International Price Dispersion in State-Dependent Pricing Models[†]

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Abstract

Studies of disaggregated international price data document a robust, positive relationship between nominal exchange (NER) volatility and the variability of international relative prices. This relationship is interpreted as evidence that sticky prices rather than trade frictions are the source of the large law of one price deviations across locations. This paper disputes this interpretation. We show that a micro-founded, state-dependent sticky price model generates a hump-shaped, rather than positive relationship between relative price variability (RPV) and NER volatility. The hump occurs earlier for more tradable goods, whose producers adjust more frequently in a model calibrated to match the excessively large volatility of nominal exchange rates observed in the data. We find strong support for the model's predictions using a dataset of actual good prices collected in 12 Eastern European cities in a very volatile environment. In contrast, the model performs poorly when tested using a dataset of CPI-based real exchange rates for 74 countries. Lack of a hump-shaped relationship in the aggregate data is only consistent with a micro-founded sticky price model if international goods market are sufficiently segmented so that large nominal exchange rate movements have little effect on firm profits.

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1. Introduction

In an attempt to uncover the source of the large swings in the real exchange rates across countries, Engel (1993, 1999) and Rogers and Jenkins (1995) have forcefully argued that deviations from the law of one price for identical goods across countries rather than movements of relative prices within a country account for most of the volatility of real exchange rates. These authors' findings have challenged the ability of the Balassa-Samuelson hypothesis to explain short-term real exchange rate fluctuations and gave rise to a surge in empirical work using disaggregated relative price data. This line of research, pioneered by Engel and Rogers (1996) in their work with US-Canadian relative prices, has documented that national borders are an important source of market segmentation. The border between Canada and US generates as much volatility of relative prices for two cities as 75,000 miles of distance would for cities within a country[1]. The US-Japan border is even wider: 70,000 times the distance from the Earth to the Sun[2].

Tariff and non-tariff barriers, language and differences in distribution networks, as well as relative wage movements across countries have been used to explain the large impact borders have on deviations from the law of one price, but no story has received as much attention as the one of sticky prices coupled with nominal exchange rate shocks. Relative prices across countries, at all levels of aggregation, track nominal exchange rates closely, as documented by Isard (1977), Giovaninni (1988) or Mussa (1986). More recent work using bilateral disaggregated price data has documented that the volatility of nominal exchange rates has a positive effect on the variability of relative prices across countries and accounts for a large portion of the effect of national borders on international price dispersion[3]. This finding, coupled with the small effect distance is found to have on the volatility of relative prices across locations, is responsible for the popularity of nominal rather than real (trade-based) frictions as an explanation for international relative price movements.

The purpose of this paper is to ask whether the pervasive positive relationship between nominal exchange rate and relative price variability, as well as the small role tradeability is found to play in generating relative price movements is indeed consistent with a sticky price model. To answer this question we depart from standard models frequently used in open economy macroeconomics[4], models which assume an exogenous

[1] Engel and Rogers (1996).
[2] Parsley and Wei (2001a).
[3] Engel and Rogers (1996, 2001), Parsley and Wei (2001a, 2001b)
[4] See Lane (2001) for a survey of the New Keynesian Open Economy literature.

frequency and timing of price changes, and therefore predict a positive relationship between nominal exchange rate volatility and the volatility of relative prices.

In a model in which nominal rigidities arise endogenously, due to fixed (menu) costs firms pay to adjust their prices, higher nominal exchange rate volatility has two effects on the volatility of international relative prices. One is the direct effect: nominal exchange rates are used to compute relative prices and higher NER volatility increases relative price variability. On the other hand, as the volatility of nominal exchange rates increases, firms tend to adjust prices more frequently and therefore more readily correct relative price fluctuations induced by nominal shocks. Which of these two effects is stronger[5]?

To answer this question, we embed two frictions, trade costs and menu costs of price adjustment, in a small open economy model with differentiated goods and monopolistic competition in the goods market. Trade costs play an important role in the model. Firms that face small trade costs have high export shares, and allow smaller law of one price deviations because they reprice more frequently: relative price fluctuations are costlier for these firms as, given the lower costs of arbitrage, they translate into more volatile movements in quantities. In contrast to standard sticky price models, we find that the state-dependent model generates a hump-shaped relationship between the volatility of nominal exchange rates and that of relative prices. In environments in which the volatility of nominal exchange rates is small, prices adjust infrequently: relative prices inherit the properties of nominal exchange rates and higher NER volatility increases the size of deviations from the law of one price. As NER volatility increases, firms find it optimal to pay the menu costs and adjust more readily. More frequent adjustment breaks the tight link between nominal and real relative prices and higher NER volatility, through its effect on the frequency of price changes, lowers the variability of relative prices.

Alvarez, Atkeson and Kehoe (2002) have documented that high inflation countries have a lower volatility of real exchange rates relative to that of nominal exchange rates and have shown that a model with endogenously segmented markets can replicate this behavior. Moreover, Burstein (2002) shows that the effect of nominal shocks on real activity is smaller in high inflation environments as firms adjust more frequently in the presence of fixed costs of adjusting prices, and argues that a state-dependent pricing model can also

[5] Devereux and Yetman (2002) tackle a related question (the relationship between exchange rate pass-through and inflation) in the context of a time-dependent model in which firms adjust prices at random intervals, in a Calvo fashion, but choose the (ex-ante) optimal frequency of price changes given a cost of price adjustment.

reproduce the lower ratio of real to nominal exchange rate volatility in high inflation environments. Our open economy state-dependent pricing model delivers however a stronger result. According to the model, real exchange rate volatility itself should eventually decrease with higher nominal exchange rate volatility.

We employ both aggregate and disaggregated price data to test the predictions of the model. We first use a dataset of actual goods prices and show that the model's results are not simply a theoretical curiosum. The model predicts that in environments in which the frequency of price changes is high, larger NER volatility should decrease the volatility of relative prices. We turn to the market for agricultural products sold in open-air markets, a volatile environment in which prices are changed frequently, to test this prediction of the model. Using data on the prices of 58 products collected biweekly over a three year period in 12 Eastern European cities, we indeed find, using gravity-type equations, that city-pairs that are subject to larger shocks to the nominal exchange rates experience smaller deviations from the law of one price. Moreover, using period-by-period measures of relative price variability first employed by Parsley and Wei (2000) in an open-economy context, we also find that periods in which the volatility of nominal exchange rates is larger are characterized by less relative price variability, consistent with the predictions of the menu costs model.

The model fares worse however when applied to aggregate data. Using monthly CPI and nominal exchange rate data for 74 countries for the post-Bretton-Woods period, we find that the volatility of monthly changes in real exchange rates increases with larger nominal exchange rate volatility even for pairs of countries that have suffered from excessive nominal exchange rate fluctuations. Only for changes in real exchange rates at horizons of a year or more do we observe a hump-shaped relationship between real and nominal exchange rate volatility. We conclude thus that international goods markets are sufficiently segmented and nominal exchange rate movements are only a minor determinant of the firms' price adjustment strategies[6].

This paper is organized as follows. Section 2 presents the model. Section 3 discusses the relationship between relative price variability and nominal exchange rate volatility that the model generates. Section 4 discusses the data and presents the results of the gravity equations. Section 5 concludes. Appendices discuss the solution method used to solve the model and present some robustness checks.

[6]Obstfeld and Rogoff (2000) also emphasize the role of trade costs in explaining several major puzzles in international macroeconomics.

2. The Model

A. Setup

The world consists of two countries, Home and Foreign. Each country is inhabited by a continuum of identical consumers and a continuum of firms. Consumers buy goods from both Home and Foreign firms and supply labor to firms in their own country. They own local firms and consume their profits. In addition, consumers in both countries have access to a complete set of state-contingent securities denominated in Home's currency. We assume that Home is small relative to Foreign so that Foreign variables are unaffected by shocks in the Home economy. We make this assumption in order to maintain computational tractability and to allow the use of non-linear solution techniques.

We describe the problem of the agents in Home. Foreign agents solve identical problems and asterisks are appended to their variables. Throughout, let s_t denote the event realized at time t, $s^t = \{s_0, s_1, ..., s_t\}$ the history of events up to this period and $\pi(s^t)$ the probability of a particular history as of time 0.

Consumers

A representative consumer has preferences over a continuum of goods indexed by $z \in (0, 1)$. For each good z, two closely substitutable varieties, one produced in Home and another in Foreign, are available for consumption. The consumer optimally allocates her income across the different goods in the consumption basket, and decides, for each good, the optimal quantity to be imported from abroad. Letting $C(z, i; s^t)$ denote the consumption of variety $i \in \{h, f\}$ of good z in s^t, the Home consumer solves:

$$\max_{C(z,i;s^t),b(s^{t+1}),n(s^t)} \sum_{t=0}^{\infty} \beta^t \sum_{s^t} U(C(s^t), n(s^t)) \tag{1}$$

subject to

$$\int_0^1 \left(p(z;s^t)C(z,h;s^t) + \tau(z)e(s^t)p^*(z;s^t)C(z,f;s^t)\right) dz + \sum_{s^{t+1}} \chi(s^{t+1}|s^t)b(s^{t+1}) = W\left(s^t\right) n\left(s^t\right) + \Pi(s^t) + b(s^t) \tag{2}$$

where

$$C(s^t) = \left(\int_0^1 C(z; s^t)^{\frac{\theta-1}{\theta}} dz \right)^{\frac{\theta}{\theta-1}} \tag{3}$$

is a CES aggregator over the different goods and $C(z; s^t) = \left(C(z, h; s^t)^{\frac{\gamma-1}{\gamma}} + C(z, f; s^t)^{\frac{\gamma-1}{\gamma}} \right)^{\frac{\gamma}{\gamma-1}}$ is an aggregator over the two (Home and Foreign) varieties of good z. θ and γ are the elasticity of substitution across goods and varieties, respectively. $p(z; s^t)$ is the price of the domestically produced variety of z, $p^*(z; s^t)$ is the (Foreign currency) price of the Foreign variety. e is the exchange rate, denominated in units of Home per Foreign currency. $\tau(z)$ is an iceberg trade cost the consumer incurs in order to purchase goods from abroad: to consume 1 unit of the foreign good, $\tau(z) > 1$ must be purchased. We broadly interpret τ as costs of international trade, which include both physical costs (trade and non-trade barriers, shipping costs), but also subjective costs, such as a bias in preferences towards the consumption of domestically produced goods. Goods differ according to their tradeability: we calibrate $G(\tau)$, the distribution of trade costs across goods, below. $\chi(s^{t+1}|s^t)$ is the price of a bond that pays 1 unit of Home currency if state s^{t+1} is realized tomorrow. $b(s^{t+1})$ is the quantity of state-contingent bonds the consumer purchases. Finally, $\Pi(s^t)$ are the profits earned by all Home firms and $n\left(s^t\right)$ is the household's supply of labor. Implicit in this formulation of the budget constraint is the convention that consumers own only their country's firms[7].

We can solve the consumer's problem in several stages, using a duality approach. First, given prices, and a desired level of consumption of any particular good z, $C(z)$, the consumer chooses her consumption of the two (Home and Foreign) varieties of the good in order to minimize the expenditure associated with consuming $C(z)$:

$$\begin{aligned} \min_{C(z,h), C(z,f)} \quad & p(z)C(z,h) + \tau(z) e p^*(z) C(z,f) \\ \text{s.t.} \quad & \left(C(z,h)^{\frac{\gamma-1}{\gamma}} + C(z,f)^{\frac{\gamma-1}{\gamma}} \right)^{\frac{\gamma}{\gamma-1}} \geqslant C(z) \end{aligned}$$

[7]This, as well as the restriction that state-contingent bonds are denominated in Home's currency, are notational conventions, rather than assumptions. Given the availability of a complete set of state-contingent Arrow-Debreu securities, any additional asset in this economy is redundant (it does not affect equilibrium allocations), as it can be constructed using a combination of the state-contingent bonds available for trade. Chari, Kehoe and McGrattan (2002) employ a similar convention.

where we drop the dependence on s^t as this is a static problem. The first-order conditions of this problem require that, at the optimum, the marginal rates of transformation and substitution of Foreign and Home varieties be equal:

$$\frac{p(z)}{\tau(z)ep^*(z)} = \frac{C(z,h)^{-\frac{1}{\gamma}}}{C(z,f)^{-\frac{1}{\gamma}}} \tag{4}$$

Letting $P(z)$ be the Home price index for good z, defined as the minimum amount a Home consumer must spend in order to purchase one unit of $C(z)$:

$$P(z)C(z) = p(z)C(z,h) + \tau(z)ep^*(z)C(z,f), \tag{5}$$

it follows that

$$C(z,h) = \left(\frac{p(z)}{P(z)}\right)^{-\gamma} C(z),$$
$$C(z,f) = \left(\frac{\tau(z)ep^*(z)}{P(z)}\right)^{-\gamma} C(z),$$

where

$$P(z) = \left[p(z)^{1-\gamma} + (\tau(z)ep^*(z))^{1-\gamma}\right]^{\frac{1}{1-\gamma}}$$

The share of good z imported from abroad depends on the price the Foreign firm charges (in Home currency terms), times the good's shipping cost, relative to a consumption-weighted average price of the Home and Foreign varieties of this good. The problem we study is thus a generalization of the standard trade-cost model, exposited, for example, in Sercu, Uppal and van Hulle (1995), in which Home and Foreign varieties of a good are perfect substitutes ($\gamma \to \infty$), and consumers purchase only the cheapest variety of a given good.

