2.5 implied by the Klenow and Willis (2006) calibration.

28. For more details, see the calibration in Gopinath and Itskhoki (2010a).

29. We use a sample of 5,000 retail firms and 6,000 wholesale firms, 1,800 of which are affected directly by the exchange rate shock. With this sample size, low frequency of price adjustment, and large idiosyncratic price movements, the sampling error in average inflation series is nontrivial.

30. This feature is consistent with aggregate data where inflation is well approximated by an ARMA(1, 1) process with both large AR and MA roots (see, e.g., Stock and Watson 2007). While monetary shocks are likely to be responsible for the AR component (long memory, low-frequency movements), there needs to be a source of relatively large transitory shocks to explain the MA component.

31. This finding is consistent with the results in Carvalho and Nechio (2008) on the persistence of real exchange rates.

32. The approach to computing the conditional half-life of output in the menu cost model is different. Since we cannot simply shut down other sources of shocks in the menu cost model (as its dynamics is nonlinear in shocks), we estimate econometrically the impulse response of output $y_t$ to current and lagged monetary shocks $\varepsilon_{mt}^j$, and then use it to compute the projection of output series onto these shocks. The reported number is based on the AR(1) coefficient for the projected series.

33. Since we look at the real profit function, that is, a profit function normalized by the price level in the economy, it is without loss of generality to assume that the nominal variables enter the sufficient state vector $S$ only normalized by the price level. Therefore, we can treat $S$ as having a stationary distribution even though monetary variables may be trending. For an example, see app. C.

34. This is the step that requires stationarity of the right-hand-side variables.

35. This is under the assumption of a unit root in $\xi_t$; otherwise, the difference equation is second order in $s_t$.

36. If $\xi_t$ is absent from the model, the process for $s_t - m_t$ is an AR(2) and therefore $s_t$ exhibits both high short-run and long-run persistence.

37. To ensure stationarity of the grids, we normalize all nominal variables in the model by $m_t / C_0$.

References


