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Do Sticky Prices Increase Real Exchange Rate Volatility at the Sector Level?*

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Abstract

We introduce the real exchange rate volatility curve as a useful device to understand the role of price stickiness in accounting for deviations from the Law of One Price at the sector level. In the presence of both nominal and real shocks, the theory predicts that the real exchange rate volatility curve is a U-shaped function of the degree of price stickiness. Using sector-level European real exchange rate data and frequency of price changes, we estimate the volatility curve. The results are consistent with the predominance of real effects over nominal effects. Nonparametric analysis suggests the curve is convex and negatively sloped over the majority of its range. Good-by-good variance decompositions show that the relative contribution of nominal shocks is smaller at the sector level than what previous studies have found at the aggregate level. We conjecture that this is due to significant averaging out of good-specific real microeconomic shocks in the process of aggregation.

JEL Classification: E31; F31; D40

Keywords: Real exchange rates, Law of One Price, Sticky prices, Nonparametric test for monotonicity.

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1. Introduction

Among international macroeconomists, it is widely believed that the variability of real exchange rates is increasing in the degree of local currency price rigidity. The reasoning is found in passages describing exchange rate overshooting in prominent textbooks:

> "Exchange rate overshooting results from the rapid response of exchange rates to monetary policy and the sluggish adjustment of prices. A monetary expansion will lead to an immediate depreciation but only a gradual increase in prices. Exchange rate overshooting implies that real exchange rates are highly volatile."
> Dornbusch, Fischer, and Startz (2004, p. 534)

The basic idea is that the nominal exchange rate is an asset price (since currencies are actively traded in the foreign exchange market) and like all asset prices, its value adjusts instantaneously in response to unexpected changes in exogenous variables in current and future periods. In contrast, many goods and services have prices which economists have documented to be fixed in local currency for extended periods of time. By definition, this means nominal and real exchange rates will be highly correlated with each other over time and will have a comparable level of time series variability, at least at the highest frequencies. The expectation, then, is a *positive* correlation between the volatility of real exchange rates and the degree of price stickiness if *nominal* shocks dominate the landscape, as they do in much theorizing on the topic.

The New Open Economy Macroeconomics (NOEM) framework provides fully articulated models of the channels through which monetary shocks drive transitory fluctuations in real exchange rates. Quantitative investigations of this framework have been undertaken, by Chari, Kehoe, and McGrattan (2002) who focus on the aggregate real exchange rate in a one-sector model and by Kehoe and Midrigan (2007) who focus on Law-of-One-Price (LOP) deviations in a multisector model.

The empirical evidence emanating from the VAR literature is mixed on the question of the relative contribution of real and nominal shocks to real exchange rate variability. Using a structural VAR model, Clarida and Galí (1994, Table 3) find that the relative contribution of nominal shocks to the 1-period-ahead forecast error variance of the quarterly bilateral U.S. real exchange rate is 47 percent for the German mark, 36 percent for the Japanese yen, but a mere 2 percent for British pound and less than 1 percent for the Canadian dollar. At much longer horizons – 31 to 36 months – Eichenbaum and Evans (1995, Table 1a) find the contributions of money shocks to the forecast error variance of real exchange rates are 42.9, 38.1, 37.5, 26.2 and 23.0 percent, for Germany, Italy, France, the UK and Japan, respectively (see also Rogers, 1999).

An early advocate for the role of real shocks in the equilibrium determination of real exchange rates is Stockman (1980). Stockman casts his model in a flexible price setting, so that nominal shocks make no contribution to real exchange rate volatility. Crucini, Shintani, and Tsuruga (2010), on the other hand, neutralize the effect of nominal shocks by focusing on intranational trade and investigate the role of real shocks on good-level real exchange rate volatility across cities in the presence of local currency price rigidity. Unlike models emphasizing the role of the nominal shocks, their model predicts a *negative* correlation between price stickiness and real exchange rate variability because only *real* shocks affect real exchange rates across locations *within* a country.

The current paper puts these two views of real exchange rate determination on the same playing field by combining the model of Kehoe and Midrigan (2007) which emphasizes money shocks and nominal exchange rates with the model of Crucini, Shintani, and Tsuruga (2010) which emphasizes productivity shocks and trade costs. These models rely on the time dependent pricing assumption of Calvo, but allow the frequencies of price changes to vary across goods, as measured in the micro-data. Under the synthesized framework, we theoretically explore the *cross-sectional* relationship between price stickiness and real exchange rate volatility at the level of individual goods. We refer to this relationship as the real exchange rate volatility curve: the functional relationship between the forecast error variance of the real exchange rate and the infrequency of price changes at the level of a good. When real shocks are absent, the volatility curve is upward-sloping: an increasing function of the price stickiness parameter and the good with the least flexible price should exhibit the greatest amount of real exchange rate variability. When nominal shocks are turned off, the volatility curve is downward-sloping: a decreasing function of the price stickiness parameter and the good with the most flexible price has the greatest amount of real exchange rate variability. When both real and nominal shocks are present, the real exchange rate volatility curve becomes U-shaped and could result in a zero unconditional correlation between real exchange rate volatility and the frequency of price adjustment.

We estimate the volatility curve using sector-level real exchanges of Austria, Belgium, France, and Spain vis à vis the US, constructed by Kehoe and Midrigan (2007). We find that the shape of the estimated curve is consistent with the predominance of real shocks over nominal shocks, in the sense that the curve is minimized at the level of price stickiness which corresponds to a more than 4 month duration between price changes. Nonparametric analysis suggests the convexity of the curve as well as a negative relationship between price rigidity and real exchange rate volatility over the vast majority of empirical frequencies of price change found in the cross-section of goods. The negative correlation together with the theoretical prediction of our model suggests that sector-specific real shocks explain the bulk of short-run volatility in real exchange rates. We further conduct variance decompositions of sector-level real exchange rates to evaluate the relative contribution of nominal and real shocks at various horizons. For almost all goods, the contribution of nominal shocks are smaller than that of real shocks, and real shocks rise in dominance as the forecast horizon lengthens.

Our findings on the role of sector-specific real shocks are consistent with recent micro evidence by Bergin, Glick and Wu (2009), but in stark contrast with the traditional view that aggregate real exchange rate variability is attributable mostly to nominal shocks (Rogoff, 1996). To reconcile the microeconomic evidence with the macroeconomic evidence, it seems necessary to allow for large idiosyncratic real shocks at the sector-level such that these microeconomic sources of variation average out in the move to the CPI-based real exchange rate.