An argument similar to the one above can be used to show that the rule according to which consumers allocate their income across the different goods, z, is given by:

$$C(z) = \left(\frac{P(z)}{P}\right)^{-\theta} C$$

where P is the aggregate price level in Home, defined as a consumption-weighted average of the prices of all goods:

$$P = \int_0^1 \frac{C(z)}{C} P(z) dz = \left(\int_0^1 P(z)^{1-\theta} dz\right)^{\frac{1}{1-\theta}}$$

Finally, standard first-order conditions govern the optimal supply of labor and demand for state-contingent securities:

$$n(s^t) : \frac{W(s^t)}{P(s^t)} = -\frac{U_n(s^t)}{U_c(s^t)}$$

$$b(s^{t+1}) : \frac{P(s^{t+1})}{P(s^t)} \chi(s^{t+1}|s^t) = \beta\pi(s^{t+1}|s^t) \frac{U_c(s^{t+1})}{U_c(s^t)}, \tag{6}$$

where U_n, U_c are the derivatives of the utility function with respect to its arguments.

Foreign consumers face a similar problem. In our exercises, we assume that the Foreign economy is large compared to the Home economy. We do so formally by assuming that there is a unit mass of home agents and the Foreign's population is N times larger, where $N \to \infty$. We also assume that Foreign consumers place little weight on their consumption of Home varieties of goods:

$$C^*(z; s^t) = \left(\omega^* C^*(z, h; s^t)^{\frac{\gamma-1}{\gamma}} + C^*(z, f; s^t)^{\frac{\gamma-1}{\gamma}}\right)^{\frac{\gamma}{\gamma-1}},$$

and $\omega^* \to 0$. Without this assumption, Home firms will de-facto sell only abroad independent of the size of their trade costs and this would preclude us from studying the role of tradeability in the model. Foreign consumer's demand for the Home-produced variety of a given good is:

$$C^*(z,h;s^t) = \omega^{*\gamma}\left(\frac{\tau(z)p(z;s^t)}{e(s^t)P^*(z;s^t)}\right)^{-\gamma} C^*(z;s^t),$$

where

$$P^*(z;s^t) = \left(p^*(z;s^t)^{1-\gamma} + \omega^{*\gamma}\left(\frac{\tau(z)p(z;s^t)}{e(s^t)}\right)^{1-\gamma}\right)^{\frac{1}{1-\gamma}}$$

is the aggregate price of good z from the perspective of the Foreign agent. Note here the consequence of our assumption that $\omega^* \to 0$: Foreign prices and quantities are independent of disturbances originating in Home. Nevertheless, because Foreign is large relative to Home, total demand for Home goods is non-zero.

To pin down the nominal exchange rate in this model, we make use of the Euler equation that governs the Foreign consumer's demand for state-contingent securities[8]:

$$b^*(s^{t+1}) : \frac{P^*(s^{t+1})}{P^*(s^t)}\frac{e(s^{t+1})}{e(s^t)}\chi(s^{t+1}|s^t) = \beta\pi(s^{t+1}|s^t)\frac{U_c^*(s^{t+1})}{U_c^*(s^t)} \tag{7}$$

Equations (6) and (7) together imply that the ratio of the marginal utilities of consumption in Home and Foreign are proportional to the real exchange rate:

$$\frac{U_c(C(s^t), n(s^t))}{U_c(C^*(s^t), n^*(s^t))} = \Lambda\frac{P(s^t)}{e(s^t)P^*(s^t)},$$

where Λ is a constant that depends on initial conditions. We normalize it to ensure that $e = 1$ in the steady-state.

[8]The problem we study here is standard in many aspects and we only present the model's key equations. We refer the reader to Chari, Kehoe and McGrattan (2002) for additional details.

Firms

Firms set a unique price $p(z;s^t)$ or $p^*(z;s^t)$ denominated in their own country's currency[9]. Both countries' consumers buy at this price. Customers from abroad pay an additional (iceberg) trade cost $\tau(z)-1$: in order to consume 1 unit of the good, $\tau(z)$ must be purchased.

Labor is the sole variable input in production: the firm's technology is $y(z) = l(z)^\alpha$ where $l(z)$ is the quantity of labor the firm uses for production. Firms also have menu costs of adjusting prices: the firm must hire $\xi(z)$ additional units of labor every time it resets its price. ξ differs across firms: more on this point below. A typical firm in Home solves:

$$\max_{p(z;s^t)} \sum_{t=0}^{\infty} \sum_{s^t} \pi(s^t) Q(s^t) \Pi(z;s^t) \tag{8}$$

where $Q(s^t) = \beta^t \frac{U_c(C(s^t),n(s^t))}{U_c(C(s^0),n(s^0))}$ is the t-period ahead stochastic discount factor, and $\Pi(z;s^t)$ are the firm's profits in state s^t

$$\Pi(z;s^t) = \frac{p(z;s^t)}{P(s^t)} D(z;s^t) - \frac{W(s^t)}{P(s^t)} D(z;s^t)^{\frac{1}{\alpha}} - \xi(z) \frac{W(s^t)}{P(s^t)} \mathcal{I}_{p(z;s^t) \neq p(z;s^{t-1})} \tag{9}$$

where $\xi(z)\frac{W(s^t)}{P(s^t)}$ is the additional cost the firm must pay every time it reprices, $D(z;s^t)$ is total demand (we make use of the demand functions derived above, and of the fact that the measure of Foreign agents is N):

$$D(z;s^t) = \left(\frac{p(z;s^t)}{P(z;s^t)}\right)^{-\gamma} \left(\frac{P(z;s^t)}{P(s^t)}\right)^{-\theta} C(s^t) + N\tau(z)\omega^{*\gamma} \left(\frac{\tau(z)p(z;s^t)}{e(s^t)P^*(z;s^t)}\right)^{-\gamma} \left(\frac{P^*(z;s^t)}{P^*(s^t)}\right)^{-\theta} C^*(s^t)$$

Note that per-capita demand from Foreign consumers is $\tau(z)C^*(z,h)$, as, given a desired level of consumption $C^*(z,h)$, the consumer must purchase a total of $\tau(z)C^*(z,h)$ units, of which a fraction melts in transit. We assume parameter values in the Foreign economy that ensure that per-capita consumption and aggregate price levels are equal in both countries at the steady state: $C = C^*$, $P = P^*$ and that $\omega^{*\gamma}N = 1$. These

[9]Here we depart from the standard local currency pricing New Keynesian open economy models in that we assume that producers cannot segment across markets. The wedge between the foreign and domestic price of a Home-produced variety arises solely due to the transport cost. We do so because assuming local currency pricing implictly rules out international goods market arbitrage and does not allow us to study the role of tradeability.

assumptions imply that, at the steady-state, a Home firm which faces no trade costs treats the Home and Foreign markets symmetrically, as its export share is close to $\frac{1}{2}$[10].

B. Equilibrium

We introduce money by imposing a quantity theory money demand and assuming unitary velocity: $P(s^t)C(s^t) = M(s^t)$, and similarly for the Foreign economy. The Foreign money supply is assumed constant. This, as well as the assumption that Foreign preferences are heavily skewed towards the consumption of Foreign goods ($\omega^* \longrightarrow 0$), implies that Foreign aggregate variables are constant at their steady-state values. The only source of uncertainty in the economy are shocks to the growth rate of the Home money supply: $g(s^t) = \log\left(\frac{M(s^t)}{M(s^{t-1})}\right)$.

The equilibrium is a sequence of prices $P(s^t), P(z; s^t), p(z; s^t), e(s^t), W(s^t), \chi(s^{t+1}|s^t)$, and allocations $n(s^t), C(z, i, s^t), b(s^{t+1})$ that are consistent with firm and household maximization, and satisfy the market-clearing and resource constraints.

We normalize all nominal variables by the money supply in this economy, e.g., $\tilde{P}(s^t) = \frac{P(s^t)}{M(s^t)}$, in order to render the state-space of this problem bounded. Let $\tilde{p}_{-1}(z, s^t) = \frac{\tilde{p}(z, s^{t-1})}{M(s^t)} \in \mathcal{P}$ be a firm's (normalized) last period's price and $\mathcal{T} = [\tau_{\min}, \tau_{\max}]$ the support of the distribution of transport costs. The aggregate state of this economy is an infinite-dimensional object, consisting of the growth rate of money in this economy: g, but also of the endogenously varying joint distribution of last period's firm prices and transport costs. Let $\boldsymbol{\mu} : \mathcal{P} \times \mathcal{T} \to [0, 1]$ denote this distribution and Γ its law of motion: $\boldsymbol{\mu}' = \Gamma(g, \boldsymbol{\mu})$. We follow Krusell and Smith (1997) and approximate $\boldsymbol{\mu}$ with a finite-dimensional vector of moments in order to allow use of numerical solution techniques to solve the model. The unknown functions that characterize the equilibrium of this economy are solved for via collocation, a functional approximation technique. In particular, we approximate the unknown functions (aggregate prices and quantities, the endogenous law of motion for $\boldsymbol{\mu}$, as well as the firm's value functions) with linear combinations of orthogonal polynomials, and solve for the unknown coefficients on these polynomials by requiring that the equilibrium conditions are satisfied at a finite number of nodes along the state-space.

To solve the firm's problem, let $\Xi = \{g, \boldsymbol{\mu}\}$ collect the aggregate state variables, $\Pi(\hat{p}) = D(\hat{p})\frac{\hat{p}}{\hat{P}} -$

[10]The exact value of the export share depends on α, the elasticity on labor in the production function.

$D(\hat{p})^{\frac{1}{\alpha}} \frac{\hat{W}}{\hat{P}}$ be the firm's profits (exclusive of menu costs), $\tilde{p}_{-1} = \frac{p_{-1}}{M}$ be the firm's last period's price, $V^a(\tau; \Xi)$ denote the value of adjustment of a firm with trade cost τ, and $V^n(\tilde{p}_{-1}, \tau; \Xi)$ be the firm's value of inaction. Then, letting $V = \max(V^a, V^n)$ denote the firm's value, the following two functional equations characterize the firm's problem recursively:

$$\begin{aligned} V^a(\tau; \Xi) &= \max_{\tilde{p}} \left[U_c \left(\Pi(\tilde{p}) - \xi \frac{\tilde{W}}{\tilde{P}} \right) + \beta E V \left(\tilde{p}'_{-1}, \tau; \Xi' \right) \right] \\ V^a(\tilde{p}_{-1}, \tau; \Xi) &= U_c \Pi(\tilde{p}_{-1}) + \beta E V \left(\tilde{p}'_{-1}, \tau; \Xi' \right) \end{aligned} \tag{10}$$

where we follow Khan and Thomas (2001) and normalize the firm's value functions by the marginal utility of consumption in order to render the discount factor time-invariant. The law of motion for $\tilde{p}_{-1}$ is

$$\begin{aligned} \tilde{p}'_{-1} &= \frac{\tilde{p}}{\exp(g)} \text{ if adjust} \\ \tilde{p}'_{-1} &= \frac{\tilde{p}_{-1}}{\exp(g)} \text{ otherwise} \end{aligned}$$

An appendix discusses the solution method in more detail.

3. The relationship between RPV and NER volatility

A. Parametrization and calibration

We parameterize the utility function as

$$U(C, n) = \frac{1}{1-\kappa} C^{1-\kappa} - \psi n$$

This specification follows Hansen (1985) by assuming indivisible labor decisions implemented with lotteries. The length of the period is one month and we set the discount factor equal to $\beta = .997$. The value of ψ is chosen to ensure that households work 30% of their discretionary time in the steady state. We follow Chari, Kehoe and McGrattan (2002) and set $\kappa = 5$ so that the model produces sufficiently large movements in real exchange rates. The elasticity of substitution across goods is set equal to $\theta = 3$ to highlight the fact that goods are imperfect substitutes, a lower bound of the range of elasticities used in closed-economy models[11]. On the other hand, varieties of a good are closely substitutable, and we set $\gamma = 15$, somewhat higher than the estimates reported by Hummels (2001) using disaggregated international trade data, in order to study the behavior of relative prices for almost identical Home and Foreign varieties of a given good. We interpret the production function $y = l^{\alpha}$ as reflecting the fact that other factors of production are fixed in the short-run. The share of labor is set to $\alpha = \frac{2}{3}$, a typical choice in the business cycle literature.

