International Competition and Inflation: A New Keynesian Perspective*

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Abstract

We develop and estimate an open economy New Keynesian Phillips curve (NKPC) in which variable demand elasticities give rise to changes in desired markups in response to changes in competitive pressure from abroad. A parametric restriction on our specification yields the standard NKPC, in which the elasticity is constant, and there is no role for foreign competition to influence domestic inflation. By comparing the unrestricted and restricted specifications, we provide evidence that foreign competition plays an important role in accounting for the behavior of inflation in the traded goods sector. Our estimates suggest that foreign competition has lowered domestic goods inflation about 1 percentage point over the 2000-2006 period. Our results also provide evidence against demand curves with a constant elasticity in the context of models of monopolistic competition.

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1 Introduction

An important question in macroeconomics is the extent to which global factors influence the behavior of aggregate prices. While it is widely recognized that import prices have a direct effect on consumer prices, there is less agreement about the extent to which global factors influence domestic prices. One prominent view is that the prices of U.S. domestic producers mainly depend on domestic variables, with international factors having only a limited impact. Recent work has challenged this view, arguing that the intensifying trend of global economic integration has changed the behavior of inflation, and international considerations have become an important determinant of inflation dynamics.[1]

We address this question in the context of a structural model of inflation in the spirit of Dornbusch and Fischer (1984) and Dornbusch (1985), who emphasized how variations in the desired markups of domestic firms could arise in response to changes in competitive pressures from abroad. These competitiveness effects arise in our model, because a firm faces an elasticity of demand as in Kimball (1995), which depends on its price relative to its competitors. As a result, a reduction in the prices of foreign competitors can induce domestic firms to lower their desired markups. We embed these non-constant elasticity preferences into a short-run model of inflation in which firms only infrequently re-optimize their prices due to the presence of Calvo (1983) contracts.

We derive a specification for domestic inflation that depends not only on real marginal cost, but on the prices of imported or foreign goods relative to domestic prices.[2] A parametric restriction on our specification yields the standard New Keynesian Phillips curve (NKPC) in which the elasticity of demand is constant, and there is no role for competition abroad to directly influence inflation.[3] By comparing the unrestricted and restricted versions of our model, we are

[1] For arguments in favor of view that global factors have changed the behavior of inflation, see Borio and Filardo (2006) and Rogoff (2003). For evidence that the effect has been limited, see Ihrig, Kamin, Lindner, and Marquez (2007). Ball (2006) takes an even more extreme position, arguing that there is no effect of foreign variables on U.S. inflation.

[2] Our paper is related to a longstanding literature that includes import prices in the estimation of reduced-form Phillips curves such as Gordon (1973) and Dornbusch and Fischer (1984). However, our paper differs from these earlier works by providing estimates from a structural model.

[3] Important work estimating the standard NKPC includes Galí and Gertler (1999), Galí, Gertler, and Lopez-

able to evaluate the extent to which foreign competition influences the behavior of inflation. In addition, we empirically assess the hypothesis of a constant elasticity of substitution (CES), which is often used by researchers due to its analytical convenience rather than its empirical validity.

Our methodology for estimating inflation closely parallels the present-value approach used in the empirical finance literature. To estimate our model, we use data on the prices of U.S. domestic tradable goods rather than a broader price measure. While this choice represents a departure from most of the empirical literature on inflation, it is motivated by two considerations. First, tradable prices are appropriate given our theoretical model, which focuses on the interactions between foreign and domestic producers of tradable products. Second, the behavior of domestic tradable prices should be particularly revealing regarding the influence of global factors on the domestic economy. We view substantiating that domestic tradeable prices are influenced by global factors as an important first step in building a similar case for broader measures of domestic inflation that include non-tradables.

Our results provide evidence that foreign competition has played an important role in explaining the behavior of traded goods inflation. For instance, we estimate that foreign competition, by reducing the desired markups of domestic producers, has lowered the annual inflation rate for domestic goods about 1 percentage point, on average, over the 2000-2006 period. In addition, movements in relative import prices associated with changes in foreign competition account for over $\frac{1}{3}$ of the volatility of goods price inflation over the 1983-2006 sample period.

Our benchmark estimate for the degree of nominal rigidities are consistent with firms that re-optimize prices, on average, once every 3 to 4 quarters.[4] We also find that once we account for the endogenous changes in desired markups, there is a limited, if not negligible, role for backward-looking price setting behavior in explaining the dynamics of traded goods inflation. In contrast, much of the NKPC literature including Galí and Gertler (1999) and Eichenbaum and Fischer (2007) estimate degrees of backward-looking behavior that are significantly different

Salido (2001), and Sbordone (2002).

[4] This estimate is broadly consistent with the micro evidence of Nakamura and Steinsson (2007), who find a median duration of non-sale prices of 8-11 months using prices for both consumers and producer's finished goods.

from zero. The difference in our results with these earlier papers reflects our focus on inflation for tradeable goods, which inherits a considerable degree of persistence from movements in relative import prices.

Another contribution of our work is that we show that in an open economy the variability in desired markups can be separately identified from changes in markups arising from nominal rigidities.[5] As demonstrated by Eichenbaum and Fischer (2007), in a closed economy, it is not possible to separately identify the frequency of price re-optimization from the real rigidity associated with changes in desired markups. To estimate the frequency of price adjustment in closed economy models, researchers frequently resort to calibrating the parameter governing the variation in the demand elasticity with little empirical guidance. In an open economy, relative import prices are informative about the competitive interaction between foreign and domestic firms, and can shed light on the nature of the demand curve. In this context, our estimates provide evidence against CES demand curves. In particular, we find a large and statistically significant departure from a constant elasticity of substitution, and our estimates for the demand curve are consistent with the calibrated values used in closed economy contexts by Eichenbaum and Fischer (2007), Coenen, Levin, and Christoffel (2007), and Dotsey and King (2005).

The rest of this paper proceeds as follows. Section 2 describes our open economy model with a variable demand elasticity and discusses the issue of identification. Section 3 and 4 describe our data and empirical methodology. Section 5 discusses our estimation results, while Section 6 concludes.

2 An Open Economy Model with a Variable Demand Elasticity

In this section, we describe the analytical framework that leads to the open economy New Keynesian Phillips curve that we estimate. Our framework can be viewed as part of a general equilibrium model which also includes households and the producers of non-tradable goods and

[5] Moving to an open-economy is not a sufficient condition for separately identifying these two sources of markup variations. For instance, Bouakez (2005), who also allows for changes in desired markups, is not able to separately distinguish between these two sources of markup fluctuations.

services. However, in order to help minimize model misspecification, we employ a limited information approach in estimating traded goods inflation, and only describe the part of the model that is relevant for our estimation approach. In doing so, we emphasize how the international dimension of our model allows us to separately identify the degree to which markups vary due to nominal rigidities from variation in desired markups arising from changes in competition.

2.1 Final Good Producers

At time t, an aggregate final good, A_t, is produced by perfectly competitive firms. The representative firm combines a continuum of intermediate goods produced at home and another continuum produced abroad. The firm chooses domestically-produced goods, $A_{Dt}(i)$, $i \in [0, 1]$, imported goods, $A_{Mt}(i)$, $i \in [0, 1]$, and A_t to maximize profits:

$$\max P_{At}A_t - \left[\int_0^1 P_{Dt}(i)A_{Dt}(i)di - \int_0^1 P_{Mt}(i)A_{Mt}(i)di\right], \tag{1}$$

subject to $\int_0^1 D\left(\frac{A_{Dt}(i)}{A_t}, \frac{A_{Mt}(i)}{A_t}\right) di \geq 1$. In maximizing profits, the firm takes the prices of the domestic, $P_{dt}(i)$, imported goods, $P_{Mt}(i)$, and the final good, P_{At}, as given.

For $\int_0^1 D\left(\frac{A_{Dt}(i)}{A_t}, \frac{A_{Mt}(i)}{A_t}\right) di$, we adopt the aggregator used by Gust, Leduc, and Vigfusson (2006), who extend the one discussed in Dotsey and King (2005) to an international environment. This aggregator is given by:

$$\int_0^1 D\left(\frac{A_{Dt}(i)}{A_t}, \frac{A_{Mt}(i)}{A_t}\right) di = \left[V_{Dt}^{1/\rho} + V_{Mt}^{1/\rho}\right]^{\rho} - \frac{1}{(1-\nu)\gamma_t} + 1. \tag{2}$$

In turn, V_{Dt} is an aggregator of domestically-produced goods given by

$$V_{Dt} = \int_0^1 \frac{(1-\omega)^{\rho}}{(1-\nu)\gamma_t}\left[\frac{1-\nu}{1-\omega}\frac{A_{Dt}(i)}{A_t} + \nu\right]^{\gamma_t} di, \tag{3}$$

and V_{Mt} is an aggregator of imported goods given by

$$V_{Mt} = \int_0^1 \frac{\omega^{\rho}}{(1-\nu)\gamma_t}\left[\frac{(1-\nu)}{\omega}\frac{A_{Mt}(i)}{A_t} + \nu\right]^{\gamma_t} di. \tag{4}$$

In the above, ρ influences the substitutability between domestic and foreign goods. The share parameter ω is related to the degree of home bias in preferences and can be thought of as indexing the degree of trade openness.

Our estimation strategy explicitly requires us to model an error to our structural equation for inflation. We let γ_t be an exogenous shock influencing the elasticity of substitution between varieties produced within a given country, which, as we discuss later, introduces exogenous variations in markups and in aggregate inflation. We specify that γ_t evolves according to:

$$\gamma_t = \gamma \exp(\epsilon_{\gamma t}), \tag{5}$$

where $\epsilon_{\gamma t}$ is an identically and independently distributed (*iid*) process with zero-mean and standard deviation, σ_γ. Later, we verify that once you take into account endogenous variations of the markup, this error is in fact white noise and thus makes no contribution to inflation persistence. In contrast, recent empirical applications such as Ireland (2004) and Smets and Wouters (2007) have generally assumed that the exogenous movements in the markup are serially autocorrelated.

To understand our aggregator, it is useful to abstract from the *iid* markup shock. In that case, when $\nu > 0$ and $\gamma_t = \gamma$, the elasticity of demand is variable (VES) and the (absolute value of the) demand elasticity can be expressed as an increasing function of a firm's relative price. When $\nu = 0$ and $\gamma_t = \gamma$, the demand aggregator has a constant elasticity of substitution (CES) and can be thought of as the combination of a Dixit-Stiglitz and Armington (1969) aggregators. In particular, in this case, our aggregator can be rewritten as:

$$A_t = \left[(1-\omega)A_{Dt}^{\frac{\gamma}{\rho}} + \omega A_{Mt}^{\frac{\gamma}{\rho}}\right]^{\frac{\rho}{\gamma}},$$

where $A_{Dt} = \left(\int_0^1 A_{Dt}(i)^\gamma di\right)^{\frac{1}{\gamma}}$ and $A_{Mt} = \left(\int_0^1 A_{Mt}(i)^\gamma di\right)^{\frac{1}{\gamma}}$.

As shown in Appendix A, profit maximization by the representative final good producer implies that its demand for domestic good i is given by:

$$A_{Dt}(i) = (1-\omega)\left[\frac{1}{1-\nu}\left(\frac{P_{Dt}(i)}{P_{Dt}}\right)^{\frac{1}{\gamma_t - 1}}\left(\frac{P_{Dt}}{P_{Ft}}\right)^{\frac{\rho}{\gamma_t - \rho}} - \frac{\nu}{1-\nu}\right]A_t. \tag{6}$$

In these demand curves, P_{Mt}, and P_{Dt} are price indices of domestic and imported goods given by:

$$P_{Dt} = \left(\int_0^1 P_{Dt}(i)^{\frac{\gamma_t}{\gamma_t - 1}} di\right)^{\frac{\gamma_t - 1}{\gamma_t}} \quad \text{and} \quad P_{Mt} = \left(\int_0^1 P_{Mt}(i)^{\frac{\gamma_t}{\gamma_t - 1}} di\right)^{\frac{\gamma_t - 1}{\gamma_t}}, \tag{7}$$

while P_{Ft} is a price index consisting of all the prices of a firm's competitors:

$$P_{Ft} = \left[(1-\omega) P_{Dt}^{\frac{\gamma_t}{\gamma_t - \rho}} + \omega P_{Mt}^{\frac{\gamma_t}{\gamma_t - \rho}} \right]^{\frac{\gamma_t - \rho}{\gamma_t}}. \tag{8}$$

As in Dotsey and King (2005), when $\nu \neq 0$, these demand curves have a linear term which implies that the elasticity of demand depends on a firm's price relative to the prices of its competitors, P_{Ft}.

2.2 Intermediate Good Producers

Intermediate good i is produced by a monopolistically competitive firm, whose technology is Cobb-Douglas over capital and labor. Intermediate goods producers face perfectly competitive factor input markets within a country. Capital and labor are assumed to be immobile across countries but completely mobile within a country. Thus, within a country, all firms have the same marginal cost, MC_t.