2. The Model

The theory combines the key features of Kehoe and Midrigan (2007) and Crucini, Shintani and Tsuruga (2010).[1] Both of these models assume heterogeneous price stickiness across goods, but the former relies on nominal exchange rate variations whereas the latter focuses on the labor productivity variations in explaining the real exchange rate volatility at sector level. In what follows, the core implications of the model are discussed focusing on the cross-sectional volatility of the (log) real exchange rate for a bilateral pair of countries, defined as:

$$q_{it} = s_t + p_{it}^* - p_{it}, \tag{2.1}$$

where p_{it} (p_{it}^*) denotes the (log) sectoral price index in the home (foreign) country and s_t is the (log) nominal exchange rate. Throughout the paper, variables marked with an asterisk denote foreign analogs of home variables.

For ease of exposition, some simplifying assumptions are made on the sources of real exchange rate variation. The first assumptions concern nominal shocks and exchange rates. The nominal shocks in the model are the home and foreign money growth rate, μ_t and μ_t^*, which are independent and identically distributed (i.i.d.). The semi-log household preferences over consumption and leisure, combined with a local-currency cash-in-advance constraint, leads to the equality of the money growth rate differential and the nominal exchange rate growth rate (i.e., $\mu_t - \mu_t^* = \Delta s_t$).[2] These assumptions are taken from Kehoe and Midrigan

[1] The full model is presented in the technical appendix of this paper, which is available from the authors upon request.

[2] To be specific, semi-log period utility is given by $\ln C_t - \chi L_t$, where C_t, L_t and $\chi (> 0)$, denote aggregate consumption, hours worked, and marginal disutility of labor supply, respectively.

(2007) and are convenient since the nominal exchange rate s_t becomes a random walk, consistent with the seminal paper of Meese and Rogoff (1983).

The second set of assumptions concern real shocks and trade costs. Monopolistically competitive firms set prices of their goods, which are produced using a technology that is linear in labor and subject to productivity shocks. Due to our microeconomic focus, the productivity shock for each good consists of three components: a global component, a nation-specific component and a good-specific component. To be precise, the productivity (in logs) for production of good i at time t, at home a_{it}, and in the foreign country, a^*_{it}, are given by:

$$a_{it} = z_t + \eta_t + \varepsilon_{it}, \qquad a^*_{it} = z_t + \eta^*_t + \varepsilon^*_{it}, \tag{2.2}$$

respectively. The three components, then, are: (i) z_t, a common global stochastic trend, following $z_t = z_{t-1} + \varepsilon_{zt}$, ε_{zt} ~i.i.d.$(0, \sigma_z^2)$, (ii) η_t ~i.i.d.$(0, \sigma_\eta^2)$ and η^*_t ~i.i.d.$(0, \sigma_\eta^{*2})$, capture country-specific productivity shocks common to all goods, and (iii) ε_{it} ~i.i.d.$(0, \sigma_\varepsilon^2)$ and ε^*_{it} ~i.i.d.$(0, \sigma_\varepsilon^{*2})$ are idiosyncratic shocks to the production of each good in each country.[3] These labor productivity shocks are the real shocks in the model. Finally, firms in each country are required to pay an iceberg transportation cost τ to send good across the border. This transportation cost leads to home bias in consumption because the home variety of each good is cheaper than the imported variety.

The focal equation of the model is the k-period-ahead forecast error variance of the sector-level real exchange rate:

$$Var_{t-k}(q_{it}) = \Lambda_{ik}[\lambda_i^2 Var(\mu_t - \mu^*_t) + (1-\lambda_i)^2(1-\lambda_i\beta)^2\psi^2 Var(a_{it} - a^*_{it})] \tag{2.3}$$

where $\Lambda_{ik} = \sum_{j=1}^{k} \lambda_i^{2(j-1)}$, λ_i is the good-specific degree of price stickiness in the sense of Calvo (1983) and the parameter, $\psi = \left(1 - (1+\tau)^{1-\theta}\right) / \left(1 + (1+\tau)^{1-\theta}\right)$, captures the asymmetric transmission of productivity to the relative price of good i, across countries due

[3]A common stochastic trend z_t among two countries can be specific to the sector i, $z_{it} = z_{it-1} + \varepsilon_{izt}$, without changing the substance of our results. Also, for simplicity, we assume that σ_ε^2 and σ_ε^{*2} are common across i. In the empirical part of the paper, the possiblity of heterogenous variances of real shocks across sectors is considered.

to the home bias in expenditure on home and foreign varieties of good i. The veracity of this asymmetry depends positively on the trade cost, τ, and the elasticity of substitution among differentiated products, θ.

Equation (2.3) attributes the forecast error variance of the sectoral real exchange rate to the variance of the money growth differential, $\mu_t - \mu_t^*$ (the nominal shocks) and the variance of the cross-country productivity differential, $a_{it} - a_{it}^*$ (the real shocks).[4] The first term of Equation (2.3) will be referred to as the nominal effect and the second term as the real effect, alluding to the effects of nominal and real shocks on the volatility of the sectoral real exchange rate.

Under time dependent pricing, monopolistically competitive firms cannot change prices with probability λ_i which is assumed to be common across countries but differs across goods.[5] That is, some sectors are allowed to change their prices more frequently than others. Viewed as a function of the sector-specific degree of price stickiness, $\lambda_i \in [0, 1]$, equation (2.3) is called: *the real exchange rate volatility curve.* Note that $\lambda_i^2 \Lambda_{ik}$ is increasing in λ_i, while $(1 - \lambda_i)^2(1 - \lambda_i\beta)^2\Lambda_{ik}$ is decreasing in λ_i. Fixing the forecast horizon k, an increase in λ_i ('stickier prices') increases the contribution of the nominal effect and decreases the contribution of the real effect to the total forecast error variance of the sectoral real exchange rate.

To gain some intuition for how price stickiness amplifies the impact of nominal shocks or mitigates the impact of real shocks on the real exchange rate volatility represented by (2.3), recall the definition of the real exchange rate given in equation (2.1). To see the impact of nominal shocks in (2.1), consider a positive money growth rate shock in the home country, holding fixed foreign money growth. The model predicts an immediate depreciation of the nominal exchange rate, that is, an increase in s_t in (2.1). The responses of local currency

[4]Note that we allow shocks to the money supply and labor productivity to have permanent effects on the levels of prices and outputs which appear to fit the data. However, the relative price (real exchange rate) is stationary as a consequence of equilibrium adjustment of the nominal exchange rate with respect to nominal shocks and as a consequence of our cointegration restriction on the productivity across countries.