To calibrate the distribution of trade costs, we turn to the US Imports and Exports Data compiled by Robert Feenstra. This dataset contains information on the total value of exports by manufacturing firms at the SIC 4-digit level for 1958-1994. We supplement this dataset with the NBER Productivity Database which contains, among others, data on total shipments for these industries. Using the two sets of data, we calculate export shares for 396 industries for 1993[12]. Based on these shares, we calculate (weighted by size of the industry) second, third and fourth moments of the distribution of export shares that we ask the model to match. We do not use the US data to match the mean of this distribution, because US is a relatively closed economy and because the disaggregated data is available only for the manufacturing sector. Instead, we turn to the OECD STAN Indicators Database for 1993 for shares for most of the OECD countries. Table 1 presents the OECD STAN data we use: the average OECD economy exports 11.87% of its production abroad. We assume a parametric functional form for the distribution of trade costs $G(\tau)$: the beta distribution, for its flexibility. The two parameters of the distribution are chosen so that the model delivers moments of the export shares distribution close to those in the data. Table 2 presents the moments in the data we target and the ones implied by the distribution of trade costs we work with: $G(\tau) =\text{beta}(a_1, a_2)$ with $a_1 = 3.4$ and $a_2 = 13$. Trade costs range in the model from 2% to 48%, with a mean of 20% ($\bar{\tau} = 1.2$). This large dispersion of trade costs across varieties is consistent with results from earlier work that has provided direct measurements of the size of transport costs. A notable example is Hummels (2001) who has assembled an extensive dataset of freight rates for 2-digit SITC industries for US and several other countries. He finds large dispersion in freight rates, both across commodities, but also across transportation modes: average freight

[11]See Obstfeld and Rogoff (2000) for a brief survey.

[12]We have discovered several errors for the 1994 dataset downloaded from the UC Davis website and went back one year in time to avoid them.

rates range from 3.5% (Office Machines) to 28.6% (Coal, Coke) for US, and are as large as 50% for other countries.

We finally assume that menu costs of price adjustment $\xi(z)$ differ across firms: firms pay 1.2% of their steady state revenues each time they adjust prices. Firms that sell more abroad face larger menu costs because of their larger market shares. We make this assumption in order to ensure that our result that more tradeable firms adjust more frequently is not simply due to the fact that these firms have smaller menu costs relative to their revenues. Our choice of menu costs implies that firms adjust every 7 months on average for monetary shocks of a size similar to that in the US data, an average price duration somewhat larger than that reported by Bils and Klenow (2002) for a large dataset of consumer prices in the US economy, and is consistent with the evidence presented by Zbaracki et. al. (2004) who study the price adjustment practices of a large US manufacturing firm.

We use the model to ask how goods prices respond to movements in exchange rates, and require the model to generate nominal exchange rate fluctuations consistent with those observed in the data. We therefore assume a process for the growth rate of the money supply to ensure that nominal exchange rates follow, in equilibrium, a random walk process:

$$\log e(s^t) = \log e(s^{t-1}) + \epsilon(s^t) \tag{11}$$

where $\epsilon \sim N(0, \sigma^2)$[13]. We vary σ in simulations to study how the variability of relative prices depends on the volatility of changes in nominal exchange rates. To be clear, even though the nominal exchange rate follow a random-walk subject to random disturbances, it does not constitute an exogenous variable in the model. Rather, the growth rate of the money supply is calibrated in order to generate realistic nominal exchange rate fluctuations.

B. Results

The only source of aggregate fluctuations in this economy are shocks to the growth rate of the money supply. The mechanism through which a monetary impulse propagates in the Home Economy is reminiscent

[13]Using the risk-sharing condition, and the functional form of the utility function assumed above, the growth rate of the money supply follows: $\log \frac{M(s^t)}{M(s^{t-1})} = (1-k) \log \frac{C(s^t)}{C(s^{t-1})} + \epsilon(s^t)$.

of that at play in standard models with nominal rigidities: a monetary expansion, given sluggish price adjustment, increases the purchasing power of Home's agents, whose consumption increases. As a result, the real wage rate increases, as the increase in consumption lowers the marginal utility of every additional (real) dollar earned in the labor market and households require higher real wages. Faced with higher marginal costs of production (arising both because of the higher real wages, but also because of a movement along an upward sloping marginal cost curve, as $\alpha < 1$), those firms whose prices are too far away from the optimum choose to pay the menu costs and reset their nominal prices. Finally, the nominal exchange rate depreciates, initially overshooting the increase in the money supply, and thus exhibiting more volatility than fundamentals, a result similar to that of Dornbusch (1976). We quantify, in Table 3, the strength of each of these effects, by calculating the volatility of macroeconomic aggregates (filtered using a first-difference operator) for simulations of the model in environments where the volatility of nominal shocks is $\sigma = .01$ and $\sigma = .03$, respectively. Note that, when the volatility of nominal shocks is low, consumption growth rates are 7.4 times less volatile than nominal exchange rate changes. while real wages, real exchange rates, and money supply are almost 1.5 times less volatile than nominal exchange rates. As the environment becomes more volatile, firms find it optimal to adjust prices more frequently, and business cycle fluctuations are dampened in the $\sigma = .03$ economy (relative to the volatility of nominal shocks), while the ratio of real to nominal exchange rate volatility decreases, results similar to those of Alvarez, Atkeson and Kehoe (2002) in a model with endogenously segmented markets.

We next turn to the behavior of producers. Figure 1 depicts a firm's value of inaction: V^n, and adjustment, V^a, as a function of the log-deviation of the firm's price from its optimum[14], when all other variables are at their steady-state levels. We plot these functions for two firms: one that sells tradeable goods ($\tau = 1.02$), and one whose good is virtually untradeable, at the upper end of the distribution of trade costs in the model: $\tau = 1.48$. The two functions are normalized so that the maximum a firm's value of not changing its price can reach is 1.

If the firm adjusts, it resets the price to its optimal level: the firm's value of adjustment is therefore independent of p_{-1}. The best a firm can do by not adjusting is if its past price coincides with today's optimum:

[14] The standard deviation of nominal shocks is $\sigma = .01$ in the exercises of Figures 1 and 2.

the firm's value of inaction in this case exceeds the value of adjustment by the size of the menu cost. As the firm's price deviates away from the optimum, the value of inaction falls and is eventually lower than the value of changing its price; the intersection of the two value functions determines the region of inaction – the (S, s) bands.

Notice also, in the right panel of Figure 1, that a firm's value of inaction is much more sensitive to deviations of the firm's price from its optimum if the firm's product is more tradeable. Intuitively, if trade costs are low, Home and Foreign agents switch more easily towards the cheapest variety of a particular good: relative price changes therefore generate larger fluctuations in demand, and therefore marginal costs for producers of more tradeable products. This in turn implies that the inaction region is wider for the high transport cost firm (despite our assumption that menu costs are proportional to steady-state revenues and therefore lower for high trade-cost firms): it tolerates prices as far as 4% away from its optimum. A low trade-cost firm adjusts more readily, every time its price deviates from the optimum by 1.5% in absolute value.

Figure 2 plots the two firms' price functions, conditional on adjustment, p, as well as the (S, s) bands of price adjustment, as a function of the nominal shock, ε. Prices are expressed as deviations of the price from the steady-state optimum. Note that optimal price functions are similar to those the firm would employ in a flexible price world (a 45-degree line) in which prices increase one-for-one with the nominal shock. This result is driven by the fact that nominal wages and exchange rates are proportional to each other (because of our assumption of an infinitely elastic Frisch labor supply): $W \sim \frac{P}{U'(C)} \sim e$, and exhibit a random-walk behavior. An important component of the firm's marginal costs is therefore unforecastable based on current information. Moreover, even though a, say, monetary expansion increases aggregate consumption and thus the marginal cost of production, firms that do adjust in times of a monetary expansion set prices that are higher than those of other firms in the economy who have not yet responded to the nominal shock: substitution by consumers towards the cheaper goods thus reduces these firms' total sales and hence their marginal costs. These two opposite effects cancel each other out for our choice of parameter values. Because adjusting firms respond one-for-one to a nominal shock, the law of one price holds for firms who have just reset their prices: relative price variability is thus solely a function of the frequency of price adjustment and the volatility of nominal exchange rates in this environment. Note again in the figure that (S, s) bands of price adjustment are wider for firms that sell less tradeable goods. Moreover, the width of these bands is unaffected by the size of the nominal disturbance.

We illustrate the key predictions of the model in panels B and C of Table 3. Note that differences in trade costs induce large differences in the frequency of price changes: firms with small ($\tau = 1.02$) trade costs adjust every 2.7 months when the volatility of nominal shocks is equal to 1%. In contrast, high trade cost firms adjust much less frequently: every 29 months on average. In environments with higher nominal exchange rate volatility ($\sigma = 3\%$) all firms find it optimal to pay the menu costs and adjust their prices more frequently: low trade cost firms adjust prices every 1.4 months, whereas firms with large trade costs change prices every other 5 months. Note in panel C of the table how these differences in the frequency of price changes translate into differences in relative price variability, where we measure relative price variability by the time-series standard deviation of changes in the relative price charged by Home and Foreign producers of a given good: $q_t(z) = \log \frac{e_t p_t^*(z)}{p_t(z)}$. Relative prices are twice more volatile for high-trade cost firms than for producers of more tradeable goods when the volatility of nominal exchange rates is low. In more volatile environments, relative price variability falls for low trade cost goods (from 0.53% to 0.42%), while it increases for goods that are traded little (from 1.1% to 2.7%).

It is helpful, at this point, to relate the predictions of this model to those of the standard trade-cost model that assumes that goods are perfectly substitutable and arbitrage eliminates law of one price deviations whenever these are too large. A key prediction of the standard trade-cost model is that the relative price of the good can deviate away from unity, but is bounded by the size of the transport cost: $-\log \tau \leqslant q_t \leqslant \log \tau$. Our discussion of Figure 2 should convince the reader that a similar behavior is implied by the menu costs model in which goods are less-than-perfect substitutes. The firm will allow its price to deviate away from its optimum, as long as the deviation does not exceed the S, s bands, and charge a price that (approximately) ensures that the law of one price holds otherwise. The behavior of relative prices can then be approximated by:

$$q_t = 0 \text{ if } \quad q_t^* \notin [s(\tau), S(\tau)]$$

$$q_t = q_t^* \text{ if } q_t^* \in [s(\tau), S(\tau)]$$

$$q_t^* = q_{t-1} + \epsilon_t$$

where q_t^* is the relative price that would prevail in the absence of price adjustment. In the menu costs-model the inaction bands are non-linear functions of the trade cost, elasticities of substitution, as well as the volatility of the environment, but the similarity with the standard trade-cost model is evident. In particular, more tradeable goods command narrower inaction regions in the menu-costs model, and allow smaller relative price fluctuations.

Figure 3 examines the relationship between relative price variability and nominal exchange rate volatility in more detail. We simulate the model for a range of σ and calculate $std(\Delta q_t)$ for each simulation. The upper panels of Figure 3 present results for firms with the lowest trade cost: $\tau = 1.02$ (left), mean trade cost: $\tau = 1.20$ (middle), and highest trade cost: $\tau = 1.48$ (right). The lower panels plot the average number of periods between two consecutive adjustments. The hump-shaped relationship between relative price variability and nominal exchange rate volatility is evident in this figure. When the volatility of nominal exchange rates is sufficiently low, firms adjust infrequently and relative prices inherit the properties of the nominal exchange rate. An increase in NER volatility then increases the volatility of relative prices, a result similar to that one obtains in standard sticky price models in which the frequency and timing of price changes is assumed exogenous. Notice however that when pricing is state-dependent the positive relationship between nominal exchange rate volatility and relative price variability holds only locally, in environments in which firms change prices sufficiently infrequently. As the frequency of price changes increases, the tight link between real and nominal international relative prices breaks down and higher nominal exchange rate variability, through its effect on the frequency of price changes, decreases the variability of relative prices. In our calibration of the model, whenever nominal exchange rates are sufficiently large to induce firms to adjust prices every 2 to 3 months, higher nominal exchange rate variability decreases the volatility of relative prices.