Intermediate goods producers sell their products to the consumption goods distributors, and we assume that markets are segmented so that firms can charge different prices at home and abroad (i.e., price to market). The domestic price is determined according to Calvo-style contracts. In particular, firm i faces a constant probability $1-\theta$ of being able to re-optimize its price. This probability is assumed to be independent across time, firms, and countries. If firm i can not re-optimize its price at time t, the firms resets its price based on lagged inflation as in Christiano, Eichenbaum, and Evans (2005) and Smets and Wouters (2003). In particular, $P_{Dt}(i) = \pi^{1-\delta_D} \pi_{t-1}^{\delta_D} P_{Dt-1}(i)$, where $\pi_{t-1} = P_{Dt}/P_{Dt-1}$, and the parameter $0 \leq \delta_D \leq 1$ captures the degree of indexation to past inflation. In this specification $\delta_D = 0$ corresponds to indexation to steady state inflation (π), and $\delta_D = 1$ implies full indexation to past inflation. When firm i can re-optimize in period t, it maximizes:

$$E_t \sum_{j=0}^{\infty} \xi_{t+j} \theta^j \left[I_{Dt+j} P_{Dt}(i) - MC_{t+j} \right] A_{Dt+j}(i), \tag{9}$$

taking MC_{t+j}, its demand schedule, and the indexing scheme, $I_{Dt+j} = \Pi_{h=1}^{j} \pi^{1-\delta_D} \pi_{t+h-1}^{\delta_D}$ as given. In the above, ξ_{t+j} is the stochastic discount factor with steady state value, $\beta \in (0,1)$, and E_t denotes the conditional expectations operator at date t. The first-order condition from this problem is:

$$E_t \sum_{j=0}^{\infty} \xi_{t+j} \theta^j \left[1 - \left(1 - \frac{MC_{t+j}}{I_{Dt+j} P_{Dt}(i)} \right) \epsilon_{t+j}(i) \right] A_{Dt+j}(i) = 0, \tag{10}$$

where the elasticity of demand for good i in the domestic market is:

$$\epsilon_t(i) = \frac{1}{1-\gamma_t} \left[1 - \nu \left(\frac{P_{Dt}(i)}{P_{Dt}} \right)^{\frac{1}{1-\gamma_t}} \left(\frac{P_{Dt}}{P_{Ft}} \right)^{\frac{\rho}{\rho-\gamma_t}} \right]^{-1}. \tag{11}$$

This elasticity results in a time-varying markup of the form:

$$\mu_t(i) = \frac{\epsilon_t(i)}{\epsilon_t(i) - 1} = \left[\gamma_t + \nu(1-\gamma_t) p_{Dt}(i)^{\frac{1}{1-\gamma_t}} p_{Ft}^{\frac{\rho}{\gamma_t - \rho}} \right]^{-1}, \tag{12}$$

where the lower case variables denote relative prices (i.e., $p_{Dt}(i) = \frac{P_{Dt}(i)}{P_{Dt}}$ and $p_{Ft} = \frac{P_{Ft}}{P_{Dt}}$).

To understand variations in the desired markup (i.e., the markup in the absence of price rigidities and the exogenous shock γ_t), it is useful to log-linearize this expression around a steady state in which relative prices are equal to one and write it as:

$$\hat{\mu}_t(i) = \hat{\mu}_{Dt}(i) - \varphi_\mu \hat{\gamma}_t. \tag{13}$$

where $\hat{\mu}_{Dt}(i)$ is the log-linearized desired markup and $\varphi_\mu = (\mu - 1)\frac{\gamma}{1-\gamma}$. The desired markup is given by:

$$\hat{\mu}_{Dt}(i) = - \left[\frac{\partial \epsilon(i)}{\partial p_D(i)} \frac{1}{\epsilon} \right] (\mu - 1) \hat{p}_{Dt}(i) + \left[\frac{\partial \epsilon(i)}{\partial p_M} \frac{1}{\epsilon} \right] (\mu - 1) \hat{p}_{Mt}. \tag{14}$$

The steady state markup of an intermediate good producer is given by

$$\mu = \frac{1}{\gamma + (1-\gamma)\nu} > 1, \tag{15}$$

and $\epsilon = \frac{1}{(1-\gamma)(1-\nu)}$ is the steady state demand elasticity.

According to equation (14), there are two sources of variations in desired markups. The first reflects variations arising from deviations in a firm's price relative to the prices of its domestic competitors. Variations in desired markups arising from this source depend on $\frac{\partial \epsilon(i)}{\partial p_D(i)} \frac{1}{\epsilon} = \nu \epsilon$, which is the elasticity of the elasticity with respect to a firm's relative price. For $\nu > 0$, this elasticity measures how much $\epsilon_t(i)$ rises when a firm raises its price above the prices of its domestic competitors. In that case, a firm will lower its desired markup so that its desired price does not deviate too far from those of its domestic competitors. If $\nu = 0$, then the demand curves are CES absent the markup shock, and $\frac{\partial \epsilon(i)}{\partial p_D(i)} \frac{1}{\epsilon} = 0$.

The second source of variation in a firm's desired markup arises from foreign competition. This effect depends on $\frac{\partial \epsilon(i)}{\partial p_M} \frac{1}{\epsilon} = \nu \epsilon_A \omega$, where

$$\epsilon_A = \frac{\rho}{(\rho - \gamma)(1 - \nu)} > 0 \tag{16}$$

is the elasticity of substitution between home and foreign goods. The elasticity of the elasticity with respect to the relative import price, $\frac{\partial \epsilon(i)}{\partial p_M} \frac{1}{\epsilon}$, measures how much $\epsilon_t(i)$ rises when relative import prices fall. In that case, a firm faces stiffer competition from abroad and will lower its desired markup. For $\nu = 0$, the CES case, there is no effect of foreign competitiveness on domestic markups and $\frac{\partial \epsilon(i)}{\partial p_M} \frac{1}{\epsilon} = 0$. The importance of foreign competitiveness on the desired markups of domestic firms depends on the degree of trade openness (ω) and the elasticity of substitution between home and foreign goods. International competition has a larger influence on desired markups when an economy is more open or its goods are closer substitutes with foreign goods.

Substituting out $\frac{\partial \epsilon(i)}{\partial p_D(i)} \frac{1}{\epsilon}$ and $\frac{\partial \epsilon(i)}{\partial p_M} \frac{1}{\epsilon}$, the desired markup can be expressed as:

$$\hat{\mu}_{Dt}(i) = -\frac{\Psi}{1 - \Psi} \hat{p}_{Dt}(i) + \frac{\Psi}{1 - \Psi} \frac{\epsilon_A}{\epsilon} \omega \hat{p}_{Mt}, \tag{17}$$

where the parameter Ψ reflects the variations in the desired markup associated with competition from other firms and is given by:

$$\Psi = \frac{(\mu - 1) \frac{\partial \epsilon(i)}{\partial p_D(i)} \frac{1}{\epsilon}}{1 + (\mu - 1) \frac{\partial \epsilon(i)}{\partial p_D(i)} \frac{1}{\epsilon}} = \frac{\nu \mu}{1 + \nu \mu}. \tag{18}$$

In our empirical work, we focus on estimating Ψ while calibrating the values of μ and ϵ_A. These three parameters uniquely determine the demand curve parameters discussed earlier – ρ, γ, and ν – via equations (15), (16), and (18).

2.3 Inflation Dynamics

To understand the role of variations in desired markups for inflation, we log-linearize the firm's first-order condition for price re-optimization, equation (10). As detailed in Appendix A, after some algebraic manipulation, a first-order approximation to this equation yields:

$$\hat{\pi}_t - \delta_D \hat{\pi}_{t-1} = \beta E_t[\hat{\pi}_{t+1} - \delta_D \hat{\pi}_t] + \kappa \left[(1-\Psi)\hat{s}_t + \Psi\omega \frac{\epsilon_A}{\epsilon} \hat{p}_{Mt} + \varphi \widehat{\gamma}_t \right], \quad (19)$$

$$\text{where } \kappa = \frac{(1-\beta\theta)(1-\theta)}{\theta},$$

and where $\hat{\pi}_t$ is domestic price inflation expressed as a log deviation from steady state, s_t represents real marginal cost (defined using P_{Dt}), and the composite parameter, φ, influences the sensitivity of inflation to exogenous variations in the markup and is given by $\varphi = 2\Psi - 1$.[6]

Since we allow for partial indexation to lagged inflation, current inflation is affected by inflation in the previous period. Similar to a standard new Keynesian Phillips curve (e.g., Galí and Gertler (1999)), the Calvo price setting parameter, θ, affects the responsiveness of inflation to real marginal cost through its effect on κ. However, equation (19) differs from the standard specification, since relative import prices also affect inflation. In an open economy, a domestic firm must take into account the prices of its foreign competitors on its desired markup.[7] If foreign goods become relatively less expensive, then domestic firms will respond by lowering their desired markups in order to maintain a competitive price; hence, this puts downward pressure on π_t.

[6] For simplicity, we assume that the underlying dynamic general equilibrium model implies zero growth in both the tradable and non-tradable sectors. As a result, β has the interpretation as the representative household's discount factor. If we did allow for positive and differential growth rates in these two sectors, the appropriate discount rate in equation (19) would be $\tilde{\beta} = \frac{\beta\pi}{v_A \pi_A}$, where π is the steady state inflation rate for tradable goods, π_A is the steady state inflation rate of consumer prices, and v_A is the growth rate of consumption. For U.S. data, consumption growth and the inflation differential between traded goods prices and consumer prices are small enough that taking into account these extra terms in equation (19) has a negligible effect on our results.

[7] Our specification has some similarities to Vega and Winkelried (2005), who derive a NKPC in which world prices affect domestic inflation in a small open economy with Rotemberg (1982) style price contracts and a translog demand curve. Our analysis is different from theirs, mainly because they do not explore the empirical implications of their model. Instead, we focus on the empirical relevance of foreign competition on domestic inflation, paying special attention to the issue of identification of real and nominal rigidities. Also, Sbordone (2007) analyzes how the entry of new competitors affects the slope of the NKPC in a closed economy context using the preferences of Dotsey and King (2005).

The importance of this foreign competitiveness effect on domestic inflation depends on the degree of trade openness (ω) and the import price elasticity (ϵ_A) as well as Ψ. We use Ψ to gauge the extent of the real rigidity associated with pricing complementarities between firms. A higher value of Ψ reduces the sensitivity of inflation to real marginal cost and raises the sensitivity of inflation to relative import prices.

Identifying the Real Rigidity. Equation (19) nests two important cases. With $\Psi = 0$, the CES case, there is no direct effect of international competition on domestic prices. Equation (19) is observationally-equivalent to the specification estimated by Galí and Gertler (1999) among others. Another interesting case is the one considered by Eichenbaum and Fischer (2007) in which $\omega = 0$. In this case, the domestic economy does not import foreign goods, and a domestic firm, while willing to vary its desired markup in response to domestic competition, need not be concerned with foreign competition. Accordingly, relative import prices do not affect domestic price inflation.

As discussed by Eichenbaum and Fischer (2007), one can not separately identify Ψ and θ in the closed economy ($\omega = 0$) using equation (19). As a result, many researchers opt to calibrate the value of Ψ with little empirical guidance. However, when $\omega > 0$, relative import prices are informative about the extent to which firms vary their desired markups, and it is clear from equation (19) that it is possible to jointly identify both Ψ and θ.

Coenen, Levin, and Christoffel (2007) alter the standard Calvo framework and show how one can separately identify real and nominal rigidities in a closed economy framework in which there are nominal pricing contracts of different durations. Their approach exploits the more complex dynamics between inflation and real marginal cost induced by their contracting structure and they use simulated methods of moments to estimate the parameters. Instead, we use the baseline Calvo model and exploit variation in relative import prices to provide information regarding the nature of demand curves and endogenous changes in desired markups.

Building on the work of Bergin and Feenstra (2001), Bouakez (2005) examines the ability of a sticky price model with a Kimball aggregator to explain the persistence of the real exchange rate. Unfortunately, Bouakez (2005) can not separately identify variations in desired markups from variations in markups associated with nominal rigidities. Instead, our framework allows us to separately identify these two sources of markup fluctuations.[8] Moreover, our focus is on the empirical relevance of foreign competition for domestic inflation.

2.4 Firm-Specific Capital

We now extend the analysis to incorporate firm-specific capital. To do so, we assume that the production function for intermediate good i is given by:

$$Y_t(i) = \bar{K}^{\alpha}(Z_t L_t(i))^{1-\alpha}, \tag{20}$$

where $L_t(i)$ is a firm's demand for labor and Z_t is a common technological factor. Finally, $\bar{K}$ denotes each firm's fixed stock of capital. As discussed in Coenen, Levin, and Christoffel (2007), the firm specific level of capital can be interpreted more broadly as production factors that remain fixed in the short run (such as land and overhead labor), while $L_t(i)$ can be interpreted as those factors which are variable in the short run.

Under these assumptions, firm i's marginal cost is given by:

$$MC_t(i) = \frac{Q}{1-\alpha}\frac{W_t}{Z_t}Y_t(i)^{\frac{\alpha}{1-\alpha}}, \tag{21}$$

where $Q = \bar{K}^{\frac{\alpha}{\alpha-1}}$ and $\frac{\alpha}{1-\alpha} > 0$ can be interpreted as the short-run elasticity of a firm's marginal cost to output. Because capital specificity implies that a firm's marginal cost is an increasing function of its output, it acts as another source of real rigidity. In particular, following an increase in nominal demand, a firm with the opportunity to raise its price will have a weaker incentive to do so, since the fall in the relative demand for its good reduces its marginal cost.

In the benchmark economy, a domestic producer may set different prices at home and abroad, and its pricing decision in its home market is completely independent of its pricing decision in its foreign market. With firm-specific capital, this is no longer true. A firm's export price affects

[8] Our aggregator also has the attractive feature of implying similar behavior for the desired prices of international goods as the game-theoretic models of Atkeson and Burstein (2005) and Bodnar, Dumas, and Marston (2002). See the appendix of Gust, Leduc, and Vigfusson (2006) for a discussion.

a firm's domestic price through its effect on the demand for its product, $Y_t(i)$, which alters its marginal cost. To keep our analysis tractable, we abstract from these effects and assume that the domestic firms who compete with foreign firms in the domestic market are distinct from those firms which export. While this assumption simplifies our analysis, it is also more in line with the empirical evidence discussed in Bernard, Eaton, Jensen, and Kortum (2003) than the standard assumption in which all domestic goods producers export their goods. In particular, Bernard, Eaton, Jensen, and Kortum (2003) document that in 1992 only 21 percent of manufacturing establishments were exporters.