[5]Much of the existing empirical work on the topic has emphasized the dominance of good-dependent frequencies over location-dependent frequencies. However, when inflation rates and/or exchange rate properties differ substantially across bilateral pairs, this may change.

prices, though, depend on the good-specific frequencies of price adjustment. For goods with prices that change every period, their local currency price adjustment completely offsets the impact of the nominal exchange rate depreciation and preserves the original LOP deviation. At the other end of the continuum, goods with prices that are extremely sticky will have real exchange rates that basically follow the path of the nominal exchange rate with negligible pass-through of the nominal shock to local currency prices. Simply put: the nominal effect on real exchange rate variability is amplified by slow local currency price adjustment.

Turning to real shocks, consider a positive shock in home productivity in sector i. Since this productivity shock is isolated to a single sector, it is assumed to have no equilibrium consequences for the nominal exchange rate, s_t in (2.1). What it does is reduce both home and foreign price indexes of sector i, because firms in the home country sell the goods produced in this sector in both countries. However, due to home bias generated by trade costs, the home sectoral price index will decrease more than the foreign sectoral price index which increases the price index differential $p_{it}^* - p_{it}$ in (2.1). Because this economic channel requires prices to actually change and thereby induce asymmetric price changes across locations, it is more quantitatively important when prices are relatively flexible. Conversely, the real effect is mitigated by slow local currency price adjustment. This discussion should make clear that the conventional wisdom of a positive relationship between real exchange rate volatility and price stickiness is predicated on the assumption that the nominal effect dominates the real effect. The next section provides some numerical examples of how different intensities of real and nominal shocks alter the shape of the real exchange rate volatility curve.

3. Numerical Examples

This section uses numerical examples to show how the shape of the real exchange rate volatility curve, as function of $\lambda_i \in [0, 1]$, depends upon the relative volatility of real and nominal shocks and a few key structural parameters.

We focus on the one-period-ahead forecast error variance by setting $k = 1$ in equation

(2.3),

$$Var_{t-1}(q_{it}) = \lambda_i^2 Var(\mu_t - \mu_t^*) + (1-\lambda_i)^2(1-\lambda_i\beta)^2\psi^2 Var(a_{it} - a_{it}^*). \tag{3.1}$$

and make note of the fact that the k period ahead forecast is proportional to the one-period ahead forecast: $Var_{t-k}(q_{it}) = \Lambda_{ik} \cdot Var_{t-1}(q_{it})$ for any k.

The structural parameters are calibrated as follows: i) the data is monthly, so the discount factor is set to $\beta = 0.96^{1/12} = 0.9966$; ii) trade costs, broadly defined at the retail level, are in the neighborhood of $\tau = 0.5$; and iii) the elasticity of substitution is set at $\theta = 10$. The multiplier on the productivity differential, reflecting home bias, then, is $\psi^2 = 0.9$. While trade costs and elasticities of substitution are expected to differ across goods, the salient features of the volatility curve are not very sensitive to these parameters, leaving the interesting dimension as the interaction of the frequency of price adjustment λ_i and the variances of nominal and real shocks.

For purposes of discussion, the expression is simplified by noting that since the discount factor and the trade-bias factor are close to 1, the variance decomposition of the real exchange rate is well approximated by:

$$Var_{t-1}(q_{it}) = \lambda_i^2 Var(\mu_t - \mu_t^*) + (1-\lambda_i)^4 Var(a_{it} - a_{it}^*)\ . \tag{3.2}$$

Simply put, the variance of the sectoral real exchange rate is a weighted average of nominal and real shocks. The share of the nominal shock (the money growth differential) in real exchange rate variability rises from zero toward 1 as prices become less flexible ($\lambda_i \to 1$) at a rate equal to the square of the infrequency of price changes. In contrast, the share of the contribution of the real shock (the productivity differential) toward real exchange rate variability rises from zero toward 1 as prices become more flexible ($\lambda_i \to 0$) at a rate equal to 4th power of the frequency of price changes $(1-\lambda_i)$. The theory conveniently encompasses the varied role of nominal and real shocks in the cross-section of goods and the feature that these two opposing forces give rise to a real exchange rate volatility curve that is U-shaped over the support $\lambda_i \in [0, 1]$.[6]

[6] To prove this, evaluate the first derivative of the total variance with respect to λ_i at $\lambda_i = 0$ and 1.

Turning to numerical examples, Figure 1 uses equation (3.1) to construct conditional variances as a function of λ_i separately for three distinct stochastic environments (each with the $Std(\mu_t - \mu_t^*)$ normalized to 1 percent): (a) $Std(a_{it} - a_{it}^*)/Std(\mu_t - \mu_t^*) = 5$; (b) $Std(a_{it} - a_{it}^*)/Std(\mu_t - \mu_t^*) = 1$; and (c) $Std(a_{it} - a_{it}^*)/Std(\mu_t - \mu_t^*) = 1/5$. Each panel shows the variance due to the real effect (in blue) and variance due to the nominal effect (in red). When the distribution of goods across frequencies of price change is uniform on the interval $[0, 1]$, these areas fully summarize the role of real and nominal shocks across the entire cross-section of goods.

The extent to which the role of real shocks is more quickly mitigated by price rigidity than is the role of nominal shocks by price flexibility is clearly evident in the middle panel where the two shocks have equal variance. The critical good for which real and nominal shocks contribute equally to real exchange rate variability has a frequency of price change, λ_i, of about 0.4 (rather than 0.5, the midpoint of the unit interval), indicating more price flexibility is needed to handicap the advantage given to nominal shocks by the theory. However, as evident in panels (a)-(c), the overall size of each area is sensitive to the relative variance of real and nominal shocks.

The asymmetries across goods are, of course, greatest at the extremes of price flexibility and inflexibility. Consider fixing a particular good, indexed by its frequency of price change, λ_i. The height of the volatility curve is the model's prediction for the total one-step ahead forecast error variance of the real exchange rate for that good. The blue and red segments of the vertical line drawn from that point in the λ distribution up to the volatility curve gives the partition of the variance into the contributions of real and nominal shocks. At the extremes, real exchange rate fluctuations of goods with fully flexible price are driven solely by real shocks while goods with completely rigid prices are driven solely by nominal shocks.