Note also that the hump in the relationship between NER volatility and relative price variability occurs for lower values of σ if goods are more tradeable. Relative price variability for goods characterized by small trade costs decreases with additional increases in the volatility of nominal exchange rates whenever the volatility of nominal exchange rates exceeds 1%. In contrast, for firms with $\tau = 1.48$, larger NER volatility increases relative price variability for most of the range of our simulations: only for values of σ close to 6% will larger NER volatility cause lower relative price variability. The difference arises because large nominal exchange rates fluctuations are necessary in order to induce high trade cost firms to adjust price sufficiently frequently (2-3 months) to break the link between real and nominal international relative prices.

We next turn to other measures of law of one price deviations. One of the reasons first-differences, rather than levels of relative prices are used in work with disaggregated price data is the interest in the volatility of short-run LOP deviations. Suppose that the relative price of a good, q_t, follows an autoregressive process:

$$q_t = \rho q_{t-1} + \varepsilon_t.$$

The purpose of the gravity-type equations estimated by Engel and Rogers (1996, 2001) and Parsley and Wei (2001a, 2001b) is to explain the volatility of ε_t rather than that of q_t. Given that real exchange rates are highly persistent, first-differencing q_t provides a good estimate of short-run shocks to the relative price, embodied in ε_t. Our model, on the other hand, predicts that ρ itself varies with the volatility of nominal exchange rates. We therefore estimate an AR(1) process for q_t and plot, in Figure 4, the volatility of the in-sample forecast errors, $\hat{\varepsilon}_t$, the persistence of the relative price series, as measured by its AR(1) coefficient, $\hat{\rho}$, as well as the unconditional volatility of relative prices, $std(q_t)$. We plot these statistics for a firm with the median transport cost, $\tau = 1.2$, as a function of the volatility of nominal exchange rates in each simulation (upper panels), and, also in the τ space, for an economy in which $\sigma = .01$.

Notice, in the upper-left panel of the figure, that the persistence of relative prices declines with higher NER volatility: when σ is close to 0, firms adjust infrequently and relative prices track nominal exchange rates closely: the estimate of ρ is therefore close to 1. In contrast, ρ quickly decays to 0 for values of σ larger than 3%: in this region of the parameter space firms adjust prices every 2 months and allow only transitory

LOP deviations. As the lower-left panel of the figure indicates, differences in trade costs also induce large differences in the persistence of relative prices across goods: firms with no trade costs allow only transitory relative price fluctuations even though the volatility of nominal exchange rates in this exercise is low (ρ is close to 0.1). In contrast, high-trade cost firms adjust infrequently, and the relative price of these firms more persistent (ρ is close to 0.9). Similar results hold for other measures of LOP deviations.

We finally ask how aggregate international relative prices respond to changes in the volatility of nominal exchange rates. As Figure 5 shows, real exchange rates behave similarly to the relative price of the good with the median trade cost with a hump at $\sigma = 2.5\%$. This figure underscores the necessity of using disaggregated price data in order to test the predictions of the state-dependent model: focusing on aggregate price indices may hide the negative relationship between RPV and NER volatility for tradeable goods in the consumption baskets if a large fraction of the goods in the economy are less tradeable.

4. The Data

A. Evidence from Micro-price data

Our first empirical exercise tests the predictions of the model using a dataset of actual goods prices. The model predicts a negative relationship between relative price variability and NER volatility in environments where firms adjust prices sufficiently frequently (more frequently than every 2-3 months) and/or goods in different markets are highly tradeable/substitutable. The dataset we use is collected in precisely such an environment: we work with prices for homogenous agricultural products sold in open open air markets, an environment in which prices change frequently and are volatile due to seasonal demand and supply disturbances. The data was collected in 12 cities in six Eastern European countries by the Central Agricultural Market Information Bureau (CAMIB[15]), a NGO created by an European Union project aimed at providing informational support to food exporters in Moldova, one of the countries in the sample. Our sample covers the years of 2000 to 2002[16]. Prices are collected on average once every two weeks (usually during weekends)

[15] The data is available (in Russian and Romanian), against a nominal fee, at www.camib.org.

[16] An earlier version of this paper used two more years of data: 1998 and 1999, years characterized by the excessive nominal exchange rate volatility induced by the Russian financial crisis. We drop these years in this version of the paper because in times of crises real exchange rate movements arise due to non-monetary factors our model abstains from. Consistent with the predictions of the model, law of one price deviations are smaller during the years of the crisis than in the rest of the sample, except in the periods immediately following the devaluations. Empirical results based on time-series measures of price dispersion are not robust to alternative measures of variability as well as to differences in the estimation technique employed, perhaps due to the increase in price dispersion immediately following the devaluations. On the other hand, cross-sectional measures of price

in all years except for 2000, when prices are sampled weekly: a total of 83 time periods are available.

Goods in four good-categories are available: meat, fruit, vegetables, as well as a small number of other agricultural products (oil, honey etc.). The primary purpose of the collecting agency is to provide local entrepreneurs with information about arbitrage opportunities in domestic and foreign markets: efforts are therefore made to ensure comparability of products across locations and accuracy of the data. The agency hires representatives in all 12 cities who sample prices several times during the day, from a number of vendors of a given product. An average price, across all vendors surveyed, is then reported. Goods are all homogenous (beef ribs, beef fillet, apples, watermelon, shelled nuts, honey, tomatoes, etc.). In the case of fruits and vegetables, where quality differentials can easily lead to price differences, the agency reports the prices of both high and low quality products. Finally, most vendors set several prices for identical goods sold in a given day due to differences in the size, quality, or color of the product. The agency treats the most and least expensive varieties of a good as separate items and records their prices separately. We follow this convention as well and treat high and low-quality varieties of a given product as separate goods in our sample.

Not all prices are sampled in all periods: some goods, especially fruits and vegetables, are not sold in open-air markets during all months of the year, and, especially during the cold season, drop out of the sample. We therefore work only with those goods for which a panel of at least 6 cities is available in at least 75% of the time-periods in our sample. A total of 58 goods make this cutoff. Table A1 in the appendix lists the cities and goods available in our sample as well as basic statistics regarding the availability of data. Data for all cities and goods is available at least 80% of the time, with the exception of Minsk (Belarus) where the agency started collecting the data late in our sample. Most goods are meat products (22), and vegetables (20), while fruits are under-represented because they are mostly unavailable during winters.

We calculate, in Table 4, several statistics that capture the extent to which the law of one price is violated in the data. We first focus on absolute LOP deviations. We calculate, in the spirit of Crucini, Telmer and Zachariadis (2005), for each city and good in the sample, the log-deviation of a city's price from

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dispersion we are about to use below are not heavily influenced by the large real devaluations following large nominal exchange rate swings and show consistently a negative relationship between RPV and NER volatility. Results of these regressions are available from the author upon request.

the cross-sectional average: $d_{tg}^c = \log\left(\frac{p_{tg}^c}{\bar{p}_{tg}}\right)$, where p_{tg}^c is the price of good g in city c at time t, and $\bar{p}_{tg}$ the cross-sectional average of the price of this particular good in period t. The law of one price predicts that d_{tg}^c should be centered at 0 in the absence of frictions that prevent goods-market arbitrage. As Table 4 indicates, this proposition is grossly violated in our sample. We report, in the table, the median and interquartile range of the distribution of d_{tg}^c (across goods and time-periods), for each city in the sample. St. Petersburg (Russia) is the most expensive city: prices here exceed the average price in the 12 cities by around 30% on average. St. Petersburg is followed closely by Moscow (Russia) and Bucharest (Romania), whose prices are 26% and 23% higher, on average, than those of other cities in the sample. Cities in Ukraine and Moldova are least expensive. Note also that these median statistics are large (in absolute value) relative to the interquartile range of the distribution of d, suggesting that some cities are consistently over- (under-) priced for all goods in our sample, in most time periods. Although large, absolute law of one price deviations are consistent with economic theory. In particular, prices are more expensive in (relatively) rich countries (Russia and Belarus)[17], and cheaper in countries that are net exporters of agricultural products (Moldova and Ukraine)[18].

We next turn to relative LOP deviations, defined as fluctuations of relative prices around a trend. Let q_{tg}^i be the log-relative price of good g at time t for city pair i. We focus on the time-series properties of q in order to characterize relative LOP deviations. In particular, we calculate the time-series standard deviation of changes and levels of the relative price series in order to gauge the extent to which relative prices fluctuate over time. Note, in Table 4, that changes and levels of relative prices are very volatile in this data, with an average standard deviation (across goods/city-pairs) in excess of 0.25 (changes) and 0.37 (levels). Fruits and vegetables are more volatile than meat products and goods in the "other" product category, perhaps because of the seasonal nature of these products, but also because of their perishability and difficulty to transport. A positive, albeit small, border effect is evident in the data: international relative prices are more volatile than intranational ones: national borders increase the volatility of changes in relative prices by 2% (from 25 to 27%), and that of levels of relative prices by 9% (from 37 to 46%).

[17] According to the Penn World Tables, 6.1, output per capita in the countries in our sample is 0.28 (Russia), 0.24 (Belarus) 0.14 (Ukraine), 0.14 (Romania), 0.06 (Moldova) relative to that of the United States.

[18] According to statistics published by the World Trade Organization, Moldova and Ukraine are net exporters of foodstuffs, with a ratio of exports to imports equal to 2.8 and 2.9, respectively. Belarus, Romania and Russia are net importers of food products (with a ratio of exports to imports equal to 0.48, 0.49 and 0.35, respectively).

We also calculate, for comparison, the average volatility of nominal exchange rates (defined as the time-series standard deviation of changes in the (log) nominal exchange rates in the periods between two price surveys, usually at bi-weekly intervals). Nominal exchange rates are volatile, but much less so than relative prices, once again suggesting that the environment we work with is an extremely volatile one, goods prices are close to flexible, and nominal shocks have a small effect on the behavior of relative prices. This is, then, the type of environment where the model predicts a negative relationship between nominal exchange rate volatility and the variability of relative prices. Moreover, the model also predicts that law of one price deviations should be larger when goods are more difficult to transport across locations. We test these two predictions of the model by estimating a cross-sectional gravity-type equation relating the time series volatility of changes (or levels) of q^i_{tg} for each of the 66 city pairs in our dataset to the volatility of nominal exchange rates, a border dummy and the (great-circle) distance between the city-pairs. The border dummy captures the numerous obstacles exporters face when crossing the border (including formal tariff and other trade barriers, but also informal dues/bribes they have to pay at customs checkpoints. Trade costs should also increase in distance, as documented, for example, by Hummels (2001)[19].

A limitation of the data is the fact that in several instances the agency samples the price data at irregular intervals (1 or 3 weeks, as opposed to the usual 2 weeks)[20]. To calculate the volatility of biweekly changes in relative prices, we assume that changes in relative prices are serially uncorrelated. Although this is a strong assumption, we report results based on first-differences in relative prices for comparability with earlier work. The volatility of levels of relative prices is not subject to this limitation. The first model we estimate is

$$vol(\Delta q \text{ or } q)_{gi} = \gamma_0 + \gamma_1 Border_i + \gamma_2 \log(distance)_i + \gamma_3 std(\Delta e)_i + \mu_g + \varepsilon_{gi} \tag{12}$$

where $vol(\Delta q \text{ or } q)_{gi}$ is the time-series volatility of changes or levels of relative prices. We use two measures

[19]We have also experimented with additional variables aimed at capturing transport costs as well as homogeneity of preferences across trading partners: a common language dummy, former Soviet Union dummy, adjacency dummy, weight-to-value ratio (kg/$) of each good (Hummels (2001) documents that heavier goods are costlier to transport). None of these variables enter significantly and we have excluded them from the analysis.

[20]This, as well as the fact that price series are frequently interrupted because of missing data, precludes us from studying the persistence of the relative price series without imposing additional assumptions on the data generating process.

of relative price variability, the standard deviation, as well as the interquartile range. Given the excessive volatility of the environment, the interquartile range is a more robust measure of LOP deviations given its immunity to outliers. In addition, we include city and type of good (meat, fruit, vegetables) dummies in all regressions.

This regression pools together all 58 goods in the sample. We assume that μ_g, the good specific effect, is uncorrelated with other regressors and estimate a random effects GLS model. Table 5 presents the results. Note first that, consistent with the results of earlier studies, national borders and the distance between locations have the expected positive effect on relative price variability. The estimates of the importance of the border relative to distance in generating relative price movements depend, however, on the measure of relative prices used. National borders are an important source of relative price movements, relative to distance, for conditional measures of price dispersion (the volatility of first-differences of relative prices), but unimportant for unconditional measures (the volatility of levels of relative prices). Given that the volatility of first-differenced relative prices measures short-term deviations from the law of one price, while that of levels of relative prices is a long-run property of the series, one interpretation for this finding is that goods-market arbitrage occurs with a lag, and transport costs only affect the speed with each relative prices revert to the mean, rather than the size of shocks to relative prices, an issue beyond the scope of this paper.