With a firm's production equal to its domestic demand (i.e., $Y_t(i) = A_{Dt}(i)\ \forall i$), the first order condition for a firm that re-optimizes its price at date t is:

$$E_t \sum_{j=0}^{\infty} \xi_{t+j}\theta^j \left[1 - \left(1 - \frac{MC_{t+j}(i)}{V_{Dt+j}P_{Dt}(i)}\right)\epsilon_{Dt+j}(i)\right] A_{Dt+j}(i) = 0. \tag{22}$$

The log-linearized expression for domestic inflation in this case is given by:

$$\hat{\pi}_t - \delta_D\hat{\pi}_{t-1} = \beta E_t[\hat{\pi}_{t+1} - \delta_D\hat{\pi}_t] + \kappa_D \left[(1-\Psi)\hat{s}_t + \Psi\omega\frac{\epsilon_A}{\epsilon}\hat{p}_{Mt} + \varphi\widehat{\gamma}_t\right], \tag{23}$$

where $\kappa_D = \frac{\kappa}{1+\epsilon\frac{\alpha}{1-\alpha}(1-\Psi)}$, and Ψ and κ are defined as before. Comparing equation (23) with equation (19), it is clear that capital specificity does not alter the form of the NKPC but lowers the reduced-form slope coefficient since $\kappa_D < \kappa$ with $\alpha > 0$. An implication of this result is that we can not separately identify the real rigidity associated with firm-specific capital from the Calvo-price setting parameter. However, the real rigidity associated with variations in desired markups can still be separately identified and estimated provided information on either α or θ.

3 Data

We use quarterly data on inflation, marginal cost, and relative import prices from 1983-2006 to estimate our model. We focus on this sample period to help abstract from changes in monetary policy regimes. Since our theoretical analysis is for the prices of tradables, we construct on inflation measure based on goods prices (from NIPA Table 1.2.4). We also net out the prices of exported goods, reflecting that prices at home and abroad can differ.[9] The upper panel of Figure 1 plots goods inflation and inflation in the non-farm business sector from 1979-2006. The two series are positively correlated with each other (the correlation is 0.5). Goods price inflation, however, has been lower, on average, than overall inflation, as well as more volatile, particularly over the past 15 years.

To measure real marginal cost, s_t, we use data on the labor share in the non-farm business sector defined as nominal labor compensation divided by nominal output. This measure is the standard one used by Galí and Gertler (1999), Sbordone (2002), and Eichenbaum and Fischer (2007) among others.[10] The lower panel of Figure 1 plots the labor share in the non-farm business sector along with GDP goods inflation. The labor share declined throughout the first half of the 1990s, rose noticeably at the end of the 1990s, and then dropped sharply from 2001-2005.

We measure relative import prices by dividing the NIPA price deflator for non-oil imported goods by the deflator for domestic goods prices.[11] This series is shown in Figure 2 along with domestic goods inflation. Relative import prices are positively correlated with goods inflation, rising and falling with inflation in the 1980s and moving lower in the 1990s before trending upward in the past five years.

The top panel of Figure 3 shows the correlations between the current value of inflation and the leads of relative import prices. The correlation between traded goods inflation and relative import prices is above 0.3 for the first twelve leads of import prices. These correlations are consistent with our theoretical model, which links inflation to expected future values of relative import prices. The bottom panel of Figure 3 also shows these dynamic correlations between inflation and real unit labor costs. Inflation is also positively correlated with leads of real unit labor costs, as suggested by our theoretical model.

[9] We construct a Laspeyres index for domestic goods prices by netting out the index for export prices from the overall index for goods prices.

[10] A measure that corresponded more closely to costs in the tradable sector is the labor share for the manufacturing sector, but it is only available on an annual basis beginning in 1986.

[11] Ideally, we would like to have data on a basket of imported goods that matches the basket of domestically-produced goods. However, no such series are available, and instead our measure of imported prices excludes oil prices, reflecting that oil's share of imports is much larger than its share of domestic goods production. Later, as sensitivity analysis, we use an import price series that excludes other commodity prices.

4 Empirical Methodology

Our methodology closely parallels the present-value approach used in the empirical finance literature.[12] In particular, we rewrite equation (19) as a relationship between inflation and the expected discounted value of the future values of real marginal cost and relative import prices:

$$\widehat{\pi}_t = \delta_D \widehat{\pi}_{t-1} + \kappa_D \sum_{k=0}^{\infty} \beta^k E_t \left[(1-\Psi)\hat{s}_{t+k} + \Psi\omega \frac{\epsilon_A}{\epsilon} \hat{p}_{Mt+k} + \varphi \hat{\gamma}_{t+k} \right], \tag{24}$$

where $\kappa_D = \kappa$ if capital is not firm specific. As discussed in Appendix B, we choose to estimate our model using equation (24) rather than applying a generalized method of moments' (GMM) estimator to equation (23), because we found that the small-sample properties of our approach were superior.

To estimate the parameters of interest using (24), we need forecasts of real marginal cost and relative import prices, which we obtain through a VAR. Defining X_t as a vector of variables that includes s_t and p_{Mt}, our VAR in companion form can be written as:

$$X_t = AX_{t-1} + u_t, \tag{25}$$

where A is a matrix of VAR coefficients, and u_t is a vector of *iid* innovations that may be correlated with each other. With the VAR expressed in this way, we compute the forecasts of X_t using the relationship: $E_t\{X_{t+k}\} = A^k \, X_t$.

It is important to recognize that both real marginal cost and relative import prices are still endogenously and contemporaneously determined by equations (24) and (25), because the elements of the error vector, u_t, are allowed to be correlated with each other and with the markup shock in equation (24). The main appeal of our limited information approach relative to full information estimation of a DGE model is that we do not need to make strong assumptions about the auxiliary variables in X_t.[13] Such assumptions, if unwarranted, can lead to inconsistent estimates of δ_D, θ, and Ψ. In the context of an open economy, these misspecification problems can

[12] For a summary of this literature, see chapter 7 of Campbell, Lo, and MacKinley (1996). For an early application of this approach to inflation dynamics, see Sbordone (2002).

[13] Our use of the term "limited information approach" is the same as in Pagan (1979). He defines the limited information approach as estimating an Euler equation jointly with a statistical model for the endogenous right-hand side variables.

be pernicious, in part because modeling the determination of the exchange rate is a particularly challenging endeavor.[14] In our context, with a full information approach, relative import prices would depend on the particular assumptions for exchange rate determination.

For our benchmark specification of the VAR, we include only measures of real unit labor costs and relative import prices in X_t. Furthermore, we used the Box-Jenkins methodology to test down from an unrestricted VAR with longer lag length. We choose an AR(1) process for real unit labor costs and an AR(2) process for relative import prices. Later, we conduct sensitivity analysis in which we allow for feedback between unit labor costs and import prices in our VAR. For our benchmark specification of the VAR, the equation for inflation that we estimate is:

$$\hat{\pi}_t = \delta_D \hat{\pi}_{t-1} + \kappa_D \left[\frac{1-\Psi}{1-\beta\rho_s} \hat{s}_t + \omega \frac{\epsilon_A}{\epsilon} \frac{\Psi(1+\beta\rho_{M2}L)}{1-\beta\rho_{M1}-\beta^2\rho_{M2}} \hat{p}_{Mt} \right] + \epsilon_{\pi t}, \tag{26}$$

where L is the lag operator, and ρ_s is the autoregressive coefficient for unit labor costs, and ρ_{M1}, and ρ_{M2} are the autoregressive coefficients for import prices. We jointly estimate the VAR, equation (25), along with the process for inflation, equation (26).

Our estimation strategy explicitly requires us to model an error to equation (26). In our model, this error reflects *iid* shocks to the markup and is given by $\epsilon_{\pi t} = \kappa \varphi \hat{\gamma}_t$. Since the exogenous variation in markups may be correlated with unit labor costs and import prices, we use lagged variables as instruments. Our benchmark set of instruments includes one lag of traded goods inflation, one lag of real unit labor costs, two lags of relative import prices, two lags of a measure of commodity price inflation, two lags of quadratically-detrended output in the goods sector, and four lags of the spread between the 10-year Treasury note and the 3-month Treasury bill.[15]

Since it is possible that our instruments are only weakly correlated with the endogenous variables in our model, we follow Stock, Wright, and Yogo (2002) and Stock and Yogo (2004) and check for the presence of weak instruments based on the g_{min} statistic of Cragg and Donald

[14] For a comparison between the limited and full information approaches in an open economy setting, see Fukac and Pagan (2006). They conclude that the full information approach leads to biased estimates in part due to misspecification associated with the determination of the exchange rate.

[15] For our commodity price measure, we use the raw industrials spot commodity price series from the Commodity Research Bureau.

(1993). We compare this statistic against the critical values for the null hypothesis of weak instruments compiled by Stock and Yogo (2004). Finally, as robustness, we also use maximum likelihood estimation as an alternative to GMM.

Identification and Calibration. We estimate δ, θ, Ψ, as well as A, the coefficients from the VAR used to forecast unit labor costs and import prices (for our benchmark specification, the relevant elements of A are ρ_s, ρ_{M1}, and ρ_{M2}). We calibrate μ, ω, and ϵ_A. Given considerable uncertainty about the values of these parameters, we report results for alternative calibrations in our sensitivity analysis. Throughout our analysis, we set $\beta = 0.99$.

For our benchmark calibration, we choose $\mu = 1.2$, which is at the midpoint of the estimates surveyed by Rotemberg and Woodford (1995) but higher than the estimate of Basu and Fernald (1997). This value of μ implies $\epsilon = 6$. We choose ϵ_A, the elasticity of substitution between home and foreign goods, to be 1.5. This estimate is toward the higher end of estimates using macroeconomic data, which are typically below unity in the short run and near unity in the long run (e.g., Hooper, Johnson, and Marquez (2000)). Nevertheless, estimates of this elasticity following a tariff change are typically much higher.[16]

We choose ω based on the ratio of non-oil imported goods to total goods production. Because of a secular rise in the share of imports, it is difficult to determine an appropriate value for ω, which in our model corresponds to the steady state import share. For our benchmark calibration, we choose $\omega = 0.26$, which is the sample average for the 1983:Q1-2006:Q4 period that we use throughout our analysis. For the version of the model with firm-specific capital, following Coenen, Levin, and Christoffel (2007), we set $\alpha = 0.4$.

5 Estimation Results

Table 1 reports our estimates of θ, Ψ, and δ_D for the version of the model in which capital moves instantaneously across domestic firms. Table 2 shows these results for the version of the model in which capital is firm-specific. (For our estimates of the auxiliary VAR, see Appendix

[16] For a discussion of the macro estimates and estimates after trade liberalizations, see Ruhl (2005).

C.) Comparing Tables 1 and 2, it is apparent that with the exception of the parameter θ, the estimates are very similar.[17] With firm-specific capital, the second column of Table 2 shows that our estimate of θ is 0.75 for the model with a variable elasticity (VES), which implies that a firm, on average, re-optimizes its price every four quarters. In contrast, without capital specificity, $\theta = 0.82$, implying an average contract duration of over 5 quarters.[18] Since our estimate of θ is the only difference in results between these two specifications, and it is quite reasonable to believe that some production factors are firm-specific, we now proceed to focus exclusively on the model in which capital is firm-specific.

Table 2 shows that our estimate of Ψ implies a demand elasticity that is far from constant, as the value of Ψ is 0.73 and also statistically significant.[19] To understand what this estimate implies for an individual firm's demand, the upper left panel of Figure 4 plots the demand curve of good i for different values of $\frac{P_D(i)}{P_D}$ and compares it to the CES demand curve (i.e., $\Psi = 0$). As shown there, because the elasticity increases as a firm raises its price, demand falls more for the VES demand curve than the CES demand curve. With a rising elasticity of demand, the upper right panel shows that a firm will reduce its desired markup as its price rises above those of its domestic competitors.

Our estimate of Ψ implies that demand for good i falls about 14 percent in response to a 2 percent increase in a firm's price above its steady-state value and about 45 percent in response to a 5 percent increase. Correspondingly, these relative price movements are associated with 6 and 12 percentage point falls in desired markups, respectively. These estimates seem quite reasonable in contrast to the values discussed in Chari, Kehoe, and McGrattan (2000). They criticize the calibration of the demand curve in Kimball (1995), because 2 percent and 2.3 percent increases in a firm's price induce a 78 percent and 100 percent fall in demand.

The lower right panel of Figure 4 shows that a decrease in foreign prices relative to domestic

[17] As discussed earlier, this reflects that θ and α are not separately identified.

[18] For the CES demand curves, we exclude relative import prices from the instrument set, since the estimated system of equations no longer involves import prices.

[19] While we consider traded goods inflation the more appropriate measure of inflation, we also examined the impact of foreign competition on a broader measure. Specifically, we used inflation for the non-farm business sector, the dashed red line in Figure 1. We estimated a value of $\Psi = .85$ with a standard error of 0.16. Therefore, our main conclusions regarding the effect of foreign competition are applicable for this broader aggregate as well.

prices induces a domestic firm to lower its desired markup. However, a firm's desired markup varies much less in response to a change in foreign prices than in response to its own price, reflecting home bias in tradable consumption (i.e., the calibrated value of ω) and the lower elasticity between home and foreign goods (ϵ_A) than between home goods (ϵ). A 10 percent fall in the relative import price from its steady state value induces a firm to lower its desired markup only about 2 percentage points. Still, as discussed below, such movements in relative import prices and desired markups of firms are enough to have substantial effects on domestic price inflation.

The results in Table 2 also suggest that there is upward bias in the degree of indexation for the CES demand curves. In particular, there is a relatively large and significant coefficient on lagged inflation in this case. In contrast, in the unrestricted VES specification, the coefficient on lagged inflation is smaller and not statistically significant. Intuitively, with the VES demand curves, inflation is inheriting persistence from movements in relative import prices, and as a result, one does not need the partial indexation scheme to compensate. Later, we report results from a Monte Carlo exercise that substantiate this interpretation.