When evaluated at $\lambda_i = 0$, the first derivative of the variance due to the nominal effect is zero but that due to the real effect is negative and finite, which implies that the first derivative of the total variance with respect to λ_i is strictly negative when $\lambda_i = 0$. Analogously, we can also show that the total variance has a strictly positive slope at $\lambda_i = 1$. Because total variance is continuous in λ_i, there exists $\lambda_i \in (0, 1)$ that minimizes total variance.

Perhaps the closest real-world counterparts of such goods are crude petroleum trading on the centralized spot market and postage stamps.

Panel (a) calibrates the model such that only real shocks are of consequence and the prediction of the model is that the volatility curve is downward sloping over almost its entire range. Only as completely rigid local currency prices are approached at $\lambda_i = 1$ do we see the curve begin to slope upward, providing the first hint of a nominal effect (i.e., the red area becomes visible). The good with the lowest real exchange rate variance in this calibration – the point at which the sign of the slope changes – is $\lambda_i = 0.75$. This corresponds to an average duration between price changes of 4 months. The average contribution of real and nominal shocks to the variance of sectoral real exchange rate may be evaluated by integrating the real exchange rate volatility curve over the entire range of λ_i. The blue area, which corresponds to real effects, accounts for 93% of the total area, the red area (nominal effects) contributes a mere 7%.

Panel (c) calibrates the model such that only nominal shocks are of consequence and the prediction of the model is that the volatility curve is upward sloping over almost its entire range. Only as we approach goods with completely flexible prices do we see the curve begin to slope downward, providing evidence of real effects (i.e., the blue area becomes visible). The point which minimizes variance, and therefore the point at which the sign of the slope changes is virtually indistinguishable from complete price flexibility, $\lambda_i = 0.06$. This corresponds to an average duration between price changes of 1 month. The red area, the nominal effect, accounts for about 98% of the total area, the blue area (real effect), a trivial 2%.

The middle panel (b), with equal variances of real and nominal shocks, displays an obvious U-shape. The blue area is comparable in size to the red area, with the real effect accounting for 35% of the variation and nominal effects accounting for the remaining 65% (these are averages across goods). The minimum point on the exchange rate volatility curve is $\lambda_i = 0.40$. This corresponds to an average duration between price changes of 1.7 months.

Consider, now, what we would expect to find in terms of the correlation of price stickiness, as measured by the Calvo parameter, λ_i, and the volatility of the sector-level real exchange rate. Computing the correlation of frequency of price changes and real exchange rate variability will yield a negative correlation in panel (a) because real shocks dominate; a positive correlation in panel (c), because nominal shocks dominate; and a near zero correlation in panel (b) due to the U-shape of the volatility curve. This is a textbook example of a situation in which correlation, a linear measure of dependence, fails to reveal the true underlying economic structure. The model gives a compelling rationale for investigating a non-linear relationship between real exchange rate variability and the frequency of price adjustment.

In practice the sign of the correlation will depend on the distribution of goods in the sample (in terms of their frequencies of price adjustment) as well as the relative importance of real and nominal shocks. Consider a researcher with a limited sample of goods with prices which happen to be very sticky in local currency units. Even in the case of panel (b), this researcher will find positive correlation rather than a zero correlation, and thus will be tempted to conclude that nominal shocks predominate. For the goods in his sample, this is true, but it need not be true in general, over the entire distribution of λ_i across goods in the economy.

Note that the blue and red areas in Figure 1 represent the cross-sectional average of the sector-level variance decomposition.[7] However, individual goods may have vastly different variance decompositions. This arises from the fact that the weights λ_i^2 and $(1-\lambda_i)^2(1-\lambda_i\beta)^2\psi^2$ appearing in (3.1) are very sensitive to the sector-specific value of λ_i. Recall that the blue and red areas in panel (b) account for 35% and 65% (respectively) of the variance of real exchange rates over the entire unit interval. Consider two goods, one with $\lambda_1 = 0.25$ and the other with $\lambda_2 = 0.56$. These two real exchange rates turn out to have the same total forecast error variance as panel (b). However, due to the manner in which shocks are transmitted to

[7] It corresponds to the average when the degree of price stickiness is uniformly distributed across goods.

relative prices, the sources of the variance differs dramatically across the them: the relative contribution of nominal shocks is only 18% for the first good but 90% for the second good.

To summarize our numerical analysis, there are two striking empirical implications of our model when both real and nominal shocks are allowed to impinge on the economy. First, it is possible that the data suggest a negative correlation between total real exchange rate volatility and degree of prices stickiness, contrary to the conventional wisdom. As panels (a) and (b) in Figure 1 suggest, negative correlations can occur when the variance of productivity differentials dominate the economic environment for most goods in the cross-section. Second, since the real exchange rate volatility curve is a U-shaped curve, running a simple linear regression of the sectoral real exchange rate variance on λ_i may be a poor method for uncovering the underlying structure. It is useful to consider a flexible functional form in the regression and find the degree of price stickiness which minimizes the volatility curve.

4. Empirical Analysis

The empirical analysis focuses on (i) examining the relationship between total variance and the degree of price stickiness; (ii) finding the degree of price stickiness which minimizes the volatility curve; and (iii) assessing the relative importance of the real and nominal effects at the sectoral level. The data used here was originally obtained by Kehoe and Midrigan (2007) and consist of highly disaggregated sectoral real exchange rates for four European countries (Austria, Belgium, France, and Spain) relative to the U.S.. Essentially this involves matching monthly local currency micro-price data from Eurostat and the Bureau of Labor Statistics and converting to a common-currency using spot nominal exchange rates. The sample period is monthly from January 1996 until December 2006. The number of sectors is 66. Kehoe and Midrigan (2007) take the cross-country average monthly infrequencies of price changes within each sector. The country-specific frequencies for the U.S. are from Bils and Klenow (2004) and those for each of the European countries are taken from the individual country studies by: Baumgartner, Glatzer, Rumler, and Stiglbauer (2005) for Austria; Aucremanne and Dhyne (2004) for Belgium; Baudry, Le Bihan, Sevestre, and Tarrieu (2007) for France; Alvarez and Hernando (2004) for Spain. The details of the data construction are found in the appendix of Kehoe and Midrigan (2007).