More importantly for our present discussion, note that the volatility of nominal exchange rates always enters with a negative sign. Moreover, when the more robust interquartile range is used as a measure of dispersion, coefficient estimates are significantly different from zero at conventional significance levels. The dataset we use is therefore consistent with the predictions of the state-dependent sticky price model: larger NER volatility generates less relative price variability in an environment in which costs of changing prices are relatively low, idiosyncratic shocks volatile, and prices adjust frequently[21].

Parsley and Wei (2001a) have used an alternative measure of deviations from the law of one price to the time-series volatility of relative prices. Their paper studies the evolution of price dispersion over time and therefore use a measure of price dispersion for every given period by calculating the volatility of Δq_t across

[21]We have performed several check to ensure the robustness of this results. A negative relationship between RPV and NER volatility is present when we estimate the gravity equation above separately for different good-categories, for goods of high/low quality, and by eliminating subsets of cities from the sample. An iteratively weighted robust regression that penalizes outliers produces qualitatively similar results.

goods, rather than time. The reason movements in exchange rates generate relative price variability across goods in sticky price models is staggered price adjustment: firms that adjust in different periods have different relative prices. In our model, firm pricing is not synchronized because of differences in trade costs. For this measure of law of one price deviations, the model predicts a similar hump-shaped relationship between price dispersion and NER volatility: when NER volatility is sufficiently high, more volatile nominal exchange rates induce more frequent price adjustment, firms are more likely to synchronize their price changes and relative price are therefore less volatile.

To test the robustness of the results in Table 5 we next use the cross-good rather than time-series measure of price dispersion. Let q_{gt}^i be again the relative price in period t for good g for city pair i. When working with levels of relative prices, we detrend q_{gt}^i by its time-series mean as in the long-run some cities are more expensive than others: $Q_{gt}^i = q_{gt}^i - \bar{q}_g^i$. To compute the volatility of changes in relative prices, we work with $\Delta Q_{gt}^i = \Delta q_{gt}^i - \overline{\Delta q_g^i}$ to filter the data of good-specific trend growth in the relative price for a given city pair. To calculate the volatility of nominal exchange rates in any given period, we use daily exchange rate data and calculate the standard deviation of changes in daily nominal exchange rates in periods between price collections[22].

We next estimate the following system of time-series regressions, pooled over all city pairs:

$$vol(\Delta Q \text{ or } Q)_{it} = \beta_0 + \beta_1 Border_i + \beta_2 \log(\text{distance})_i + \beta_3 std(\Delta e)_{it} + \mu_t + \varepsilon_{it} \tag{13}$$

where i is an index over city pairs and t over time. Once again, our measures of volatility are the interquartile range and the standard deviation and we assume that time-specific effects are uncorrelated with nominal exchange rates. We also include city fixed effects and estimate the model again by using a random effects GLS specification. As Table 6 suggests, the results of this exercise are consistent with those in the cross-sectional regressions. Periods and city-pairs for which nominal exchange rates are less volatile suffer from less price dispersion, consistent with the predictions of the model. In the appendix (Tables A2 and A3) we also report results of an additional set of exercises in which we allow for a quadratic NER volatility term.

[22] Here we again assume that ΔQ_{gt}^i is serially uncorrelated in order to compare observations in different periods given the irregular intervals between price collections.

In all but one case, the (negative) effect of NER volatility on law of one price deviations is strongest when NER volatility is low and decreases towards zero as the volatility of NER increases. This pattern is again consistent with the predictions of the model: eventually, as NER volatility increases sufficiently, firms adjust each period and nominal shocks have no real effects, and suggests that the data we consider lies indeed on the downward-sloping portion of the RPV-NER volatility relationship predicted by the model.

B. Evidence from Macro-Data

In our second exercise, we turn to aggregate CPI data for 74 countries for which the International Financial Statistics have complete monthly CPI and nominal exchange rate data from 1973 to 1998. Earlier research using aggregate price data has mainly focused on developed countries in which the volatility of nominal exchange rates is relatively low and in which the model predicts a positive relationship between the volatility of changes in relative prices and that of changes in nominal exchange rates. For example, Engel and Rogers (2001) use consumer price indices in 12 European countries and find evidence of a positive relationship between RPV and NER volatility. In their sample monthly NER volatility ranges from 0 to 2%, which is exactly the region where the model predicts a positive relationship. Identification of a non-linear relationship between real and nominal exchange rate volatility therefore requires a broader sample of countries, some of which have experienced large NER movements.

Figure 6 presents scatter plots of RPV versus NER volatility for all 2700+ country pairs in the dataset for differences at different horizons[23]. The solid lines are the results of a locally-weighted regression in which we weigh observations using Cleveland's (1979) tricube function and employ a bandwidth length of 0.5. For monthly changes in real exchange rates the aggregate data shows no evidence of a hump-shaped relationship: real exchange rates are as volatile as nominal exchange rates even when the volatility of nominal exchange rates exceeds 30%. As the figure indicates, aggregate relative price data does show evidence of a hump-shaped relationship, but only for differences at large (24-month) horizons.

To formally test the predictions of the state-dependent model, we need a proxy that captures the degree to which the bilateral country pairs in our sample are integrated. We make use of the increasing returns-

[23]For monthly changes, nominal and real exchange rate volatility are defined as the standard deviation of $e_t - e_{t-1}$ and $q_t - q_{t-1}$ respectively. For annual changes, we calculate the standard deviation of $e_t - e_{t-12}$ and $q_t - q_{t-12}$ and similarly for all horizons.

monopolistic competition model of international trade due to Krugman (1980), and its key implication, the gravity equation, as derived for example, in Hummels (2001), to construct such a proxy. The typical prediction of this model is that imports of country i from country j are determined according to:

$$M_{ij} = kY_iY_j\tau_{ij}^{-\theta}\left(\frac{p_j}{P_i}\right)^{-\theta},$$

where, assuming symmetry across varieties of the good produced in a given country, p_j is the exporter's price (excluding the trade cost and expressed in the currency of the importer), P_i is the price index of country i, τ is the iceberg transport cost, Y_iY_j is the product of the two countries' GDP, and θ the elasticity of substitution across varieties. The state-dependent model predicts that real exchange rate volatility should decrease with $\tau_{ij}^{-\theta}$, as firms that have larger export shares face stronger incentives to adjust prices in the presence of LOP deviations. We use the bilateral trade dataset assembled by Rose (2004) in order to calculate a measure of $\tau_{ij}^{-\theta}$ by inverting the expression above:

$$\log\left(\tau_{ij}^{-\theta}\right) = \log\frac{M_{ij}}{Y_{ii}Y_j} + \theta\log\left(\frac{p_j}{P_i}\right) + \text{const}$$

Assuming that trade costs and elasticities of substitution are time-invariant, relative prices stationary, and trade costs are symmetric: $\tau_{ij} = \tau_{ji}$, our measure of the degree of tradeability of two country-pairs, $\log\left(\tau_{ij}^{-\theta}\right)$, is the time-series average of the (log) ratio of bilateral trade flows to the product of the two partners' GDP: $\log\frac{M_{ij}+M_{ji}}{Y_{ii}Y_j}$. Annual bilateral trade flow data is available for 1503 of the country-pairs in our sample and we use this restricted sample to conduct formal inference.

We relate the volatility of bilateral real exchange rates (at four different horizons) to nominal exchange rate volatility, its square, as well as our measure of tradeability, by estimating an iteratively reweighted regression which penalizes outliers by assigning them smaller weights[24]. Note, in Table 7, that tradeability enters significantly, with the negative sign predicted by the model: country-pairs that trade more suffer from smaller RER fluctuations. Coefficient estimates are small however: a one-standard deviation (1.67) increase

[24]We use biweight weights in the last rounds of the algorithm with a tuning constant of 7 (implemented as rreg in Stata.)

in this variable decreases the volatility of monthly changes in real exchange rates by only 0.02%[25]. Coefficient estimates of the effect of NER volatility on RER variability are consistent with those presented in Figure 6: although the coefficient on the squared term is negative at all horizons, it is small, in absolute value, and indicative of a hump-shaped relationship only at horizons larger than one year: the effect of NER volatility on RER variability is negative only when the volatility of annual changes in nominal exchange rates exceeds 90%, or that of bi-annual changes exceeds 137%.

The last column of Table 7 focuses on the persistence of monthly exchange rate series (as measured by their AR(1) coefficient) for the pairs of countries in our dataset. Consistent with the predictions of the model, real exchange rate fluctuations are more transitory for pairs of countries that trade more and have more volatile nominal exchange rates. The economic significance of these two effects is small however: a 10% increase in nominal exchange rate volatility reduces the autoregressive coefficient by only 0.013. The effect of tradeability is even smaller, much smaller than that predicted by the model (see Figure 4).

C. Discussion

We have shown above evidence that nominal exchange rate volatility lowers relative price variability in a volatile environment in which we have computed relative prices using a dataset of actual goods prices, evidence consistent with the predictions of a micro-founded sticky price model. The predictions of the model are, however, violated in the aggregate price data, at least at horizons shorter than one year. In particular, we have found no evidence of a hump-shaped relationship between RER and NER volatility in this environment. What explains, then, the inability of the sticky price model to explain the behavior of aggregate relative prices?

Recent research[26] has emphasized the importance of aggregation and substitution biases in generating a spurious correlation between nominal and real exchange rates. Because aggregate real exchange rates are constructed using consumption-based price indices, they may reflect, at least in the short-run, substitution by consumers towards cheaper, non-tradeable varieties of goods whose prices do not immediately respond to exchange rate fluctuations. Moreover, some of the properties (e.g. persistence) of the real exchange rate series

[25]This negative relationship is entirely absent in an ordinary least squares regression that weighs all observations equally.
[26]Burstein, Eichenbaum and Rebelo (2002), Campa and Goldberg (2002), Imbs et. al. (2002), Taylor (2001).

are heavily influenced by the behavior of its most volatile (i.e. non-tradeable) components. Crucini, Telmer and Zachariadis (2005) provide support for this idea. They use a large dataset of actual transaction prices to construct real exchange rate series immune to substitution biases and find that (at 5-year horizons) real and nominal exchange rates are weakly (in fact, negatively) correlated in a sample of European countries. Crucini and Shintani (2004) show that the degree of a good's tradeability affects the behavior of its relative prices: relative prices of tradeable goods are more transitory (half-lives less than 2 years) than the relative prices of non-tradeable goods (half-lives in excess of 4 years). Moreover, aggregation of micro-price data considerably over-estimates persistence due to the aggregation bias.

Unless aggregation bias is indeed important in generating the type of real exchange rate movements observed in the data–this is a question whose resolution requires an extensive investigation of highly disaggregated relative price data, a sticky price model can only rationalize the lack of a hump-shaped relationship between real and nominal exchange rate volatility if goods markets are sufficiently segmented and most goods trade little, either because of high trade costs of exporting goods, low elasticities of substitution, or bias in preferences towards domestic varieties. Note that at longer horizons, the aggregate data is indeed consistent with the model. At longer horizons movements in nominal exchange rates are more closely linked to fundamentals[27] and are associated with fluctuations in the marginal cost of production that increase the frequency of adjustment even for firms that trade little abroad[28]. At short horizons, when nominal exchange rate movements are disconnected from fundamentals and mostly affect the relative price of domestic relative to foreign goods, firms find optimal to revise the frequency with which they change prices in response to volatile nominal exchange rate movements only if they export sufficiently abroad. Trade frictions are therefore an important ingredient of a micro-founded sticky price model capable of replicating the relationship between nominal and real exchange rate volatility in the data.

[27] Mark (1995)

[28] Devereux and Yetman (2002) find evidence that pass-through elasticities are larger for countries subject to more volatile nominal exchange rate movements using annual data.