Table 2 reports the Ljung-Box Q-statistic at lags 1 and 4. For the VES specification with indexation, we can reject the presence of serially correlated markup shocks. For the CES specification without indexation, there is strong evidence that the markup shocks are serially correlated, suggesting that the model is misspecified. Although there appears to be less serial correlation for the CES specification with indexation, the Q-statistic at lag 4 still suggests model misspecification. In contrast, even if we omit indexation from the VES specification, the Q-statistics are consistent with no serial correlation.

Table 2 computes the g_{min} statistic, which can be used to test for the presence of weak instruments using the critical values from Stock and Yogo (2004). Based on both definitions of weak instruments discussed there, we can reject that the instrument set is weak for all four specifications shown in Table 2.[20]

[20] More specifically, we use the critical values from Tables 1 and 2 of Stock and Yogo (2004). For the VES specification with $\delta_D = 0$, we have 2 endogenous regressors ($n = 2$), and 9 instruments ($K_2 = 9$) excluding exogenous variables such as the constant. The critical value for the test based on a desired maximal bias of 5 percent relative to OLS is 18.76, and the critical value for a 10 percent desired maximal size of a 5 percent Wald

To assess the fit of our model without indexation (shown in the third column of Table 2), Figure 5 plots predicted inflation, $\hat{\pi}_t^F$, defined as:

$$\hat{\pi}_t^F = \kappa \left[\frac{1-\Psi}{1-\beta\rho_s} \hat{s}_t + \omega \frac{\epsilon_A}{\epsilon} \frac{\Psi(1+\beta\rho_{M2}L)}{1-\beta\rho_{M1}-\beta^2\rho_{M2}} \hat{p}_{Mt} \right], \quad (27)$$

using our estimates for θ and Ψ. The dashed red line in the figure shows a four-quarter moving-average of $\hat{\pi}_t^F$, while the solid black line shows a four-quarter moving average of actual inflation. Predicted inflation tracks the broad contours of observed inflation. In particular, the predicted series rises in the mid to late 1980s, trends downward with inflation in the 1990s, and rises and falls with actual inflation in the first half of this decade.

An important implication of our estimate of Ψ is that international competition plays an important role in influencing domestic inflation. To assess this role, the dashed blue line in Figure 5 plots predicted inflation for the CES specification in which $\Psi = 0$ and foreign prices do not influence the desired markups of domestic firms. As shown there, without this foreign competitiveness channel, the model fails to account for the increase in inflation in the late 1980s and its subsequent reversal in the early 1990s. More disconcerting, the CES specification overstates the level of inflation for the last seven years of our sample: the model predicts an average, annualized inflation rate of 0.5 percent from 2000-2006 compared to a slight deflation of 0.4 percent. In contrast, the average value of predicted inflation for the VES specification is very close to the observed value over this period. Since the difference between these two specifications reflects the influence of foreign competition on desired markups, our estimates suggest that foreign competition has lowered domestic goods inflation nearly 1 percentage point over the last seven years.

We can also assess the role of foreign competition for inflation dynamics by computing its contribution to the volatility of the four-quarter change in domestic goods prices. For the VES specification, as shown in Table 2 in the row labelled "$\frac{\sigma_{\pi^F}}{\sigma_\pi}$", predicted inflation accounts for nearly 75 percent of the volatility of observed inflation, with movements in relative import prices accounting for about a $\frac{1}{3}$ of actual inflation volatility. In comparison, the CES specification that allows for lagged indexation only accounts for 35 percent of the volatility of inflation.

test is 29.32.

Accordingly, these variance decompositions offer additional evidence that foreign competition has played an important role in accounting for movements in domestic goods prices.

Model Misspecification and Indexation. The results shown in Table 2 suggested that the CES model, by excluding import prices is misspecified. In particular, this specification appears to generate upward biased estimates of δ_D, the degree of indexation. We investigate this hypothesis by considering the following Monte Carlo experiment. We use the VES specification with pseudo-true values of $\Psi = 0.67$, $\delta_D = 0.16$, $\theta = 0.75$ to bootstrap 10,000 repetitions of artificial data, each with 96 observations (i.e., the length of 1983Q1-2006Q4 sample period).[21] For each Monte Carlo sample, we re-estimated the VES and CES specifications with indexation. We also repeated this exercise by generating bootstrapped data with an alternative parameterization of Ψ.

The top panels of Figure 6 plot the sampling distributions of our estimates for δ_D and θ for the first Monte Carlo experiment in which the pseudo-true value of Ψ is 0.67. The estimate of δ_D from the VES specification appears to be unbiased with the mass of the distribution narrowly concentrated around its pseudo-true value, while the estimate of θ displays some small sample bias and a bit wider distribution than implied by the asymptotic standard errors provided in Table 2. Still, these results suggest that our GMM estimator fares well in small samples.[22]

Figure 6 also shows that the misspecification bias of the CES formulation leads to estimates of δ_D and θ above their pseudo-true values. As shown in the bottom panels, the bias for δ_D and θ becomes more severe, when we increase the value of Ψ from 0.67 to 0.9 and lower θ from 0.75 to 0.67.[23] In particular, the mean estimate of δ_D is 0.47 compared to its pseudo-true value of 0.16 when $\Psi = 0.9$. This upward bias arises, because the misspecification associated with the omitted import price variable gives rise to serially correlated markup shocks. As a result, the

[21] These values for Ψ, δ_D, and θ correspond to the GMM estimates of the VES specification using a smaller set of instruments. In particular, we eliminated the lags of commodity price inflation and the interest rate spread from the instrument set to facilitate the Monte Carlo exercise. The results of this smaller instrument set are shown in Table 5. For additional details regarding the Monte Carlo experiment, see Appendix B.

[22] See Podivinsky (1999) for a review of the literature using Monte Carlo simulations to evaluate the small sample properties of GMM.

[23] This alternative parameterization holds fixed the value of $\kappa_D(1-\Psi)$, the reduced-form slope coefficient of real unit labor cost in equation (23).

estimate of δ_D rises above its pseudo-true value to help soak up this residual autocorrelation. Thus, an econometrician, who ignored the influence of foreign competition on inflation, may mistakenly conclude that lagged indexation plays an important role in explaining inflation.

Comparison with the Literature. As discussed earlier, Ψ can be used to gauge the degree of real rigidities associated with variations in desired markups arising from domestic competition. From equation (18), we can see that Ψ depends on both the steady state demand elasticity or markup, and the elasticity of the demand elasticity with respect to a firm's price, $\frac{\partial\epsilon(i)}{\partial p_D(i)}\frac{1}{\epsilon}$. It is therefore a useful metric to compare our estimates with calibrated values of the Kimball (1995) preferences used in the literature.

Table 3 shows our estimated value for Ψ as well as the elasticity of the elasticity with respect to a firm's price. Although our estimates suggest that those discussed in Chari, Kehoe, and McGrattan (2000) are high, a number of researchers use calibrations that are validated by our results. In contrast, Dossche, Heylen, and den Poel (2007) use scanner data from a euro-area supermarket chain to argue that most calibrations of the Kimball (1995) aggregator impose too high a value of $\frac{\partial\epsilon(i)}{\partial p_D(i)}\frac{1}{\epsilon}$, as the median estimate for the goods they consider is only 0.8. However, given that they estimate a demand elasticity with a (net) markup of 250%, their implied estimate of Ψ is 0.67, quite close to our estimate. In our view, Ψ is the relevant metric for comparing results, since $\frac{\partial\epsilon(i)}{\partial p_D(i)}\frac{1}{\epsilon}$ is not a sufficient statistic for describing the demand curve or the degree of variation in desired markups.[24] Our estimate is also much lower than Bouakez (2005), who estimates $\frac{\partial\epsilon(i)}{\partial p_D(i)}\frac{1}{\epsilon}$ by calibrating the Calvo price setting parameter to be consistent with 4 quarter contracts.

Our results are also related to Batini, Jackson, and Nickell (2005), who estimate an open economy NKPC for the United Kingdom in which foreign prices affect inflation due to both variations in desired markups and the presence of imported intermediate goods. In contrast to our results, they find that their measure of external competitiveness does not have a statistically significant role in explaining the variation in inflation. However, there are a number of important

[24] For example, for a very high markup, such as the one estimated by Dossche, Heylen, and den Poel (2007), the variation in the desired markup can be substantial without much variation in the demand elasticity.

differences in their paper from ours. Most notably, they adopt an *ad hoc* specification for variations in desired markups.

In our purely forward-looking model, we estimate a value of θ, which implies an average contract duration of four quarters. This estimate is broadly consistent with the micro evidence of Nakamura and Steinsson (2007), who find a median duration of non-sale prices of 8-11 months using prices for both consumers and producer's finished goods.[25] Our estimates are also broadly in line though slightly higher than those of Coenen, Levin, and Christoffel (2007) and Eichenbaum and Fischer (2007), who incorporate both VES demand curves and firm-specific capital into New Keynesian Phillips curves.

Our estimate of an insignificant degree of indexation are in line with two recent papers by Ireland (2004) and Coenen, Levin, and Christoffel (2007). Ireland (2004) finds no role for indexation in a closed economy model when he allows for serially autocorrelated markup shocks. In contrast, we use *iid* markup shocks to show that once we allow for endogenous variations in markups, lagged indexation is not significant. Coenen, Levin, and Christoffel (2007) estimate a closed economy Phillips curve and argue that backward-looking price-setting is not needed to explain aggregate inflation in the context of a stable monetary policy regime. Contrary to their analysis, our results do not hinge on the use of a dummy variable to account for a change in the U.S. monetary policy regime occurring in 1991.[26]

5.1 Alternative Model Calibrations

Table 4 considers the sensitivity of our estimates to the calibrated values of ω, ϵ_A, and μ. As a point of comparison, Table 4 also reports our estimates from the VES specification with firm-specific capital and lagged indexation using the benchmark calibration of these parameters.

The parameter ω determines the share of imports in goods production. For our benchmark

[25] The findings of Nakamura and Steinsson (2007) are also in line with earlier micro studies surveyed in Taylor (1999). In contrast, Bils and Klenow (2004) find a much higher frequency of price adjustment using micro data on consumer prices. The lower frequency of price changes in Nakamura and Steinsson (2007) largely reflects that they exclude temporary sales in measuring price changes, while Bils and Klenow (2004) include sales.

[26] If we included the 1991 dummy into our analysis, the estimates of θ and δ_D would fall and the overall fit of the model would improve. However, we take a more conservative approach and exclude the dummy from our analysis.

calibration, we choose $\omega = 0.26$, the sample average of the ratio of nonfuel goods imports to goods production from 1983-2006. As an alternative, we set $\omega = 0.35$, its value in 2004. With $\omega = 0.35$, our estimate of Ψ declines slightly from 0.73 in our benchmark calibration to 0.66.

The third column of Table 4 shows the effect of lowering the import price elasticity, ϵ_A, from its benchmark value of 1.5 to 0.5, a value consistent with short-run estimates. In this case, the estimate of Ψ rises to 0.84, still within the 90% confidence interval of the benchmark model. Alternatively, an increase in ϵ_A to 2 lowers our estimate of Ψ to 0.56. This fall in Ψ, however, does not necessarily imply that foreign competition has a smaller effect on the desired markups of domestic firms. In particular, for a given value of Ψ, a higher import price elasticity raises the responsiveness of domestic firms' desired markups to foreign prices. The final column of Table 4 shows the estimation results using a markup of 10%, a value in line with the estimates of Basu and Fernald (1997). In this case, our benchmark estimate for Ψ rises from 0.73 to 0.82.

Although the point estimates for the parameters governing the nominal and real rigidities are somewhat dependent on our calibration choices, the qualitative results are not. For a wide set of parameter choices, as evidenced in Table 4, we find that nominal contracts last on average between 3 and 4 quarters, and that foreign competition by inducing changes in the desired markups plays a significant role in explaining the dynamic of inflation.

5.2 Alternative Instruments, Data, and Estimation Procedures

Table 5 compares the structural estimates for the VES model assuming firm-specific capital with a number of alternatives. The first alternative examines the estimation results when we use a smaller instrument set. In particular, in this case, we include two lags of relative import prices, two lags of inflation, and one lag of real unit labor costs. While the point estimate for θ does not change, we find that the degree of real rigidities is modestly lower in this case.

In our benchmark specification for forecasting unit labor costs and import prices, we ignored any feedback between these variables by considering separate AR processes for these variables. In the third column, we consider an alternative forecasting process in which these variables are modeled as a VAR(2): $X_t = AX_{t-1} + u_t$ with $X'_t = [\hat{s}_t\ \hat{p}_{Mt}\ \hat{s}_{t-1}\ \hat{p}_{Mt-1}]$ and $u'_t = [u_{st}\ u_{p_Mt}\ 0$

$0]'$.[27] For the VAR, it is useful to define the vectors $e_1' = [1\ 0\ 0\ 0]$ and $e_2' = [0\ 1\ 0\ 0]$ to pick out the the first element, $\hat{s}_t$, and the second element, $\hat{p}_{Mt}$, from the vector X_t. Using the previous definitions, we can solve equation (23), the second order difference equation governing inflation dynamics and rewrite it as:

$$\widehat{\pi}_t = \delta_D \widehat{\pi}_{t-1} + \kappa_D[(1-\Psi)e_1'(I-\beta A)^{-1}X_t + \omega\frac{\epsilon_A}{\epsilon}\Psi e_2'(I-\beta A)^{-1}X_t] + \epsilon_{\pi t}. \quad (28)$$

The third column of Table 5 reports the estimation results under this alternative forecasting model, and Appendix C reports our estimate of A. As shown in Table 5, the estimate of Ψ is somewhat larger; however, overall, the restrictions we place on the forecasting model do not appreciably alter the estimates vis-à-vis our benchmark model.