While the Euro was officially introduced part-way through our sample, even before this the nominal exchange rate of these European countries were quite stable against each other. This is evident in the standard deviations of the nominal exchange rate growth of the U.S. dollar against Austrian Schillings, Belgian Francs, French Francs, and Spanish Pesetas, which are 2.36, 2.37, 2.35, and 2.36 percent, respectively. The similarity of nominal exchange rate volatility, effectively the nominal shocks of the theory, rationalize a pooled regression of the four country-pairs against the dollar as the benchmark in the analysis below. However, the productivity differentials may differ across bilateral pairs, so we estimate the relationship for each country separately as a robustness check.

4.1. Estimating the real exchange rate volatility curve

Let V_{ij} be the one-period-ahead forecast error variance of the real exchange rate for good i for country j, vis à vis the United States. The technical appendix of the paper proves that q_{ijt} follows an AR(1) process with an AR coefficient λ_{ij} under a set of maintained assumptions. Effectively, this means V_{ij} is equal to the sample variance of $q_{ijt} - \lambda_{ij} q_{ijt-1}$ using the observed infrequency of price changes, λ_i. When either λ_i or q_{ijt} is missing or when V_{ij} can be computed from only a short time sample, we exclude such goods from the sample.[8]

As a preliminary analysis, the first set of regression results are simple linear regressions of V_{ij} on λ_{ij} using (i) the pooled samples of all four country-pairs; and (ii) country-by-country samples. The results are reported in Table 1. In all cases, the sign of the coefficient on λ_{ij} is significantly negative based on heteroskedasticity consistent standard errors reported

[8] After excluding the samples, the number of sectors amounts to 57 for Austria, 46 for Belgium, 48 for France, and 31 for Spain.

below the point estimates. Despite its simplicity, this specification, explains 70 percent of the cross-sectional variation in the volatility of real exchange rates using the pooled regression.[9] The estimated slope coefficients are similar across the four nation-specific regressions and the cross-country pooled regression. Regression fit is especially good for Austria and France.

Table 1 also reports the minimum and maximum values of price stickiness in our data (denoted $\lambda_{\min}$ and $\lambda_{\max}$). The linear regression estimates suggest that the real exchange rate volatility curve is downward sloping within the range of the observed λ_{ij}. According to the theory, the negative correlation is consistent with the dominance of productivity differentials over money growth rate differentials, consistent with the stylized numerical example presented in panel (a) of Figure 1.

Recall, however, that the theory predicts a non-linear relationship between the frequency of price adjustment and real exchange rate variability when both nominal and real shocks are present.[10] To more adequately address this implication of the theory, the volatility curve is augmented with a quadratic term and a quartic term following the structural model:

$$V_{ij} = b_{1j}\lambda_{ij}^2 + b_{2j}(1-\lambda_{ij})^2(1-\lambda_{ij}\beta)^2 + u_{ij}, \tag{4.1}$$

where the b's are regression coefficients and u_{ij} is the regression error term for good i for country j. The second regressor is constructed by setting $\beta = 0.96^{1/12}$. According to (3.1), regression coefficient b_{1j} should capture the nominal effects, due to $Var(\mu_t - \mu_{jt}^*)$, where the money growth rate has been replaced by the variance of the bilateral nominal exchange rate to anticipate the empirical implementation that follows. The regression coefficient, b_{2j}, captures the real effects $\psi_j^2 Var(a_{it} - a_{ijt}^*)$, with the restriction that the variance of productivity differentials are common across i. Note that since the empirical work involves more than one bilateral pair (as assumed in the theory), the trade costs, demand elasticities

[9] For robustness, we also run the pooled regression with the country dummies for both intercept and slope coefficient to control for differences in trade costs paid to carry goods from a country to another country. The results are essentially unchanged.

[10] Using Ramsey's (1969) RESET test, the null hypothesis of linearity is rejected at the one percent significance level for the pooled case, as well as for the Austrian and French cases. Weaker evidence of nonlinearity is obtained for Belgian and Spanish cases possibly because the power of the test is lower in smaller samples.

and productivity shocks are allowed to vary for across bilateral pairs for each good. For example, productivity variation may be higher across Spain and the United States (the numeraire) than between France and the United States. With this nonlinear extension, it is also possible to locate the degree of price stickiness which minimizes the volatility curve using the estimates of b_{1j} and b_{2j}.

Table 2 presents the estimation results of regression model (4.1). In all cases, the estimated coefficients have positive signs which is consistent with the theory, all are statistically significant. The quartic regression is comparable to the linear regression in terms of the goodness of fit.[11] The role of nominal shocks compared to real shocks, as implied by the estimates b_{1j} and b_{2j}, indicates a much larger role for real shocks. While the regression does not separately identify the role of home bias ψ_j^2 and productivity shocks $Var(a_{it} - a_{ijt}^*)$, b_{2j} identifies their combined influence and serves as a lower bound on $Var(a_{it} - a_{ijt}^*)$ given $0 \leq \psi_j^2 \leq 1$ for $\tau_j \geq 0$ and $\theta_j \geq 1$. The same argument establishes that $\sqrt{b_{2j}/b_{1j}}$ can be used as a lower bound for $Std(a_{it} - a_{ijt}^*)/Std(\mu_t - \mu_{jt}^*)$.

Using the coefficient estimates reported in Table 2 the ratios of the standard deviation of real to nominal shocks are inferred to be: 5.28, 6.17, 4.42, 5.30 and 9.64, for the pooled case and the Austrian, Belgian, French and Spanish nation-specific cases, respectively. Thus, the quartic regression result is very comparable to the numerical example of case (a) where $Std(a_{it} - a_{it}^*)/Std(\mu_t - \mu_t^*) = 5$. This is further confirmed in the left panel of Figure 2 showing the fitted curve of the pooled quartic regression (the solid line) from Table 2, along with that of the pooled linear regression (the dashed line) from Table 1. The fitted curve of the quartic regression resembles panel (a) of Figure 1 in terms of the shape of the curve, again suggesting the importance of real effects.

The last two columns of Table 2 compare the estimated degree of price stickiness which minimizes the forecast error variance in each regression. For the pooled case, the variance of the real exchange rate is minimized at $\lambda_i = 0.76$ which is remarkably close to the value of

[11] The inclusions of country dummies into the pooled regression again did not affect the signs and magnitudes of the estimated coefficients.

0.75 from our numerical example with dominant real shocks. This frequency of price change implies that the U-shaped real exchange rate volatility curve is minimized when the duration between price changes is 4.2 months.