5. Conclusion

This paper studies the relationship between international deviations from the law of one price and nominal exchange rate volatility in the context of a state-dependent sticky price model. We show that the model generates a hump-shaped relationship between the volatility of relative prices and nominal exchange rate volatility. Firms that trade more have stronger incentives to adjust as they suffer from large shocks to relative prices induced by nominal exchange rate movements, and are willing to pay the menu costs and adjust more frequently. Tradeability therefore plays an important role in generating relative price movements in sticky price models: for values of nominal exchange rate volatility similar to those for the US[29], firms with 20% trade costs adjust three times more frequently than firms with zero costs of international trade. We find, using a dataset of actual goods prices in a relatively flexible-price environment, that pairs of cities and time-periods subject to more volatile nominal exchange rate movements tend to have smaller deviations from the law of one price, consistent with the predictions of the model. A hump-shaped relationship between real and nominal exchange rate volatility is also apparent in a dataset of CPI-based real exchange rates series for 74 countries following the Bretton-Woods period, albeit only at long horizons. At shorter horizons, real and nominal exchange rate volatility are positively correlated even for countries that have suffered for excessively large nominal exchange rate movements. Given that the model predicts a positive relationship only for firms that face large costs of international trade (or a home bias in preferences/low elasticity of substitution), we view this evidence as supportive of the view that international goods markets are far from being integrated. Frictions that limit international goods market arbitrage are therefore important in generating the type of real exchange rate movements observed in the data.

[29]The standard deviation of changes in effective nominal exchange rates for US is 1.6% in the period from 1990 to 2003. Source: IFS and author's calculations.

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Table 1: Export Share of Production

		%
1	Austria	13.5
2	Canada	14.6
3	Czech Republic	14.2
4	Denmark	16.3
5	Finland	15.3
6	France	9.9
7	Germany	11.2
8	Hungary	12.1
9	Iceland	14.5
10	Italy	10
11	Japan	4.4
12	Mexico	8.2
13	Netherlands	24.7
14	New Zealand	10.8
15	Norway	16.1
16	Poland	7.7
17	Portugal	9.3
18	Spain	6.8
19	Sweden	15.6
20	United Kingdom	10.2
21	United States	3.9
	mean	11.87

Notes: All industrial sectors. OECD STAN Indicators Database, 1993

Table 2: Moments of the Distribution of Export Shares:

	Data	Model
%		
mean	11.87	11.87
std. dev.	11.26	11.38
skewness	1.43	1.23
kurtosis	4.74	3.94

Notes: the mean of the distribution calculated based on OECD data in Table 1. The rest of the moments based on the distribution of export shares in the US Imports and Exports data at UC Davis.

Table 3: Quantitative Properties of the Model

	σ=1%	σ=3%
A. Standard Deviation of Macroeconomic Aggregates		
relative to consumption, first-differenced data		
Consumption, %	0.14	0.21
Nominal Exchange Rates, Wages	7.4	13.8
Real Exchange Rates, Wages	5.0	5.0
Money supply	4.6	12.3
Price level	4.3	12.1
B. Average Duration of Price Spells, months		
τ=1.02 firm	2.7	1.4
τ=1.20 firm	16.7	2.8
τ=1.48 firm	28.6	4.9
C. Relative Price Variability		
standard deviation of changes in relative prices, %		
τ=1.02 firm	0.53	0.42
τ=1.20 firm	0.98	1.64
τ=1.48 firm	1.10	2.27

Table 4: Summary Statistics from Micro-Price data

Absolute PPP deviations — log-deviation from cross-sectional mean (pooled over goods/time-periods)

By city

		median	iqr
St. Petersburg	(Russia)	0.30	0.57
Moscow	(Russia)	0.26	0.40
Bucharest	(Romania)	0.23	0.37
Minsk	(Belarus)	0.13	0.45
Odessa	(Ukraine)	-0.01	0.34
Kiev	(Ukraine)	-0.02	0.30
Tiraspol	(Transdniester)	-0.18	0.37
Cernovtsi	(Ukraine)	-0.20	0.52
Chisinau	(Moldova)	-0.20	0.38
Balti	(Moldova)	-0.22	0.38
Cahul	(Moldova)	-0.29	0.34
Causeni	(Moldova)	-0.35	0.40

Relative PPP deviations — Time-Series standard deviation of bilateral relative prices, Average across goods/city-pairs

		Δq	q
By good category			
	meat	0.21	0.33
	vegetables	0.32	0.59
	fruit	0.31	0.59
	other	0.16	0.27
International pairs		0.26	0.46
Intranational pairs		0.24	0.37

NER volatility (std. dev. of log-changes) 0.042
(average across city-pairs separated by national borders)

Table 5: Time-series volatility of relative prices

	1	2	3	4
	std(Δq)	**iqr(Δq)**	**std(q)**	**iqr(q)**
border	22.37	29.83	45.27	82.97
	(12.93)	(9.43)	(6.38)	(9.80)
log-distance	0.03	0.33	23.64	44.09
	(3.69)	(2.69)	(3.91)	(6.00)
std(ΔNER)	-1.90	-3.01	-0.09	-0.41
	(1.55)	(1.13)	(0.13)	(0.20)
# observations	3147	3147	3610	3610
adj. R^2	0.28	0.47	0.39	0.36

Notes:
1. Random-effects model
2. Standard errors reported
3. Coefficient estimate and standard errors on log-distance and border multiplied by 1000
4. Regressions include meat/fruit/vegetable as well as city dummies

Table 6: Cross-goods volatility of relative prices

	1	2	3	4
	std(ΔQ)	**iqr(ΔQ)**	**std(Q)**	**iqr(Q)**
border	10.72	11.34	45.10	64.83
	(4.76)	(3.91)	(3.86)	(4.40)
log-distance	-0.58	0.76	28.00	34.66
	(3.49)	(2.86)	(2.74)	(3.13)
std(ΔNER)	-0.72	-1.19	-0.72	-1.67
	(0.52)	(0.42)	(0.41)	(0.46)
# observations	6102	6102	6919	6919
adj. R^2	0.08	0.14	0.20	0.21

Notes:
1. Random effects model
2. Standard errors reported
3. Coefficient estimate and standard errors on log-distance and border multiplied by 1000
4. Regressions include city dummies

Table 7: CPI-based Real Exchange Rates

		Bilateral RER volatility				Persistence of monthly RER
		1 month	3 months	12 months	24 months	AR(1) coefficient
NER volatility		1.02	0.98	0.83	0.77	-0.13
		(0.004)	(0.01)	(0.01)	(0.01)	(0.01)
NER volatility, squared		-0.74	-0.62	-0.46	-0.28	
		(0.01)	(0.01)	(0.01)	(0.01)	
tradeability		-0.013	-0.054	-0.222	-0.385	-0.154
		(0.004)	(0.010)	(0.035)	(0.076)	(0.04)
# observations		1503	1503	1503	1503	1503
summary statistics:						
NER volatility	95th percentile	0.19	0.38	1.07	1.80	
	5th percentile	0.02	0.04	0.07	0.10	
tradeability	std. dev.	1.67				
AR(1) coefficient: monthly RER	mean	0.97				

Notes: results of a iteratively reweighted regression reported
coefficients and standard errors on "tradeability" multiplied by 100
standard errors reported in parantheses