To be consistent with our theoretical model, it would be ideal to have data on a basket of imported goods that matches the basket of domestically-produced goods. However, no such series are available, and instead our benchmark measure of imported prices excludes oil prices, reflecting that oil's share of imports is much larger than its share of domestic goods production. Since the same argument is applicable to other commodities, the third column of Table 5 presents results based on a measure of import prices that excludes oil, materials, and industrial goods. As shown there, our estimates only change slightly from the benchmark case; the most noticeable difference is for the indexation parameter, which equals 0.18 instead of the benchmark estimate of 0.1.

The last column of Table 5 presents results from estimating our system of equations (i.e., the structural inflation equation and the two AR processes for unit labor costs and import prices) using maximum likelihood estimation (MLE). Despite this different estimation strategy, the results are remarkably similar to our GMM estimates. We can also use our MLE estimates to test whether the restrictions implied by our structural model with VES demand are rejected by the data. To do this, we estimated an unrestricted VAR of order 2 that includes traded goods inflation, real unit labor costs, and relative import prices. Using a likelihood ratio test, we fail to reject the restrictions implied by our theoretical model.[28]

[27] In this case, we use the benchmark set of instruments.

[28] The log-likelihood of the unrestricted VAR is 1092.85, while the log-likelihood of the restricted VAR implied

Overall, we conclude that our results are robust to some broad changes in the instrument set, the forecasting process, the import price series, and the estimation method. In particular, in all cases, these estimates suggest that movements in relative import prices have been an important factor in accounting for traded goods inflation; once we account for endogenous markup variation, there is a limited, if not negligible, contribution of indexation to traded goods inflation.

6 Conclusions

In this paper, we developed a structural model and showed that foreign competition has played an important role in accounting for the behavior of goods inflation through changes in desired markups of domestic firms. In particular, we found that foreign competition has lowered domestic goods inflation by nearly 1 percentage point over the 2000-2006 period. In addition, our results also provided evidence in favor of demand curves which lead to endogenous variations in markups. In contrast to previous work, we found that an inflation specification without backward-looking behavior performed reasonably well in explaining movements in traded goods inflation due to endogenous changes in desired markups.

Although we view this as an important step in understanding how international factors influence domestic prices, goods production is about $\frac{1}{3}$ of overall GDP. A rough estimate would suggest that foreign competition has lowered overall GDP inflation about $\frac{1}{3}$ of 1 percentage point over the 2000-2006 period. However, this estimate does not take into account any interaction between the traded and non-traded sectors, which may magnify these effects. We leave the exploration of this issue to future research.

by our model is 1084.68. Hence, the likelihood ratio statistic is 2*(1092.85-1084.68) = 16.34. Our theoretical model imposes 12 restrictions, and the 10 percent probability of a χ^2-distribution with 12 degrees of freedom is 18.55.

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Table 1: Estimates of Open Economy Calvo Model (Capital Not Firm Specific)[a,b]
1983:Q1 – 2006:Q4

	VES with indexation	VES without indexation	CES with indexation	CES without indexation
θ	0.82	0.82	0.92	0.91
	(0.03)	(0.03)	(0.02)	(0.02)
Ψ	0.73	0.72	0	0
	(0.11)	(0.12)	–	–
δ_D	0.10	0	0.34	0
	(0.08)	–	(0.08)	–
$\frac{\sigma_{\pi F}}{\sigma_\pi}$	0.73	0.76	0.34	0.28
$\frac{\sigma_\pi^{pm}}{\sigma_\pi}$	0.33	0.40	0.00	0.00
Q-Statistic(1)	0.20	2.22	0.38	11.43
	[0.65]	[0.14]	[0.54]	[0.00]
Q-Statistic(4)	3.93	6.87	11.34	44.33
	[0.42]	[0.14]	[0.02]	[0.00]
g_{min}	71.64	62.03	72.30	62.33

[a]Standard errors are reported in parentheses. A dash in lieu of a standard error indicates that we restricted the corresponding parameter. Q-statistic refers to the Ljung-Box test for serial correlation of $\epsilon_{\pi t}$ at lags 1 and 4. Probability values of Q-statistics are reported in brackets. $\frac{\sigma_{\pi F}}{\sigma_\pi}$ refers to the ratio of the volatility of predicted inflation to the volatility of actual inflation, and $\frac{\sigma_{\pi P_m}}{\sigma_\pi}$ refers to the contribution of the relative import price to inflation volatility.
[b]The estimated inflation equation from Section 4 is:

$$\hat{\pi}_t = \delta_D \hat{\pi}_{t-1} + \kappa \left[\frac{1-\Psi}{1-\beta\rho_s} \hat{s}_t + \omega \frac{\epsilon_A}{\epsilon} \frac{\Psi(1+\beta\rho_{M2}L)}{1-\beta\rho_{M1}-\beta^2\rho_{M2}} \hat{p}_{Mt} \right] + \epsilon_{\pi t},$$

where $\kappa = \frac{(1-\beta\theta)(1-\theta)}{\theta}$.

Table 2: Estimates of Open Economy Calvo Model (Firm-Specific Capital)[a,b]
1983:Q1 – 2006:Q4

	VES with indexation	VES without indexation	CES with indexation	CES without indexation
θ	0.75	0.74	0.83	0.80
	(0.03)	(0.03)	(0.03)	(0.04)
Ψ	0.73	0.72	0	0
	(0.11)	(0.12)	–	–
δ_D	0.10	0	0.34	0
	(0.08)	–	(0.08)	–
$\frac{\sigma_{\pi^F}}{\sigma_\pi}$	0.73	0.76	0.34	0.28
$\frac{\sigma_\pi^{p_m}}{\sigma_\pi}$	0.33	0.40	0.00	0.00
Q-Statistic(1)	0.20	2.22	0.38	11.43
	[0.65]	[0.14]	[0.54]	[0.00]
Q-Statistic(4)	3.93	6.87	11.34	44.33
	[0.42]	[0.14]	[0.02]	[0.00]
g_{min}	71.64	62.03	72.30	62.33

[a]Standard errors are reported in parentheses. A dash in lieu of a standard error indicates that we restricted the corresponding parameter. Q-statistic refers to the Ljung-Box test for serial correlation of $\epsilon_{\pi t}$ at lags 1 and 4. Probability values of Q-statistics are reported in brackets. $\frac{\sigma_{\pi^F}}{\sigma_\pi}$ refers to the ratio of the volatility of predicted inflation to the volatility of actual inflation, and $\frac{\sigma_{\pi P_m}}{\sigma_\pi}$ refers to the contribution of the relative import price to inflation volatility.
[b]The estimated inflation equation is:

$$\hat{\pi}_t = \delta_D \hat{\pi}_{t-1} + \kappa_D \left[\frac{1-\Psi}{1-\beta\rho_s} \hat{s}_t + \omega \frac{\epsilon_A}{\epsilon} \frac{\Psi(1+\beta\rho_{M2}L)}{1-\beta\rho_{M1}-\beta^2\rho_{M2}} \hat{p}_{Mt} \right] + \epsilon_{\pi t},$$

where $\kappa_D = \frac{(1-\beta\theta)(1-\theta)}{\theta\left[1+\epsilon\frac{1-\alpha}{\alpha}(1-\Psi)\right]}$.

Table 3: Comparison of Benchmark Estimates and Calibrated Demand Curves in the Literature

	ϵ	μ	$\frac{\partial\epsilon(i)}{\partial p_D(i)}\frac{1}{\epsilon}$	Ψ
Benchmark Estimates	6	1.2	13.67	0.73
Chari, Kehoe, and McGrattan (2000)	10	0.11	300	0.97
Coenen, Levin, and Christoffel (2007)	5–20	0.05-0.25	10–33	0.47–0.89
Eichenbaum and Fischer (2007)	11	0.1	10–33	0.5–0.77
Dossche, Heylen, and den Poel (2007)[b]	1.4	2.5	0.8	0.67
Dotsey and King (2005)	10	0.11	60	0.87
Gust, Leduc, and Vigfusson (2006)	6	0.2	18.30	0.78
Bouakez (2005)	11	10	216	0.96

[b]Median estimated demand elasticity and curvature from Table 5.

Table 4: Estimates of VES Specification Under Alternative Calibrations

	Benchmark*	$\omega = 0.35$	$\epsilon_A = 0.5$	$\epsilon_A = 2$	$\mu = 0.1$
θ	0.75	0.77	0.72	0.77	0.69
	(0.03)	(0.03)	(0.04)	(0.03)	(0.04)
Ψ	0.73	0.66	0.84	0.56	0.82
	(0.11)	(0.14)	(0.08)	(0.14)	(0.08)
δ_p	0.10	0.10	0.10	0.10	0.10
	(0.08)	(0.08)	(0.08)	(0.08)	(0.08)

*The benchmark column refers to the model including firm-specific capital.
The estimated inflation equation is:

$$\hat{\pi}_t = \delta_D \hat{\pi}_{t-1} + \kappa_D \left[\frac{1-\Psi}{1-\beta\rho_s}\hat{s}_t + \omega\frac{\epsilon_A}{\epsilon}\frac{\Psi(1+\beta\rho_{M2}L)}{1-\beta\rho_{M1}-\beta^2\rho_{M2}}\hat{p}_{Mt}\right] + \epsilon_{\pi t},$$

where $\kappa_D = \frac{(1-\beta\theta)(1-\theta)}{\theta\left[1+\epsilon\frac{1-\alpha}{\alpha}(1-\Psi)\right]}$.

Table 5: Estimates of VES Specification Under Alternative Assumptions[a]

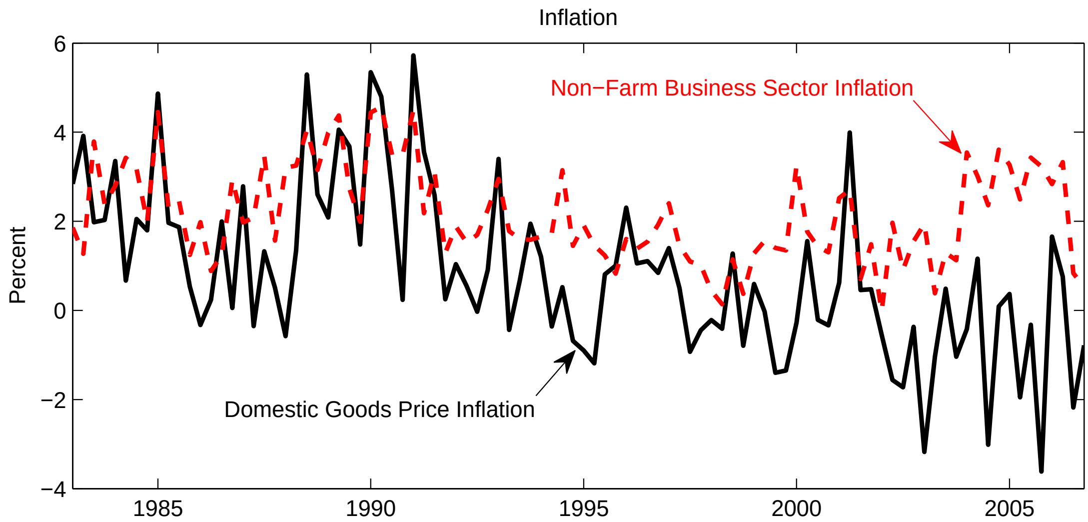
	Benchmark VES[b]	Smaller Instrument Set[c]	VAR(2) Forecasting Model[d]	Alternative Import Price Series[e]	Maximum Likelihood
θ	0.75	0.75	0.72	0.75	0.77
	(0.03)	(0.04)	(0.05)	(0.03)	(0.05)
Ψ	0.73	0.67	0.82	0.76	0.71
	(0.11)	(0.17)	(0.16)	(0.12)	(0.18)
δ_D	0.10	0.16	0.11	0.18	0.11
	(0.08)	(0.09)	(0.08)	(0.09)	(0.08)
$\frac{\sigma_{\pi F}}{\sigma_\pi}$	0.73	0.77	0.69	0.74	0.74
Q-Statistic(1)	0.20	0.00	0.29	0.12	0.15
	[0.65]	[0.99]	[0.59]	[0.72]	[0.70]
Q-Statistic(4)	3.93	3.43	3.82	5.56	3.37
	[0.42]	[0.49]	[0.43]	[0.23]	[0.50]
g_{min}	71.64	156.81	71.03	84.07	–

[a]Standard errors are reported in parentheses. Q-statistic refers to the Ljung-Box test for serial correlation of $\epsilon_{\pi t}$ at lags 1 and 4, respectively. Probability values of the Q-statistics are reported in brackets. $\frac{\sigma_{\pi F}}{\sigma_\pi}$ refers to the ratio of the volatility of predicted inflation to the volatility of actual inflation.
[b]The benchmark column refers to the model including firm-specific capital. The estimated inflation equation is:

$$\hat{\pi}_t = \delta_D \hat{\pi}_{t-1} + \kappa_D \left[\frac{1-\Psi}{1-\beta\rho_s} \hat{s}_t + \omega \frac{\epsilon_A}{\epsilon} \frac{\Psi(1+\beta\rho_{M2}L)}{1-\beta\rho_{M1}-\beta^2\rho_{M2}} \hat{p}_{Mt} \right] + \epsilon_{\pi t},$$

where $\kappa_D = \frac{(1-\beta\theta)(1-\theta)}{\theta\left[1+\epsilon\frac{1-\alpha}{\alpha}(1-\Psi)\right]}$.
[c]The benchmark instrument set includes two lags of relative import prices, commodity price inflation, and quadratically-detrended output in the traded goods sector, 4 lags of the interest rate spread, and one lag of real unit labor costs and inflation. The smaller instrument set includes two lags of relative import prices, two lags of inflation, and one lag of real unit labor costs.
[d]The benchmark system includes an AR(1) process for real unit labor costs and an AR(2) for relative import prices. The VAR(2) model refers to replacing these part of the benchmark system with an unrestricted VAR(2) model for real unit labor costs and relative import prices.
[e]The benchmark relative import price series is the NIPA price deflator for non-oil imported goods divided by domestic goods prices. The alternative series excludes import prices of industrial goods and materials in addition to fuel prices.