Note that the parametric regression (4.1) imposes a strict theoretical shape restriction on the real exchange rate volatility curve. As a robustness check, the functional form restriction is replaced with a general nonparametric regression,

$$V_{ij} = m_j(\lambda_{ij}) + u_{ij}$$

where $m_j(\cdot)$ is an unknown conditional mean function for country j. The right panel of Figure 2 shows the estimated curve using the nonparametric local linear regression estimator with pooled data.[12] The shape of the fitted curve shown as the solid line is very different from the linear regression fit shown as the dashed line. This suggests the plausibility of a nonlinear structure in the real exchange volatility curve.

Turning to a comparison of the quartic regression (4.1) and the non-parametric regression, both similarities and differences are evident. Both estimates imply convexity in the real exchange rate volatility curve. When the first derivative of the m function is evaluated nonparametrically, it tends to be increasing in λ_{ij}, which is consistent with the theoretical prediction. The slope of the curve is negative over the empirical range of λ_{ij} and it becomes flatter as λ_{ij} increases. The most notable difference between the quartic regression and nonparametric regression is the location of the bottom of the curve. The value of λ_{ij} which minimizes the forecast error variance in the nonparametric regression is close to unity, a value larger than the theoretical prediction based on the quartic regression.

To formally investigate the shape of the estimated curve, a nonparametric test of monotonicity developed by Ghosal, Sen and van der Vaart (2000) is employed – a test of the null hypothesis that the m function is an increasing (or decreasing) function over a certain interval. In the present context, the shape of the curve is examined over the observed range of the data, $[\lambda_{\min}, \lambda_{\max}]$. The test is also applied to establish the monotonicity of the first derivative

[12] In estimation, Gaussian kernel is used along with the bandwidth selected by the rule of thumb.

of the m function. The results are reported in Table 3.

Regarding the m function itself, the hypothesis of an increasing function in λ_{ij} is rejected, and that of decreasing function is not, based on a conventional significance level. For the first derivative, the test fails to reject a monotonically increasing function while a monotonically decreasing function is rejected. Simply put: the real exchange rate volatility curve is a convex function consistent with the U-shape prediction of the theory.

Establishing the inflection point on the real exchange rate volatility curve is tenuous, it is likely to be associated with a λ_{ij} larger than estimated from the quartic regression, though the structurally restricted estimate of 0.76 (a duration of about 4 months between price adjustments) is our preferred choice. Since the minimum point is largely a function of the stochastic environment and not the underlying structural parameters, the minimum need not be a crucial focus. However, it would be reassuring in terms of validating the generality of the theory to explore other samples of goods, cross-sections of countries and historical periods such that the nominal shocks play a larger role and the U-shape not be truncated at the upper boundary of price rigidity.

4.2. Variance decomposition

Let us now turn to the relative importance of the real and nominal effects at sector level by directly using equation (2.3) at various horizons along with an empirical measure of the variance of the nominal shock. According to the theory, the appropriate metric of $\mu_t - \mu^*_{jt}$ is Δs_{jt} and the nominal contribution to real exchange rate variance is $\lambda^2_{ij}\Lambda_{ijk}Var(\Delta s_{jt})$. The k-period-ahead forecast error variance, $Var_{t-k}(q_{ijt})$, is obtained from the quasi-difference $q_{ijt} - \lambda^k_{ij} q_{ijt-k}$, using observed sectoral infrequency of price changes, λ_{ij}.

The relative contribution of nominal shocks to the total forecast error variance of the real exchange is:

$$\phi(i,j,k) = \frac{\lambda^2_{ij}\Lambda_{ijk}Var(\Delta s_{jt})}{Var_{t-k}(q_{ijt})}$$

where the indices of the share function, $\phi(i,j,k)$, reflect the role of goods (infrequency of price changes), country (due to the variability of the bilateral nominal exchange rate of country j vis à vis the U.S. dollar) and horizon. As $k \to \infty$, the sample variance of q_{ijt} is used for the unconditional variance and $[\lambda_{ij}^2/(1-\lambda_{ij}^2)]Var(\Delta s_{jt})$ is used to measure the contribution of the nominal shocks. In addition to its simplicity, this good-by-good variance decomposition method has the advantage that it does not require the assumption of a common volatility of real shocks across sectors, since $1-\phi(i,j,k)$ is unrestricted and may vary due to heterogeneous variance of real shocks across sectors.

Table 4 reports the summary statistics for the contributions of the nominal shocks to the forecast error variance of sectoral real exchange rates at monthly horizons of $k = 1, 3, 6, 12$ and ∞. Note that unlike the variance decomposition of aggregate real exchange rates often reported in the literature, the decomposition is calculated for each sectoral good. The first row of the table shows the average contribution of nominal shocks, with the average taken across all goods and all four bilateral pairs. The numbers in parentheses in the second row are the standard deviations across goods and countries. The remaining rows report corresponding results for each pair of countries.

For the one-period-ahead forecast error decomposition, nominal shocks account for about 40 percent of real exchange rate variation on average and range from a high of 49 percent for Austria to a low of 35 percent for Spain. The large standard deviations in the table imply that the contributions of nominal shocks differ considerably across goods. This cross-sectional dispersion is very similar across countries. For the shortest horizon it seems safe to conclude that the contribution of real shocks is at least as large as that of nominal shocks for many goods.

The role of nominal shocks becomes smaller as the horizon lengthens. At a horizon of 6 months, the relative contribution is about one half of the 1-month horizon. The long-run contribution of nominal shocks, evaluated at $k = \infty$, is lower than 10 percent for all countries except for Austria, thus leaving 90 percent to be explained by real shocks. The cross-sectional variation (across goods) at the longest horizon is much smaller than that at shorter horizons, implying the dominance of real shocks in real exchange rate fluctuations for most goods. This finding of a dominant role of sector-specific real shocks is consistent with recent micro evidence by Bergin, Glick and Wu (2009) who claim that idiosyncratic industry price shocks account for about 80% of variation in LOP deviations in disaggregated data, with nominal exchange rate shocks playing a very small role.