Figure 1: Value Functions

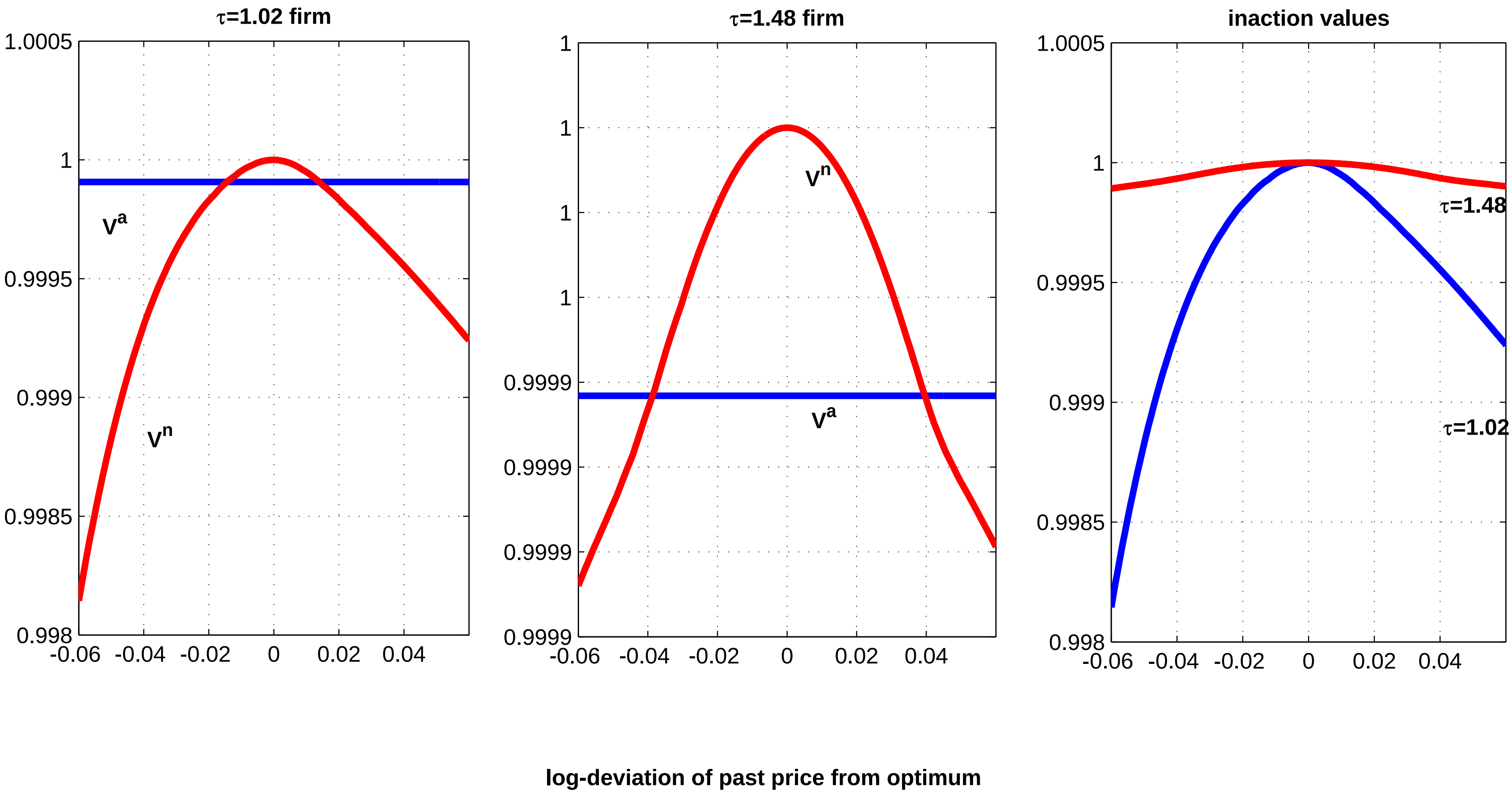


Figure 2: Price Functions and Ss bands

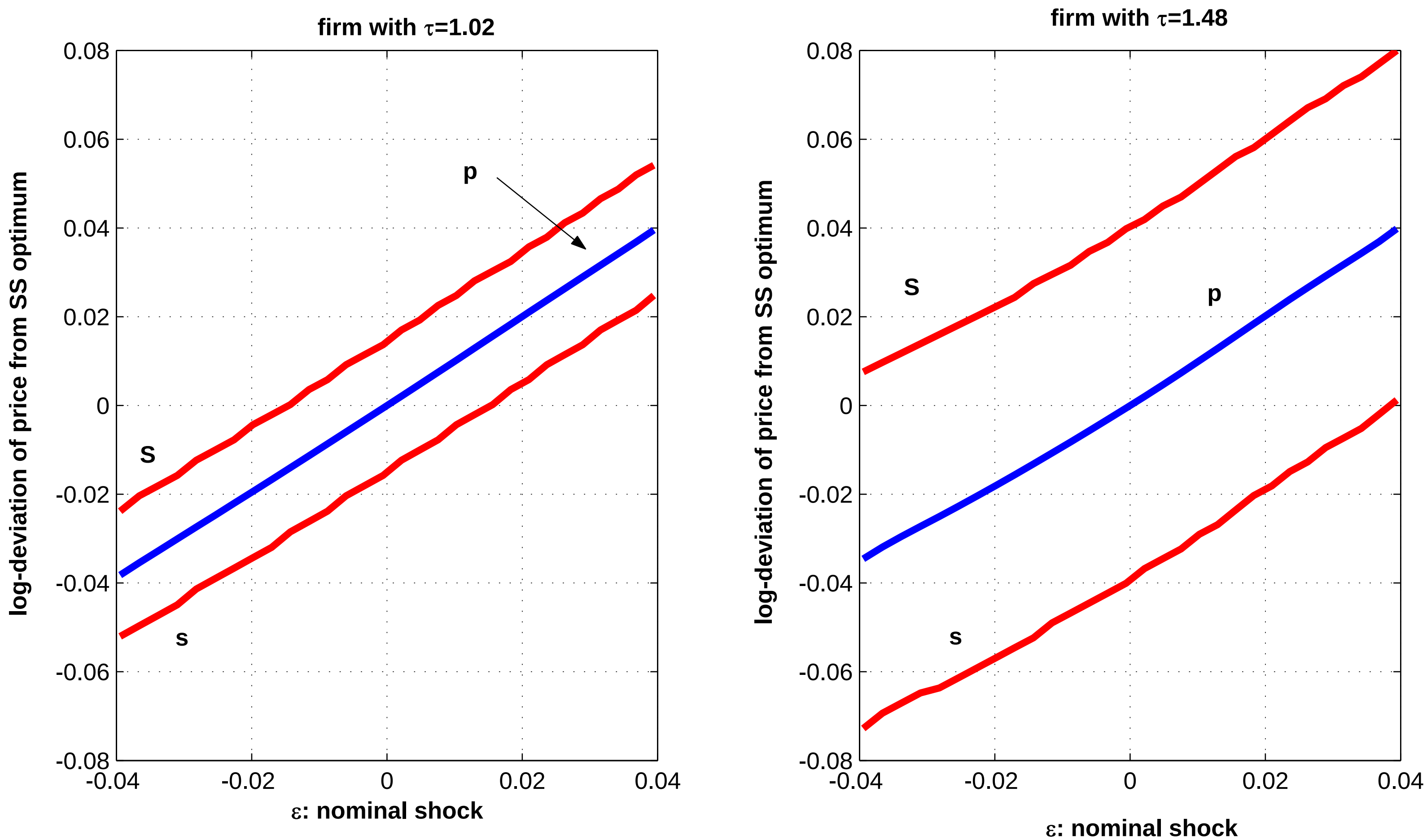


Figure 3: Relative price vs. NER volatility: Model Simulations

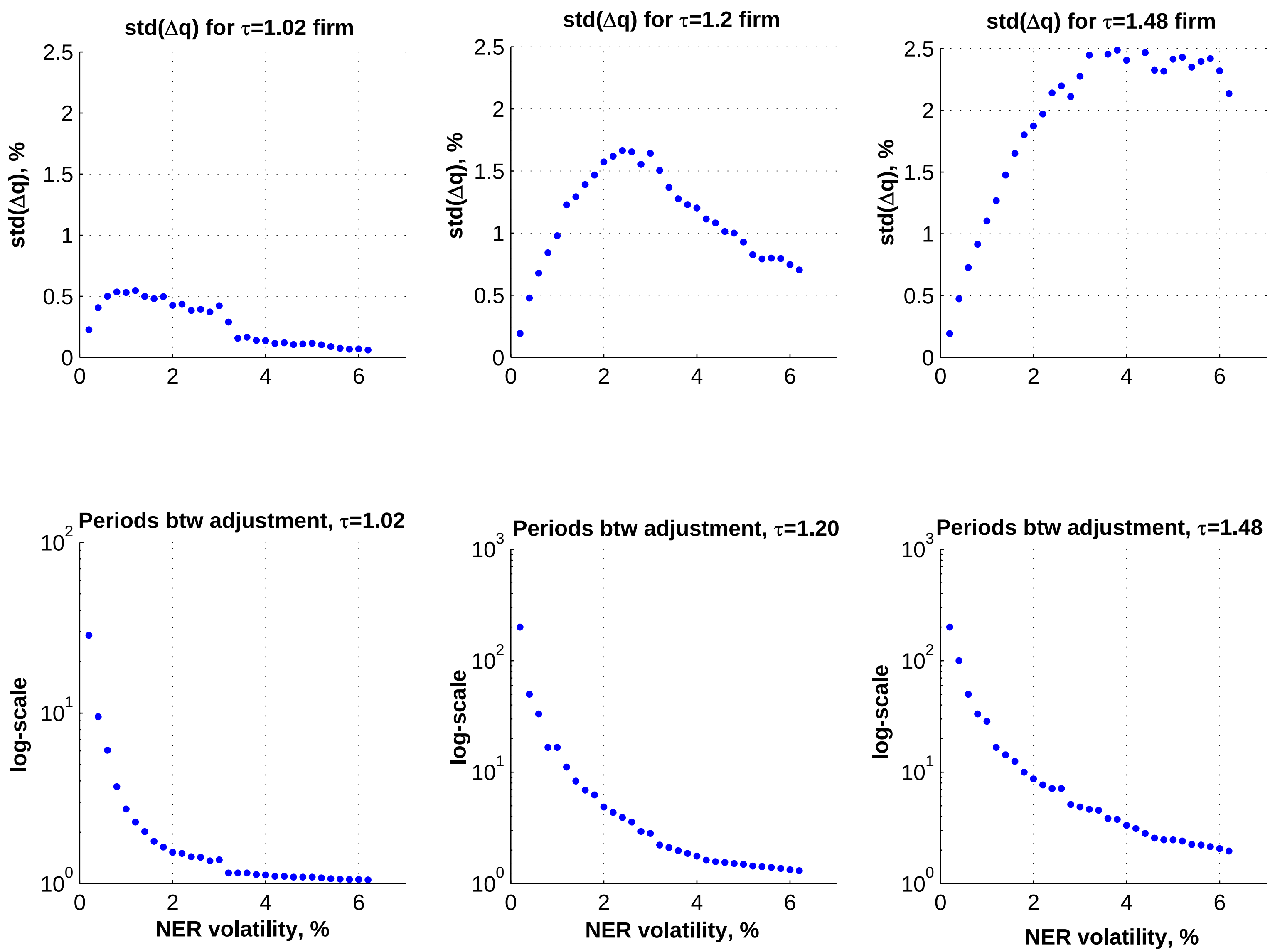


Figure 4: Other measures of LOP deviations

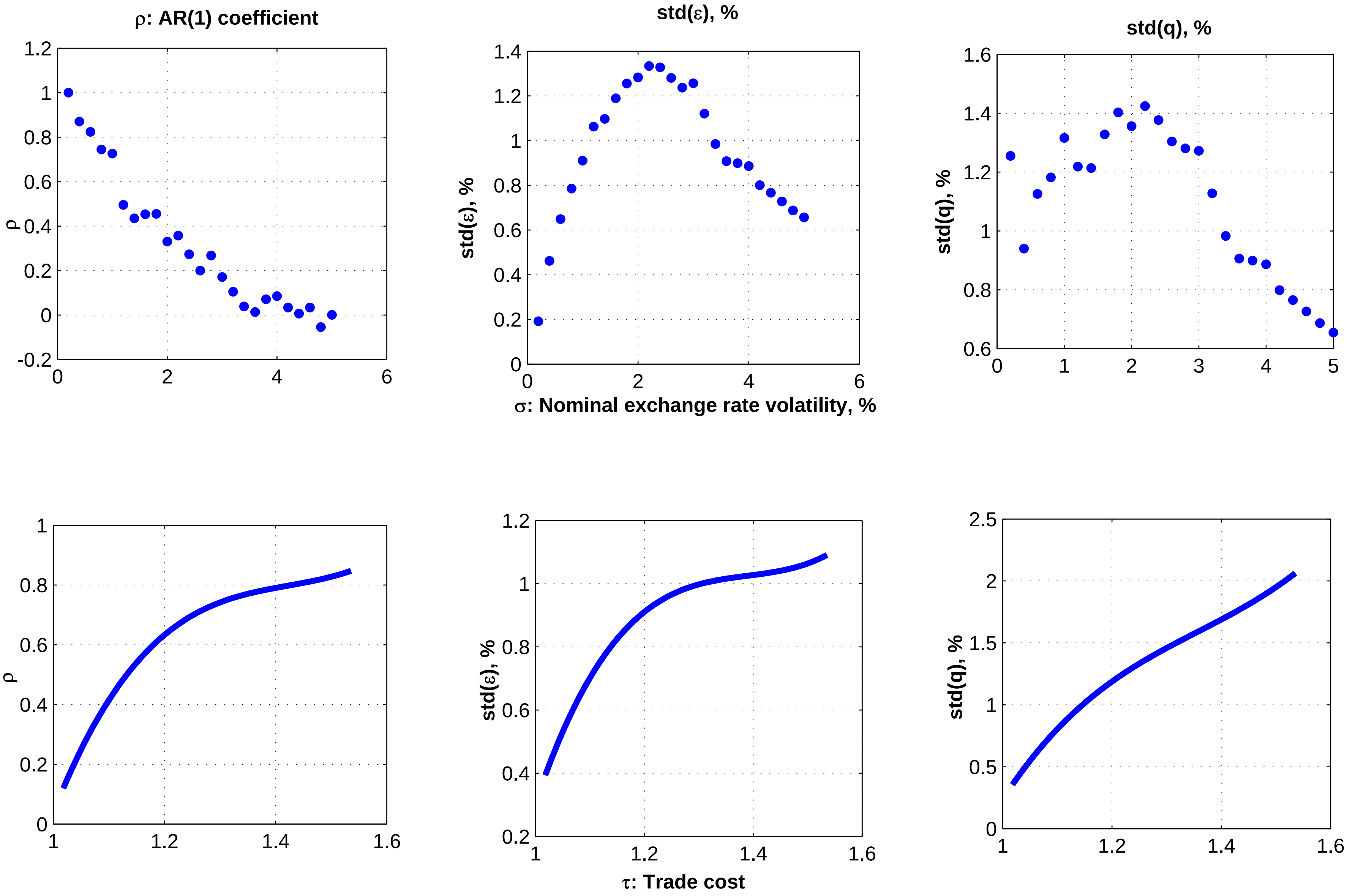


Figure 5: RER and NER volatility: Model Simulations

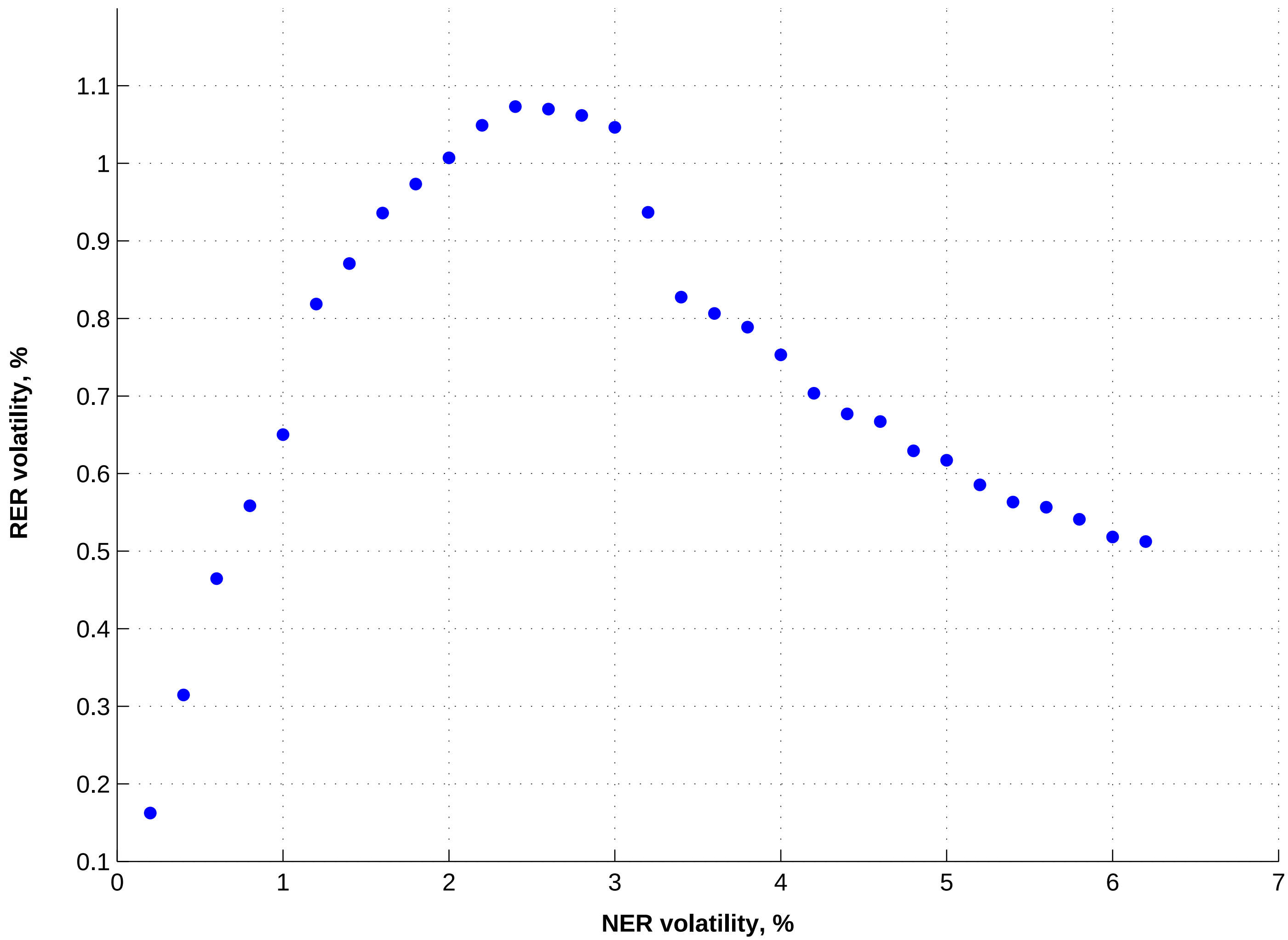
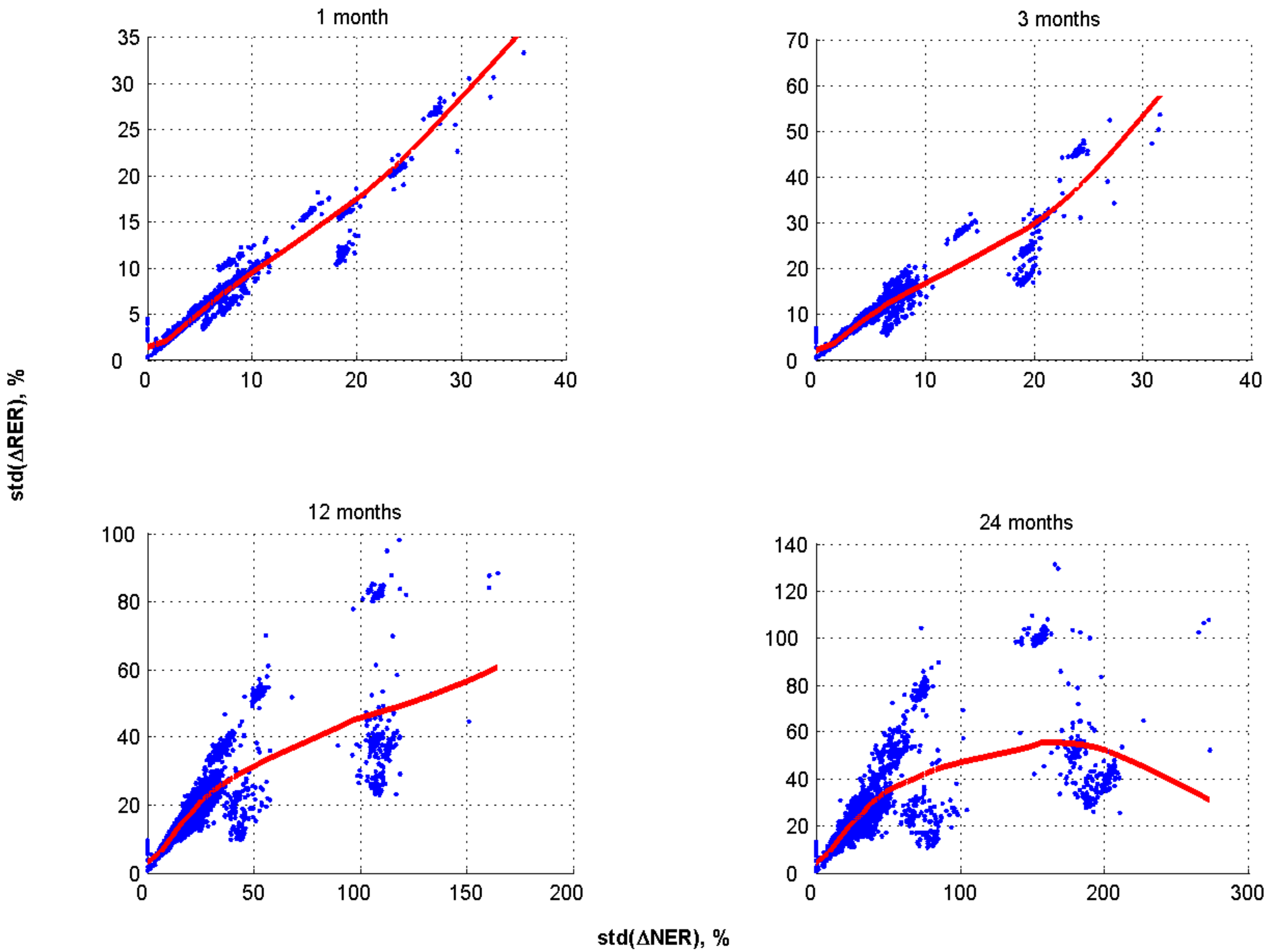


Figure 6: RER vs. NER volatility in macro data



Appendix: Solution Method

The typical approach used in solving state-dependent pricing or inventory models is the simulation technique proposed by Krusell and Smith (1997) and used by Willis(2002) and Khan and Thomas(2001) in models with non-convexities. The method involves replacing the distribution of the firms' last period's prices with a vector of its moments, postulating a linear relationship between aggregate prices and quantities and the state of the world, and solving for the unknown coefficients of these approximants by minimizing their in-sample forecast errors.

We depart slightly from the standard method, and use a solution technique free of simulations, one that draws heavily on collocation, a residual-based functional approximation method discussed at length in Miranda and Fackler (2002). A simulation-free solution technique used to solve models with heterogeneous agents was originally suggested by den Haan (1997) in the context of an uninsurable idiosyncratic risks model. The advantage of this solution method is its explicit reliance on numerical theory, as the nodes at which equilibrium conditions are solved, and the basis functions used in approximation, are chosen to ensure optimality of the approximants.

Before discussing how we solve for the aggregate prices and quantities, we turn to the solution of the firm's problem.

We solve the system of two functional equations that characterize the firm's problem in (see equation 10) using collocation. More specifically, we approximate each of the two value functions using a linear combination of N Chebyshev polynomials. To solve for the $2N$ unknown coefficients, we require that the Bellman equations hold at $2N$ nodes in the state space. This condition yields $2N$ equations we use to solve for the unknown coefficients. In addition, one needs to solve the firm's maximization problem in (10) and evaluate the expectations on the RHS of the Bellman equation by discretizing the distribution of shocks and integrating using Gaussian quadrature. We evaluate the accuracy of our solution method by calculating the difference in the two sides of the Bellman equations for points other than the collocation nodes. These residuals are small (less than 5×10^{-4} in absolute value) and are equioscillatory, a property typical of Chebyshev approximations.

We next turn to the solution of the equilibrium conditions. We have found that the most efficient approximation to the distribution of last period's prices (conditional on trade costs) is the mean of the deviations of the firms' last period's prices from their steady-state optimum, $\hat{\mu}_t = \frac{\tilde{p}_{t-1}(z)}{\tilde{p}_{ss}(z)}$ as there is little variation in this moment across firms of different types and introducing a higher moment adds little precision to our approximation to the aggregate functions. The following five functional equations (in five unknown functions: $C(s), \tilde{P}(s), \tilde{W}(s), \tilde{e}(s), \Gamma(s)$ where $s = (g, \hat{\mu})$ is the state of the world, and $\Gamma(s)$ the law of motion of μ, are sufficient to characterize the equilibrium of this economy:

$$\frac{\tilde{W}(s)}{\tilde{P}(s)} = \psi C(s)^{\kappa}$$

$$\tilde{P}(s)C(s) = 1$$

$$\frac{C(s)^{-k}}{\tilde{P}(s)} = \frac{\tilde{\Lambda}}{\tilde{e}(s)}$$

$$\tilde{P}(s) = \left(\int_0^1 \left[\left[\tilde{p}(z;s)^{1-\gamma} + (\tau(z)\tilde{e}(s)p^*(z))^{1-\gamma} \right]^{\frac{1}{1-\gamma}} \right]^{1-\theta} dz \right)^{\frac{1}{1-\theta}}$$