Figure 1: Tradable Goods Inflation and Unit Labor Costs, 1983-2006



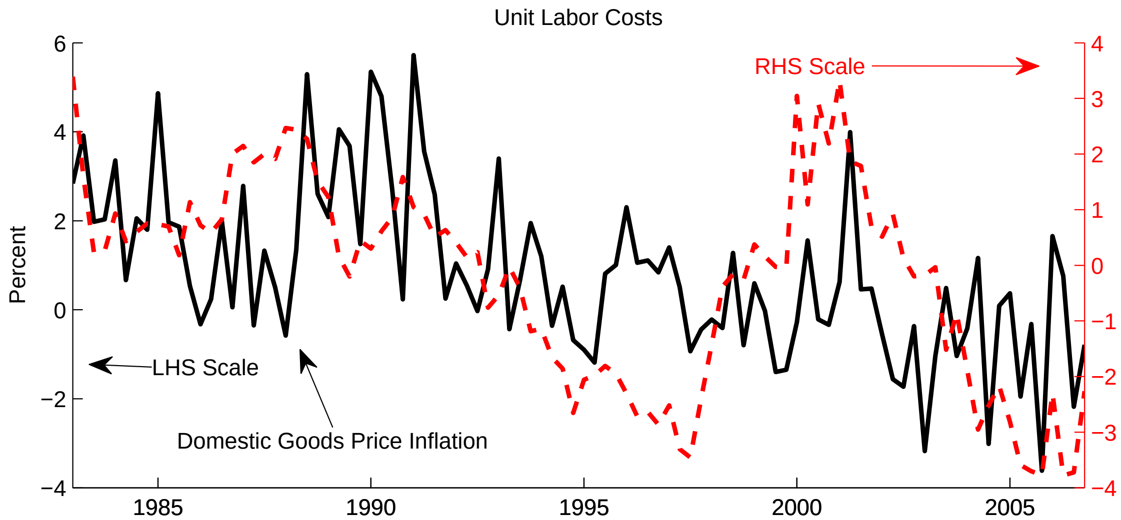


Figure 2: Relative Import Prices, 1983-2006

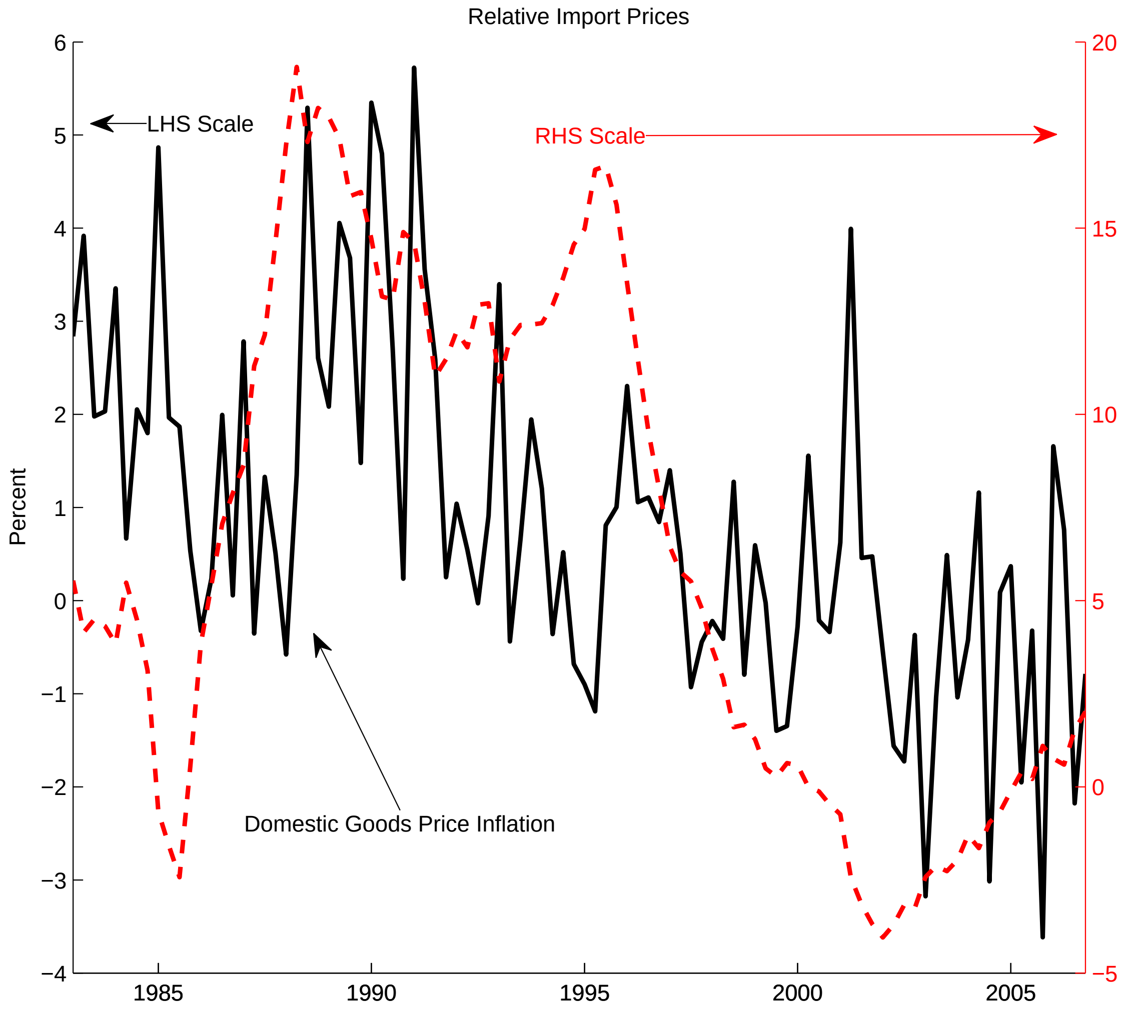


Figure 3: Cross-Correlogram for Inflation, 1983-2006

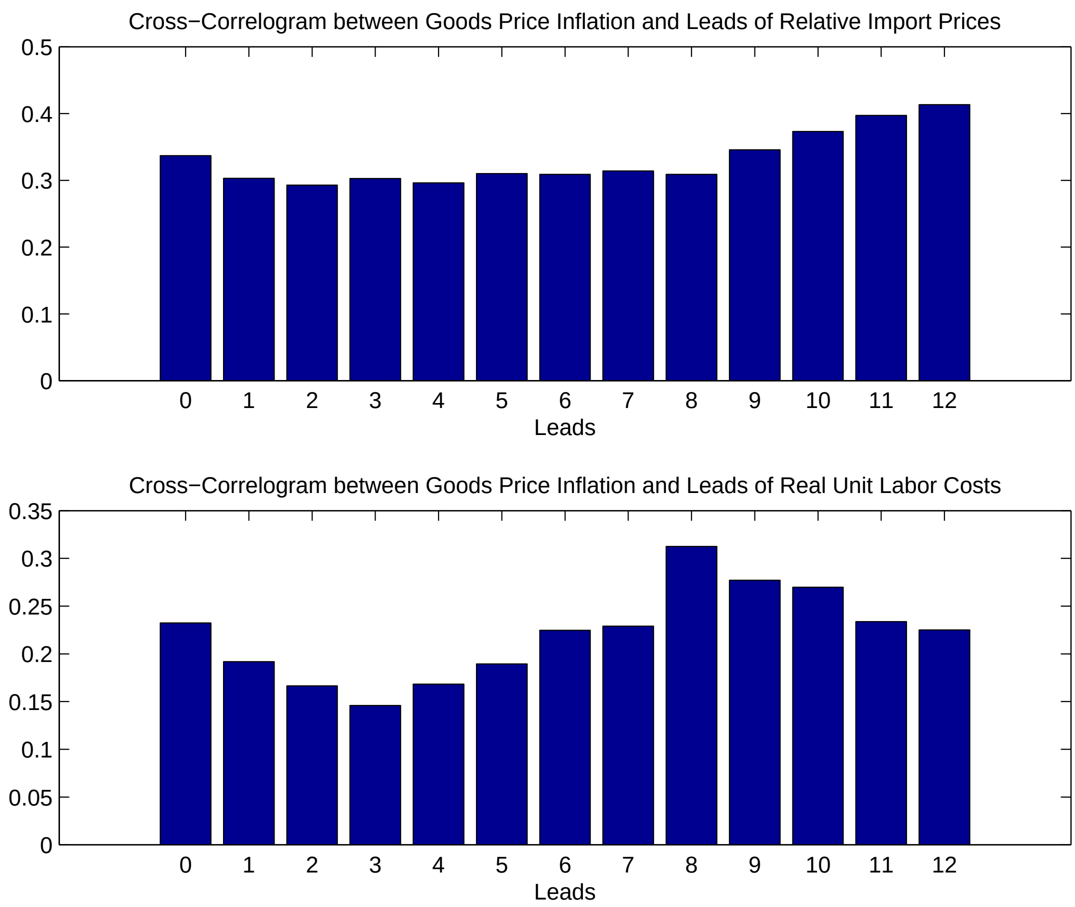


Figure 4: Properties of Estimated Demand Curve

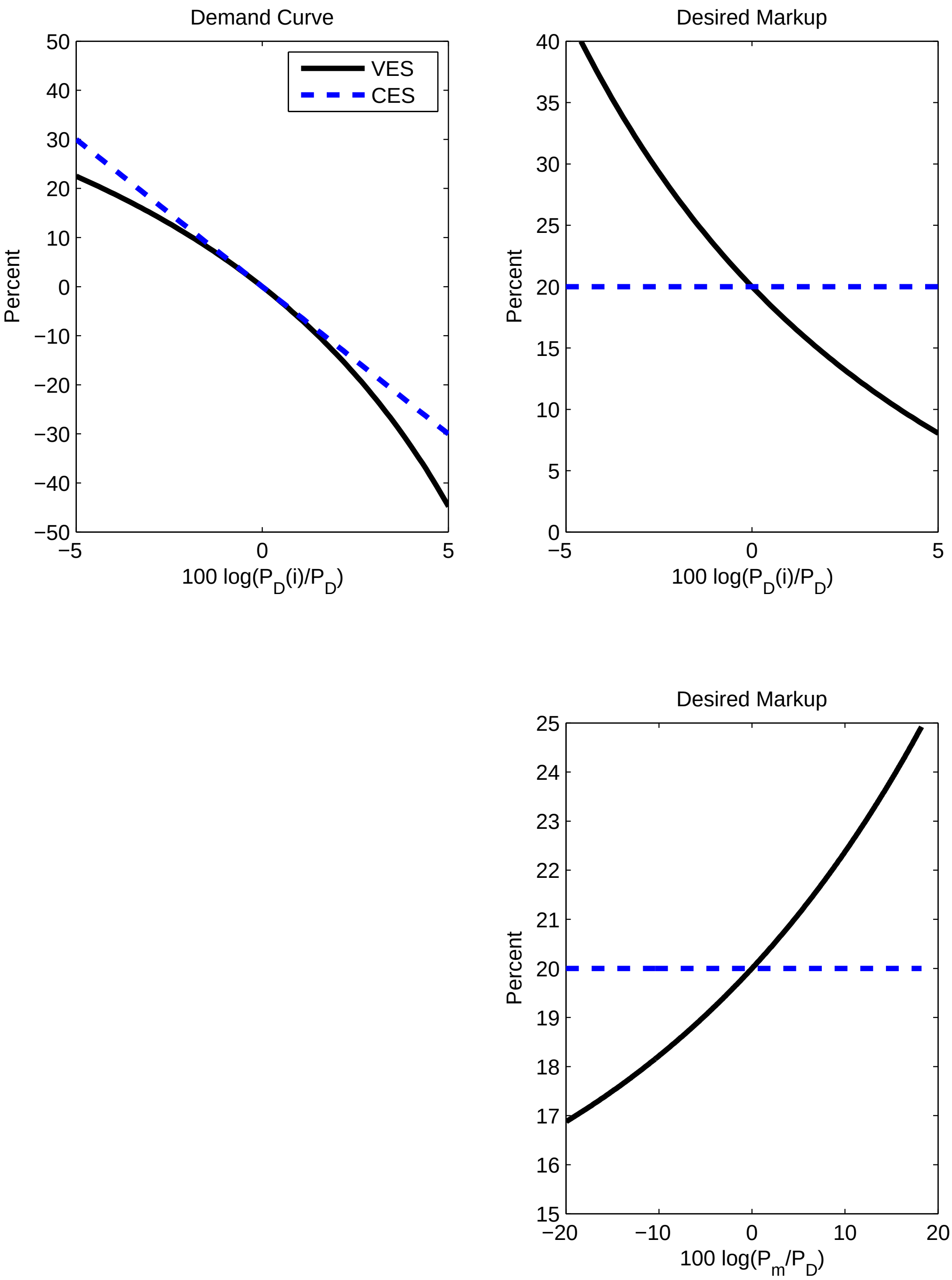
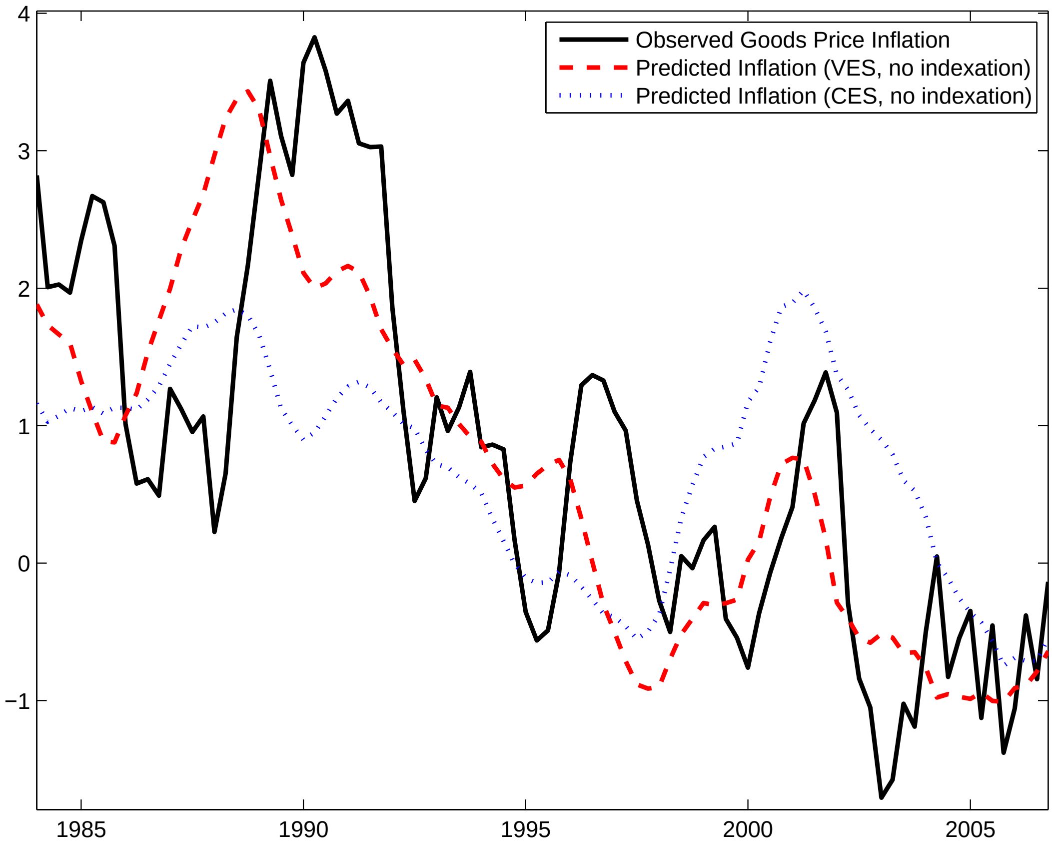
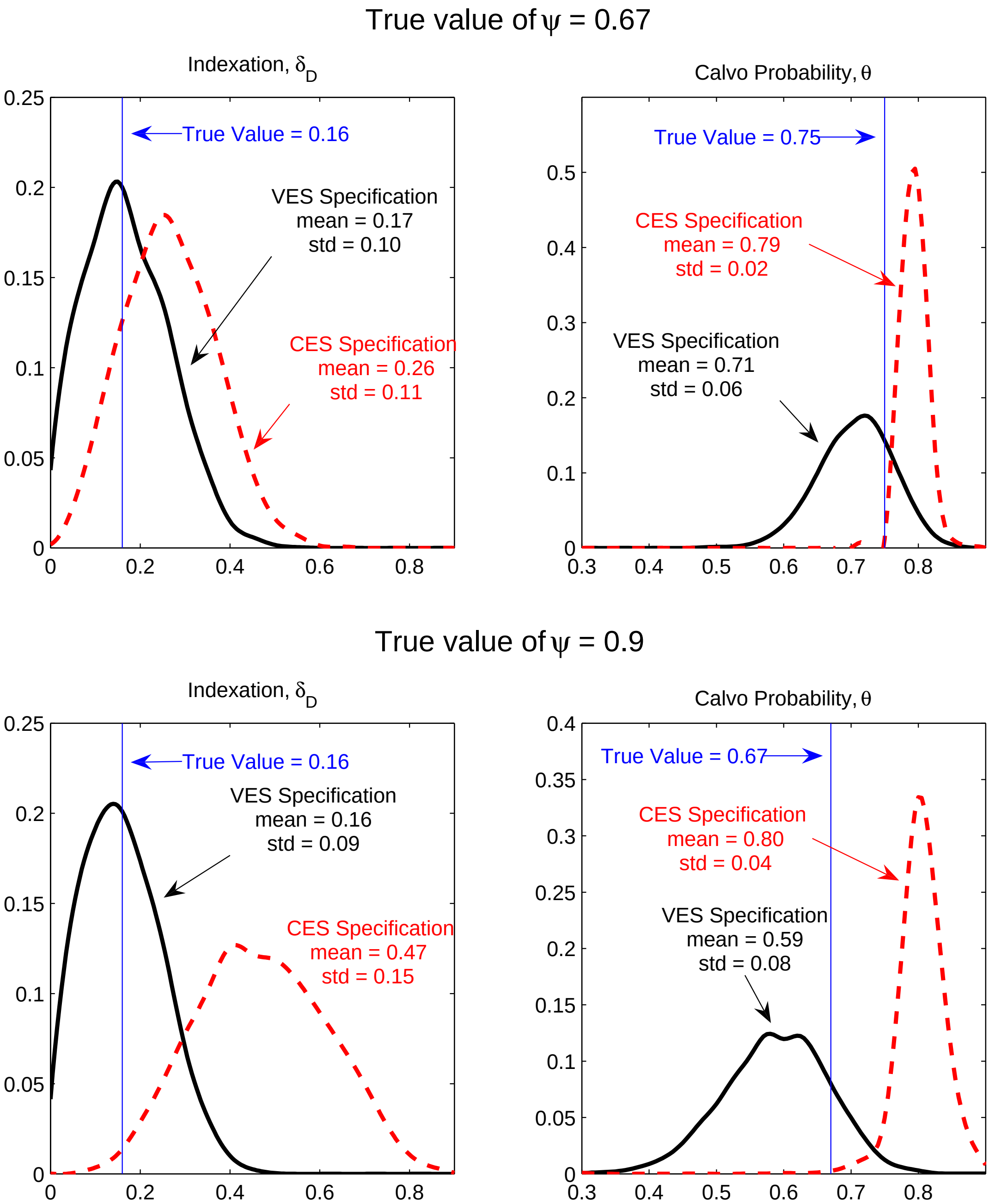


Figure 5: Actual and Predicted Inflation from Alternative Specifications (4-Quarter Moving Average)



Predicted inflation is defined in equation (27) in the text. The estimated parameters used in constructing the predicted series for the VES specification are reported in the second column of Table 2, labelled "VES without indexation." The parameters used for the CES case appear in the fourth column of Table 2, labelled "CES without indexation."

Figure 6: Sampling Distribution of Estimates from Alternative Specifications



The VES specification described in equation (26) in the text is used as the data-generating process.

Appendix

This appendix is divided into three sections. In Appendix A, we derive the demand curves of the final goods producer as well as the log-linearized expression for inflation, i.e., equation (19). Appendix B compares the small sample properties of our approach with estimating the moment condition implied by equation (23). In Appendix C, we show the results from the system estimation of the forecasting VAR.

A Theoretical Derivations

A.1 Deriving the Demand of a Domestically-Produced Good

To derive the demand curves for domestically-produced goods, recall that the representative final goods producer maximizes equation (1) subject to the demand aggregator implied by equations (2)-(4). The first order conditions associated with this problem are:

$$P_{Dt}(i) = \frac{\Lambda_t}{A_t}\left[\frac{1-\nu}{1-\omega}\frac{A_{Dt}(i)}{A_t}+\nu\right]^{\gamma_t-1}\left[V_{Dt}^{1/\rho}+V_{Mt}^{1/\rho}\right]^{\rho-1}V_{Dt}^{\frac{1}{\rho}-1}(1-\omega)^{\rho-1}, \tag{A.1}$$

$$P_{Mt}(i) = \frac{\Lambda_t}{A_t}\left[\frac{1-\nu}{\omega}\frac{A_{Mt}(i)}{A_t}+\nu\right]^{\gamma_t-1}\left[V_{Dt}^{1/\rho}+V_{Mt}^{1/\rho}\right]^{\rho-1}V_{Mt}^{\frac{1}{\rho}-1}\omega^{\rho-1}, \tag{A.2}$$

where Λ_t is the Lagrange multiplier associated with equation (2). Before deriving the demand curves, we need to define $P_{Ft} = \frac{\Lambda_t}{A_t}$ and show that P_{Ft} satisfies equation (8).

To do so, rewrite equations (A.1)-(A.2) as:

$$\left[\frac{1-\nu}{\omega}\frac{A_{Dt}(i)}{A_t}+\nu\right] = \left(\frac{P_{Dt}(i)}{P_{Ft}}\right)^{\frac{1}{\gamma_t-1}}\left[V_{Dt}^{1/\rho}+V_{Mt}^{1/\rho}\right]^{\frac{1-\rho}{\gamma_t-1}}V_{Dt}^{\frac{\rho-1}{\rho(\gamma_t-1)}}(1-\omega)^{\frac{1-\rho}{\gamma_t-1}},$$

$$\left[\frac{1-\nu}{\omega}\frac{A_{Mt}(i)}{A_t}+\nu\right] = \left(\frac{P_{Mt}(i)}{P_{Ft}}\right)^{\frac{1}{\gamma_t-1}}\left[V_{Dt}^{1/\rho}+V_{Mt}^{1/\rho}\right]^{\frac{1-\rho}{\gamma_t-1}}V_{Mt}^{\frac{\rho-1}{\rho(\gamma_t-1)}}\omega^{\frac{1-\rho}{\gamma_t-1}}.$$

Substituting these expressions into equations (3)-(4), we can express V_{Dt} and V_{Mt} as:

$$V_{Dt} = \frac{1}{(1-\nu)\gamma_t}\left(\frac{P_{Dt}}{P_{Ft}}\right)^{\frac{\gamma_t}{\gamma_t-1}}\left[V_{Dt}^{1/\rho}+V_{Mt}^{1/\rho}\right]^{\frac{\gamma_t(1-\rho)}{\gamma_t-1}}V_{Dt}^{\frac{\gamma_t(\rho-1)}{(\gamma_t-1)\rho}}(1-\omega)^{\frac{\gamma_t-\rho}{\gamma_t-1}} \tag{A.3}$$

$$V_{Mt} = \frac{1}{(1-\nu)\gamma_t}\left(\frac{P_{Mt}}{P_{Ft}}\right)^{\frac{\gamma_t}{\gamma_t-1}}\left[V_{Dt}^{1/\rho}+V_{Mt}^{1/\rho}\right]^{\frac{\gamma_t(1-\rho)}{\gamma_t-1}}V_{Mt}^{\frac{\gamma_t(\rho-1)}{(\gamma_t-1)\rho}}\omega^{\frac{\gamma_t-\rho}{\gamma_t-1}}, \tag{A.4}$$

where the price indices, P_{Dt} and P_{Mt}, are defined in equation (7). Using equations (A.3) and (A.4), the ratio of V_{Dt} to V_{Mt} is given by:

$$\left(\frac{V_{Dt}}{V_{Mt}}\right)^{\frac{1}{\rho}} = \left(\frac{P_{Dt}}{P_{Mt}}\right)^{\frac{\gamma_t}{\gamma_t-\rho}}\frac{(1-\omega)}{\omega}. \tag{A.5}$$

Since optimal behavior by a final goods producer implies that equation (2) holds with equality, we can rewrite it as:

$$\left[\left(\frac{V_{Dt}}{V_{Mt}}\right)^{\frac{1}{\rho}}+1\right]^{\rho} V_{Mt}=\frac{1}{(1-\nu)\gamma_t}. \tag{A.6}$$

It is useful to express equation (A.4) as:

$$V_{Mt}=\frac{1}{(1-\nu)\gamma_t}\left(\frac{P_{Mt}}{P_{Ft}}\right)^{\frac{\gamma_t}{\gamma_t-1}}\left[\left(\frac{V_{Dt}}{V_{Mt}}\right)^{\frac{1}{\rho}}+1\right]^{\frac{\gamma_t(1-\rho)}{\gamma_t-1}}\omega^{\frac{\gamma_t-\rho}{\gamma_t-1}}.$$

Substituting this expression and equation (A.5) into equation (A.6), we have:

$$\left[\left(\frac{P_{Dt}}{P_{Mt}}\right)^{\frac{\gamma_t}{\gamma_t-\rho}}\frac{(1-\omega)}{\omega}+1\right]^{\frac{\gamma_t-\rho}{\gamma_t-1}}\omega^{\frac{\gamma_t-\rho}{\gamma_t-1}}P_{Mt}^{\frac{\gamma_t}{\gamma_t-1}}=P_{Ft}^{\frac{\gamma_t}{\gamma_t-1}}.$$

This expression, with some manipulation, can be written as:

$$P_{Ft}=\left[(1-\omega)P_{Dt}^{\frac{\gamma_t}{\gamma_t-\rho}}+\omega P_{Mt}^{\frac{\gamma_t}{\gamma_t-\rho}}\right]^{\frac{\gamma_t-\rho}{\gamma_t}},$$

which is equation (8).