Let us now compare the variance decompositions of sector-level real exchange rates with previous studies involving the aggregate real exchange rate. Using a structural VAR model, Clarida and Galí (1994, Table 3) find that the relative contribution of nominal shocks to 1-period-ahead forecast error variance of quarterly real exchange rate is 47 percent for Germany and 36 percent for Japan. In contrast, our three-month (the counterpart to one quarter) ahead variance decomposition indicates nominal shocks account for between 19 to 31 percent, depending on the country, when results are averaged across sectors (see Table 4). Using over 100 years of annual UK-US real exchange rate data, Rogers (1999) finds that the contribution of nominal shocks to the 1 year-ahead forecast error variance ranges from 19 percent to 60 percent, with a median value of 41 percent. Our 12-month ahead forecast variance decomposition estimates indicate nominal shocks only account for about 14 percent when we average across countries and sectors. The benchmark estimates of Eichenbaum and Evans (1995, Table 1a) show a nominal shock contribution at horizons of 31- to 36-months averaging 38 percent for France, while our estimates imply long-run contributions between 9 and 12 percent for France (again, using averages across sectors). Thus, largely independent of the horizon or countries examined, nominal shocks play a more important role in accounting for aggregate real exchange rate fluctuations than in accounting for sector-level real exchange rate fluctuations.

What accounts for this difference in the microeconomic and macroeconomic evidence? Our suspicion is that the sectoral real effects tend to average out across sectors while the nominal effects, almost by definition cannot, since there is only one nominal exchange rate per bilateral pair. When researchers use aggregate level CPI-based real exchange rates, the impact of real shocks is attenuated by the aggregation process, while the impact of nominal shocks is not (since the nominal exchange rate shock is common to all goods, up to heterogenous price adjustment rates). When using less aggregated data, it is therefore perhaps not surprising that nominal and real shocks are more on par as contributors to real exchange rate variation.

In terms of the theory, recall productivity shocks found in equation (2.2) are expected to embody idiosyncratic sector-specific shocks ε_{it} and ε_{it}^{*}, at least in part. The aggregation of goods prices across sectors eliminates this idiosyncratic component and one is left with a national productivity differential shock and the money growth differential shock to account for the aggregate real exchange rate variance. This averaging-out argument is consistent with a recent finding by Crucini and Telmer (2007) who show that only a small fraction of LOP changes are common to all goods. Assigning all of the common component to nominal exchange rates in the presence of sticky-prices would, thus, leave most of the LOP variation unaccounted for. The decomposition performed here fills this gap with real shocks.

5. Conclusion

We use a time-dependent Calvo pricing model with real and nominal shocks to develop the concept of a real exchange rate volatility curve. The curve was proved to have a U-shape as a function of the degree of price stickiness, implying an ambiguous correlation between the forecast error variance of real exchange rates and price stickiness. Using US-European real exchange rate data, the correlation between the forecast error variance and the degree of price stickiness was found to be negative over most of the range of the micro-data. The good with minimal real exchange rate volatility was estimated to have a duration between the price changes of about 4.2 months. The downward sloping profile suggested that for this micro-sample of goods and countries, the variance of sectoral real exchange rates is dominated by real shocks, though nominal shocks are important as well.

These results point to the value of examining cross-sectional differences in real exchange rate variability in order to flesh out the rich quantitative predictions of models of micro-price adjustment currently under development. Differences across goods help us to disentangle heterogeneous responses to common shocks due to differences in economic propagation mechanisms such as costs of price adjustment and trade costs from heterogeneity in the underlying shocks themselves. Averaging across goods, as is inevitable in the move to an aggregate real exchange rate, is not innocuous in terms of the weight given to real and nominal shocks. The same averaging may also lead to an under-appreciation of the sources of the risks that individuals and firms face. We hope to explore these possibilities in future work. Much remains to be done.

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Table 1: Linear regressions

	Const	λ_i	Adj. R^2	Obs.	$[\lambda_{\min}, \lambda_{\max}]$
Pooled	0.013	-0.014	0.700	182	[0.223, 0.979]
	(0.001)	(0.001)			-
Austria	0.016	-0.016	0.886	57	[0.223, 0.979]
	(0.002)	(0.002)			-
Belgium	0.011	-0.011	0.454	46	[0.296,0.956]
	(0.001)	(0.001)			-
France	0.013	-0.014	0.833	48	[0.254, 0.958]
	(0.002)	(0.002)			-
Spain	0.014	-0.015	0.589	31	[0.524, 0.964]
	(0.003)	(0.003)			-

NOTES: The heteroskedasticity-consistent standard errors are in parentheses. The "Adj. R^2" denotes the adjusted R^2. The "Obs." denotes the number of observations. The last column shows the empirical range of infrequencies of price changes $[\lambda_{\min}, \lambda_{\max}]$ in our dataset.

Table 2: Structural regressions

	λ_i^2	$(1-\lambda_i)^2(1-\lambda_i\beta)^2$	Adj. R_{uc}^2	$\underline{\lambda}$
Pooled	0.0016	0.0456	0.751	0.762
	(0.0001)	(0.0039)		(0.008)
Austria	0.0012	0.0475	0.895	0.784
	(0.0001)	(0.0066)		(0.011)
Belgium	0.0021	0.0419	0.623	0.735
	(0.0004)	(0.0056)		(0.017)
France	0.0015	0.0411	0.905	0.763
	(0.0001)	(0.0040)		(0.009)
Spain	0.0017	0.1567	0.716	0.836
	(0.0002)	(0.0292)		(0.011)

NOTES:The heteroskedasticity-consistent standard errors are in parentheses. The "Adj. R_{uc}^2" denotes the adjusted uncentered R^2. $\underline{\lambda}$ presents the estimates of λ which minimize the total variance.

Table 3: Tests of monotonicity

	$m(\lambda_i)$ Null hypothesis		$m'(\lambda_i)$ Null hypothesis		Critical values		
	Increasing	Decreasing	Increasing	Decreasing	1%	5%	10%
Pooled	10.329***	-1.831	-2.324	17.014***	5.342	4.386	3.964
Austria	4.927**	-1.546	1.883	9.362***	5.530	4.488	4.027
Belgium	4.608*	-0.647	1.866	9.378***	5.737	4.609	4.111
France	6.369***	-1.212	1.397	6.675***	5.645	4.554	4.072
Spain	5.098*	-1.281	-0.025	7.731***	6.924	5.382	4.701