$$\Gamma(s) = \int_0^1 \frac{\tilde{p}_{-1}(z;s)}{\tilde{p}(z)} dz,$$

where $\tilde{p}(z;s)$ are the prices that solve the firms' problems, and $\tilde{p}(z)$ the firm's steady-state optimal price.

We solve for the aggregate functions by replacing them with a combination of Chebyshev polynomials. Given an initial guess for the coefficients on these polynomials, we solve the firm's problems and recompute a new set of aggregate quantities and prices at each state of the world used to discretize the state-space. We approximate the unknown functions using a relatively small number of basis functions (typically $K = 64$; 8 for each dimension in the aggregate state-space) but solving the model at a larger number of nodes (typically $M = 144$) and retrieving the unknown coefficients by minimizing the sum of squared residuals. For example, letting $\tilde{P}$ be a $M \times 1$ vector of home prices that satisfy the equilibrium conditions at each node, Φ be a $M \times K$ matrix of K Chebyshev polynomials evaluated at M nodes, we find the K unknown coefficients c by solving

$$\min_c \left(\tilde{P} - \Phi c \right)' \left(\tilde{P} - \Phi c \right)$$

This set of coefficients for all aggregate variables is used to re-solve the firm's problem, obtain a set of new aggregate variables at each node and calculate a new c. The convergence criterion is the norm of the difference between the last two sets of c and we typically stop when $norm < 10^{-5}$. Once the algorithm converges, our approximants produce accurate out-of sample forecasts, and explain 95% of the variation of aggregate variables in simulations of the model.

Table A1: Data Availability
(proportion available)

By City:

	city	country	
1	Chisinau	Moldova	0.97
2	Balti	Moldova	0.93
3	Cahul	Moldova	0.93
4	Causeni	Moldova	0.89
6	Tiraspol	Transdniester	0.82
7	Bucharest	Romania	0.91
8	St. Petersburg	Russia	0.80
9	Moscow	Russia	0.75
10	Kiev	Ukraine	0.81
11	Odessa	Ukraine	0.85
12	Cernovtsi	Ukraine	0.89
13	Minsk	Belarus	0.31

By good:

		grade 1	grade 2			grade 1	grade 2
	meat				**vegetables**		
1	chicken legs	0.77	0.77				
2	pork legs	0.87	0.88	16	tomatoes	0.85	0.85
3	pork w/ bones	0.87	0.87	17	sugar beet	0.87	0.87
4	beef ribs	0.85	0.85	18	potatoes	0.87	0.87
5	beef w/ bones	0.86	0.86	19	dry beans	0.76	0.76
6	beef legs	0.86	0.87	20	carrots	0.86	0.87
7	beef fillet	0.87	0.87	21	sweet pepper	0.71	0.71
8	mutton	0.70	0.69	22	garlic	0.86	0.86
9	chicken	0.76	0.76	23	onions	0.86	0.87
10	pork fillet	0.86	0.87	24	cabbage	0.87	0.87
11	pork ribs	0.85	0.86	25	cucumber	0.81	0.81
	fruit				**other**		
12	grapes	0.69	0.69	26	honey	0.87	0.87
13	apples	0.86	0.86	27	sugar	0.86	0.87
14	walnut	0.77	0.78	28	oil	0.87	0.87
15	walnut w/o shells	0.58	0.58	29	eggs	0.86	0.87

Table A2: Time-series volatility of relative prices
Non-linear Terms

	1	2
	iqr(Δq)	**iqr(q)**
std(ΔNER)	-9.76	1.23
	(2.74)	(0.83)
std(ΔNER), squared	205.51	-13.26
	(76.62)	(6.53)
Effect of std(ΔNER) on RPV when std(ΔNER) is at its		
5th percentile	-9.76	1.23
	(2.74)	(0.83)
median	-6.26	0.93
	(1.64)	(0.69)
95th percentile	-0.31	-1.59
	(1.52)	(0.61)

Notes:
1. Random-effects model
2. Standard errors reported
3. Coefficient estimate and standard errors on log-distance and border multiplied by 1000
4. Regressions include meat/fruit/vegetable as well as city dummies

Table A3: Cross-goods volatility of relative prices
Non-linear Terms

	1	2
	iqr(ΔQ)	**iqr(Q)**
std(ΔNER)	-1.93	-4.34
	(1.07)	(1.11)
std(ΔNER), squared	43.41	150.67
	(57.03)	(57.56)
Effect of std(ΔNER) on RPV when std(ΔNER) is at its		
5th percentile	-1.93	-4.34
	(1.07)	(1.11)
median	-1.78	-3.80
	(0.89)	(0.94)
95th percentile	-0.96	-0.83
	(0.51)	(0.55)

Notes:
1. Random effects model
2. Standard errors reported
3.The last 3 rows report the derivative of RPV wrt. NER volatility for 3 different values of NER volatility
4. Regressions include city dummies