With P_{Ft} defined in this way, we can now turn to deriving the demand curve for a domestically-produced good, i.e., equation (6). We begin by re-expressing equation (A.1) as:

$$\left[\frac{1-\nu}{1-\omega}\frac{A_{Dt}(i)}{A_t}+\nu\right]=\left(\frac{P_{Dt}(i)}{P_{Ft}}\right)^{\frac{1}{\gamma_t-1}}\left[1+\left(\frac{V_{Mt}}{V_{Dt}}\right)^{\frac{1}{\rho}}\right]^{\frac{1-\rho}{\gamma_t-1}}(1-\omega)^{\frac{1-\rho}{\gamma_t-1}}. \tag{A.7}$$

Note that equation (A.5) implies:

$$1+\left(\frac{V_{Mt}}{V_{Dt}}\right)^{\frac{1}{\rho}}=\frac{P_{Dt}^{\frac{\gamma_t}{\rho-\gamma_t}}}{1-\omega}\left[(1-\omega)P_{Dt}^{\frac{\gamma_t}{\gamma_t-\rho}}+\omega P_{Mt}^{\frac{\gamma_t}{\gamma_t-\rho}}\right],$$

or

$$1+\left(\frac{V_{Mt}}{V_{Dt}}\right)^{\frac{1}{\rho}}=\frac{1}{1-\omega}\left(\frac{P_{Ft}}{P_{Dt}}\right)^{\frac{\gamma_t}{\gamma_t-\rho}}. \tag{A.8}$$

Substituting equation (A.8) into equation (A.7) yields:

$$\left[\frac{1-\nu}{1-\omega}\frac{A_{Dt}(i)}{A_t}+\nu\right]=\left(\frac{P_{Dt}(i)}{P_{Ft}}\right)^{\frac{1}{\gamma_t-1}}\left(\frac{P_{Dt}}{P_{Ft}}\right)^{\frac{\gamma_t}{\gamma_t-\rho}\frac{1-\rho}{\gamma_t-1}}.$$

Rearranging this expression, we get equation (6):

$$A_{Dt}(i)=(1-\omega)\left[\frac{1}{1-\nu}\left(\frac{P_{Dt}(i)}{P_{Dt}}\right)^{\frac{1}{\gamma_t-1}}\left(\frac{P_{Dt}}{P_{Ft}}\right)^{\frac{\rho}{\gamma_t-\rho}}-\frac{\nu}{1-\nu}\right]A_t.$$

A.2 Deriving the Log-Linearized Pricing Equation

To derive equation (19), we begin by defining the contract price, $P^c_{Dt}(i) = \frac{P_{Dt}(i)}{P_{Dt}}$, for a firm that optimally chooses its price at date t. Using this definition in equation (10) and log-linearizing, we get:

$$\hat{P}^c_{Dt}(i) = \sum_{j=1}^{\infty}(\beta\theta)^j(\hat{\pi}_{Dt+j} - \delta_D\hat{\pi}_{Dt+j-1}) + (1-\beta\theta)\sum_{j=0}^{\infty}(\beta\theta)^j\left[\hat{s}_{t+j} - \frac{1}{\epsilon-1}\hat{\epsilon}_{t+j}(i)\right]. \quad \text{(A.9)}$$

In the above equation, $\hat{\epsilon}_t(i)$ is the log-linearized version of the elasticity of demand for good i given by:

$$\hat{\epsilon}_{t+j}(i) = \nu\epsilon\left(\hat{P}^c_{Dt}(i) - \sum_{k=1}^{j}(\hat{\pi}_{Dt+k} - \delta_D\hat{\pi}_{Dt+k-1})\right) - \nu\epsilon_A\hat{p}_{Ft+j} + \frac{\gamma}{1-\gamma}\hat{\gamma}_{t+j}, \quad \text{(A.10)}$$

where $\hat{p}_{Ft}$ is the log-linearized price index consisting of all of the prices of a firm's competitors relative to the domestic price index, (i.e., $p_{Ft} = \frac{P_{Ft}}{P_{Dt}}$). Substituting this expression for the elasticity of demand into equation (A.9), we have:

$$\hat{P}^c_{Dt}(i) = \sum_{j=1}^{\infty}(\beta\theta)^j(\hat{\pi}_{Dt+j} - \delta_D\hat{\pi}_{Dt+j-1}) + \frac{1-\beta\theta}{1+\frac{\nu\epsilon}{\epsilon-1}}\left[\hat{s}_{t+j} + \frac{\nu\epsilon}{\epsilon-1}\frac{\epsilon_A}{\epsilon}\hat{p}_{Ft+j} - \left(\frac{1}{\epsilon-1}\right)\left(\frac{\gamma}{1-\gamma}\right)\hat{\gamma}_{t+j}\right].$$

Using the definition of the steady state markup (i.e., $\mu = \frac{\epsilon}{\epsilon-1}$) and the definition of Ψ (i.e., $\Psi = \frac{\nu\mu}{1+\nu\mu}$), this expression, after quasi-differencing, can be rewritten as:

$$\hat{P}^c_{Dt}(i) - \beta\theta\hat{P}^c_{Dt+1}(i) = \beta\theta(\hat{\pi}_{Dt+1} - \delta_D\hat{\pi}_{Dt}) + (1-\beta\theta)\left[(1-\Psi)\hat{s}_t + \Psi\frac{\epsilon_A}{\epsilon}\hat{p}_{Ft} + (2\Psi-1)\hat{\gamma}_t\right]. \quad \text{(A.11)}$$

From the log-linearized version of the first expression in equation (7), the contract price at date t can be related to traded goods inflation via:

$$\hat{P}^c_{Dt}(i) = \frac{\theta}{1-\theta}(\hat{\pi}_{Dt} - \delta_D\hat{\pi}_{Dt-1}). \quad \text{(A.12)}$$

Substituting this expression into equation (A.11), we get an expression relating domestic price inflation to real marginal cost and p_{Ft}:

$$\hat{\pi}_{Dt} - \delta_D\hat{\pi}_{Dt-1} = \beta(\hat{\pi}_{Dt+1} - \delta_D\hat{\pi}_{Dt}) + \kappa\left[(1-\Psi)\hat{s}_t + \Psi\frac{\epsilon_A}{\epsilon}\hat{p}_{Ft} + (2\Psi-1)\hat{\gamma}_t\right]. \quad \text{(A.13)}$$

The log-linearized version of equation (8) implies that

$$\hat{p}_{Ft} = \omega\hat{p}_{Mt}.$$

Using this expression in equation (A.13) yields equation (19).

B Small Sample Properties of our GMM Estimator

This appendix investigates the small-sample properties of our estimator through a Monte Carlo exercise. We bootstrapped 10,000 Monte Carlo samples for traded goods inflation, relative import prices, and the labor share using equation (26), and its auxiliary forecasting processes, which we have reproduced below:

$$\hat{\pi}_t = \delta_D \hat{\pi}_{t-1} + \kappa_D \left[\frac{1-\Psi}{1-\beta\rho_s} \hat{s}_t + \omega \frac{\epsilon_A}{\epsilon} \frac{\Psi(1+\beta\rho_{M2}L)}{1-\beta\rho_{M1}-\beta^2\rho_{M2}} \hat{p}_{Mt} \right] + \epsilon_{\pi t} \tag{B.1}$$
$$\hat{s}_t = \rho_s \hat{s}_{t-1} + u_{st}$$
$$\hat{p}_{Mt} = \rho_{M1}\hat{p}_{Mt-1} + \rho_{M2}\hat{p}_{Mt-2} + u_{p_M t}.$$

As pseudo-true parameter values for our Monte Carlo exercise, we used the GMM estimates of θ, Ψ, δ_D, ρ_s, ρ_{M1}, and ρ_{M2} from equations (B.1) with the following instrument set: one lag of the labor share, two lags of goods inflation, and two lags of relative import prices as instruments (see the second column of Table 5). In bootstrapping, we jointly sampled with replacement from this model's estimates of $\epsilon_{\pi t}$ and the residuals from the two autoregressive processes, which maintains the correlation structure across residuals. Each bootstrapped sample had a length of 96 observations, which is the same length as the 1983Q1-2006Q4 sample.

For each sample of artificial data, we used the generalized method of moments to re-estimate θ, Ψ, and δ_D based, again, on equations (B.1) using the same instrument set as above. The solid lines in Figure 7 show the parameters' sampling distributions. As discussed earlier in the text, the estimates of θ and δ_D based on the closed form solution do not display much small-sample bias. The estimate of Ψ does display some modest upward bias, and the small-sample confidence intervals appear a little wider than intervals based on asymptotic derivations. Overall, however, our GMM estimator performs well in small samples.

Figure 7 also compares the performance of our baseline estimator with an alternative GMM estimator that uses equation (23) as the moment condition (reproduced below):

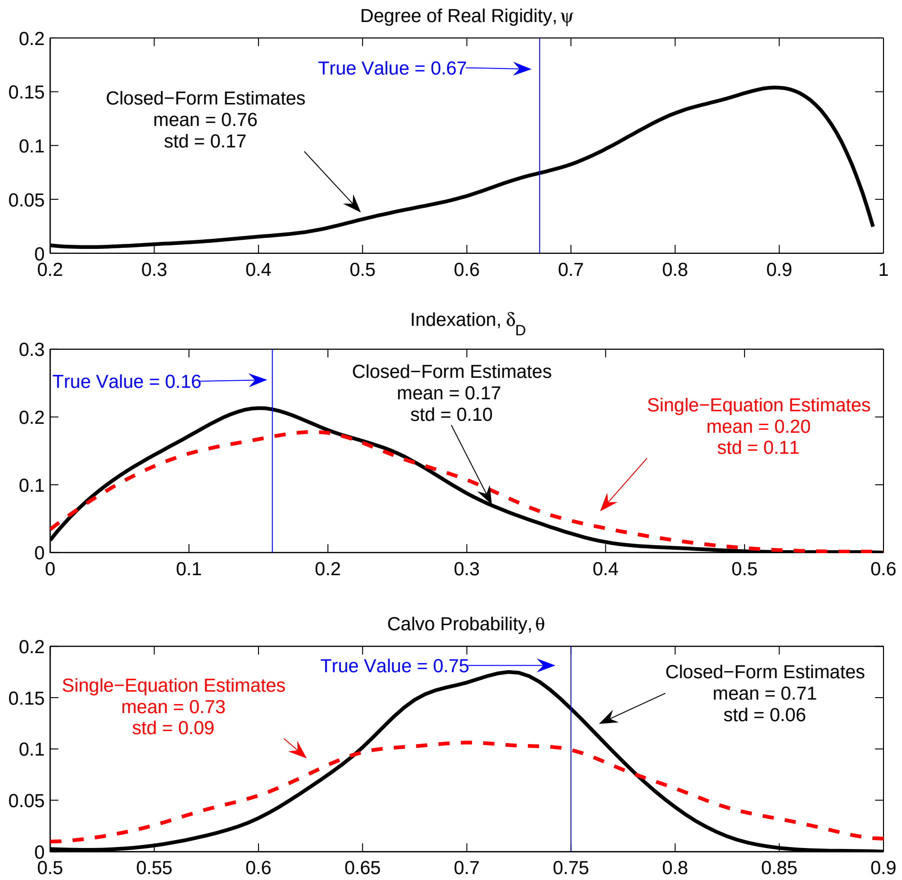
$$\hat{\pi}_t - \delta_D\hat{\pi}_{t-1} = \beta E_t[\hat{\pi}_{t+1} - \delta_D\hat{\pi}_t] + \kappa_D \left[(1-\Psi)\hat{s}_t + \Psi\omega\frac{\epsilon_A}{\epsilon}\hat{p}_{Mt} + \varphi\widehat{\gamma_t} \right] \tag{B.2}$$
$$\kappa_D = \frac{(1-\beta\theta)(1-\theta)}{\theta[1+\epsilon\frac{\alpha}{1-\alpha}(1-\Psi)]}.$$

For convenience, we shall refer to our baseline estimator as "closed-form", while we dub the alternative, based on equation (B.2), "single-equation."

In contrast to our closed-form estimator, we find that single-equation GMM estimator performs poorly in small samples. In particular, we found that estimates of Ψ were severely upward-biased with the mass of the distribution near unity – the upper bound of feasible values for Ψ. Accordingly, we do not show sampling distributions when we jointly estimate Ψ with θ and δ_D. Instead, the dashed lines labelled "Single-Equation Estimates" in the middle and lower panels of Figure 7 show the sampling distributions for θ and δ_D, conditional on ψ being at its pseudo-true value of 0.67. We find it remarkable that, despite using the pseudo-true value of

Ψ, the GMM estimator based on equation (B.2) still performs worse in small samples than the GMM estimator based on the closed form solution. The former yields estimates of δ_d that are more upward-biased and the distribution of estimates for θ has fatter tails.

Figure 7: Comparison of Small Sample Properties of Alternative Estimators



C Empirical Estimates of the Forecasting VAR

Before we show the estimates of the forecasting VAR, recall that the VAR that we use can be rewritten in companion form as:

$$X_t = AX_{t-1} + u_t,$$

where $X_t' = [\hat{s}_t \; \hat{p}_{Mt} \; \hat{s}_{t-1} \; \hat{p}_{Mt-1}]$, and $u_t' = [u_{st} \; u_{p_M t} \; 0 \; 0]'$. Accordingly, the first row of the matrix A corresponds to the estimated process for the labor share ($\hat{s}_t$) and the second row corresponds to the estimated process for relative import prices ($\hat{p}_{Mt}$).

Table 6 shows the estimates of A for three different specifications: the benchmark VES, CES with lagged indexation, and the VES specification using an unrestricted VAR(2). (We show only the results in which capital is firm-specific, since the estimates of A are unchanged under the alternative assumption of full capital mobility within a country.) The table confirms that both the labor share and relative import prices are well-approximated by simple, univariate processes. In particular, for the benchmark VES specification, there is no evidence of feedback between the two variables (i.e., $A_{12} = A_{14} = A_{21} = A_{23} = 0$).

Table 6: System Estimates of VAR

	Benchmark VES	CES with indexation	VAR(2) Forecasting Model
Labor Share Equation:			
A_{11}	0.89	0.91	0.71
	(0.03)	(0.03)	(0.14)
A_{12}	0	0	0.00
	–	–	(0.05)
A_{13}	0	0	0.21
	–	–	(0.16)
A_{14}	0	0	0.00
	–	–	(0.05)
R^2	0.83	0.83	0.83
Durbin-Watson Statistic	2.26	2.32	1.78
Relative Import Price Equation:			
A_{21}	0	NA	-0.01
	–	–	(0.11)
A_{22}	1.37	NA	1.39
	(0.06)	–	(0.06)
A_{23}	0	NA	0.00
	–	–	(0.1)
A_{24}	-0.39	NA	-0.41
	(0.06)	–	(0.06)
R^2	0.98	NA	0.98
Durbin-Watson Statistic	1.91	NA	1.92

[a]Standard errors are reported in parentheses.