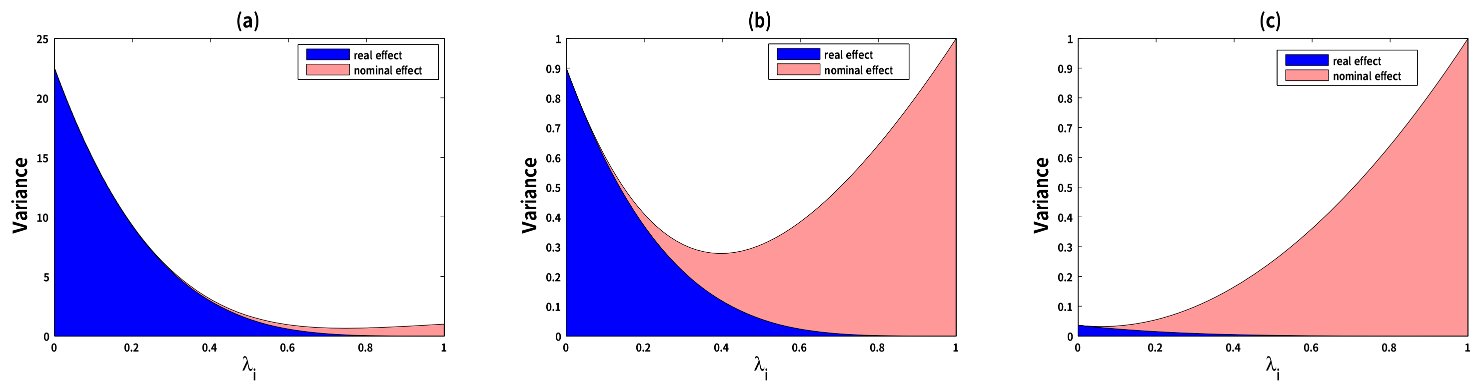
NOTES: The first two columns correspond to the hypothesis testing for $m(\lambda_i)$ and the second two columns correspond to the test for the first derivative of $m(\lambda_i)$ with respect to λ_i. Critical values shown in the last three columns are computed from the method by Ghosal, Sen, and van der Vaart (2000).

Table 4: Percentage of forecast error variance accounted for by nominal shocks

k	1	3	6	12	∞
Pooled	40.6	23.6	18.7	14.2	11.4
	(24.1)	(16.5)	(15.7)	(13.5)	(11.8)
Austria	48.6	30.5	25.7	20.3	17.1
	(24.4)	(16.9)	(17.1)	(15.7)	(16.3)
Belgium	34.9	19.9	15.3	11.4	8.9
	(23.0)	(14.9)	(13.6)	(11.2)	(8.2)
France	40.2	21.8	16.2	11.7	9.2
	(22.7)	(15.2)	(13.5)	(10.2)	(7.6)
Spain	35.2	18.9	14.6	11.1	7.9
	(24.5)	(16.4)	(15.6)	(13.3)	(8.0)

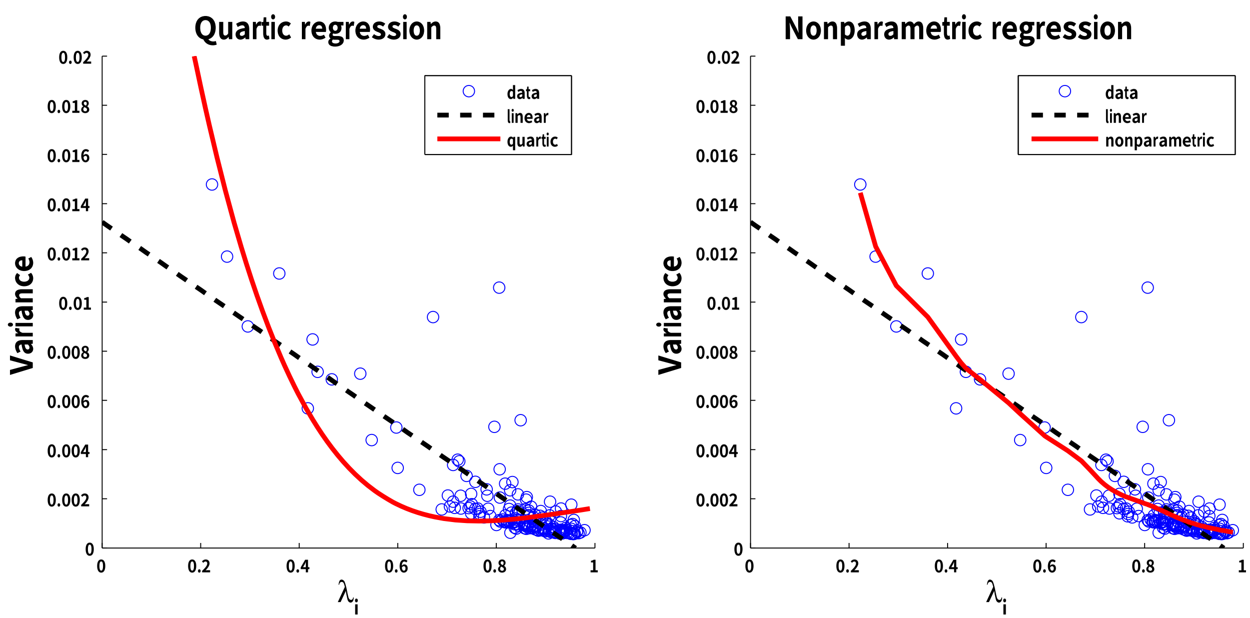
NOTES: Numbers are in percent. Each column corresponds to the cross-sectional average of the k-period-ahead forecast error variance of sector-level real exchange rates accounted for by nominal shocks. Numbers in parentheses are standard deviations.

Figure 1: Simulated real exchange rate volatility curves



NOTES: Each panel of the figure shows contributions of the nominal and real shocks on the sector-level real exchange rate volatility over the range of the degree of price stickiness. The volatility is measured by the one-period-ahead forecast error variance. For each good, the variances due to real and nominal shocks are expressed by the heights of blue and red areas. The combined height of the two areas corresponds to the total variance. Panels (a),(b) and (c) show the cases of $Std(a_{it} - a_{it}^*)/Std(\mu_t - \mu_t^*) = 5$, 1, and 1/5, respectively. Standard deviation of nominal shocks is set to one.

Figure 2: Estimated real exchange rate volatility curves



NOTES: The both panels show the scatter plot of the one-period-ahead forecast error variance of the sector-level real exchange rates against the degree of price stickiness. The solid line in the left panel shows the fitted curve based on the quartic regression. The solid line in the right panel represents the fitted curve based on the nonparametric regression. The dashed line in both panels show the linear regression fit for comparisons.