Inflation Asymmetry and Menu Costs New Micro Data Evidence*

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Abstract

We use the natural experiment provided by a series of value-added tax (VAT) changes in Hungary to provide micro-data evidence on the asymmetry of inflation response to aggregate shocks. We show that even though a standard menu cost model like that of Golosov and Lucas (2007) underestimates the asymmetry, a sectoral menu cost model with multi-product firms and trend-inflation can quantitatively account for the inflation asymmetry observed in the data, thereby it provides direct evidence to the argument of Ball and Mankiw (1994). The model predicts that the effect of a positive monetary policy shock can have almost twice as large inflation effect as a negative shock.

Keywords: Menu Cost, Inflation Asymmetry, Sectoral Heterogeneity, Value-Added Tax

JEL Classification: E30

1 Introduction

The paper sets out to examine asymmetric inflation response to symmetric aggregate shocks on a new comprehensive CPI data set of Hungary, capitalizing on the natural experiment provided by a major value added tax (VAT) increase and a major tax decrease in 2006. These VAT shocks provide exceptional information about the pricing behavior of firms[1], as these exogenous cost push shocks influence a large number of firms simultaneously in an easily measurable way. The main question of the paper is what this episode would imply for the asymmetry of monetary policy shocks.

Gabriel and Reiff, 2007 have documented that the Hungarian VAT shocks had an asymmetric aggregate inflation effect (see Figure 1): while the 2006 September 5%-points increase

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[1]Especially in Hungary, where firms are required by law to quote gross prices.

of the 15% VAT rate[2] increased the sample CPI by 2.13%, the 2006 January 5%-points decrease of the 25% VAT rate[3] decreased the CPI only by 0.92%. The asymmetry is even more pronounced in the subsamples of products directly affected by the tax changes: their average price increase was 3.73% for the VAT increase and only -1.24% for the VAT decrease.

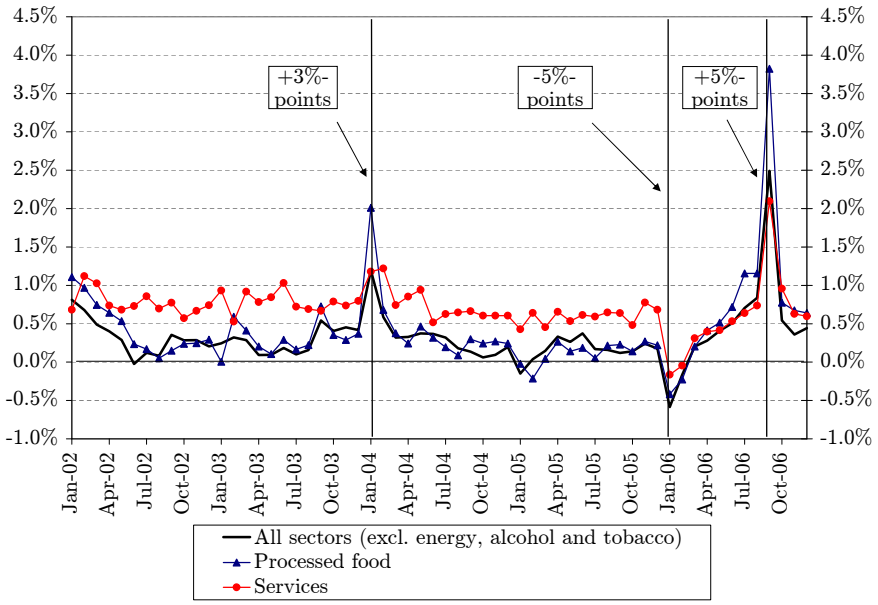

Figure 1: Monthly inflation rates and VAT rate changes (sample averages)

As standard frictionless, flexible price models can not explain the observed asymmetry, the paper departs from them by assuming sticky prices. As it was argued by Ball and Mankiw, 1994, sticky prices with positive trend inflation imply asymmetric inflation response, because forward looking firms setting their prices for several periods will incorporate the effects of the inflation; and they are going to be more responsive to a positive shock than to negative one. The conjecture that this channel is the main reason for the asymmetry is further reinforced by the observation that, perfectly in line with the prediction of the model, the asymmetry is much higher in the services sector with high inflation rate and high price stickiness than in the - similarly large - processed food sector which has lower inflation rate and lower level of price stickiness.[4]

The menu cost model – which assumes that firms face a small fixed cost for changing their prices, but choose endogenously when to change them – is well suited to explain the major

[2]Influencing 46.9% of the products in our sample.

[3]Influencing another 51%.

[4]A series of recent papers, use search frictions instead to explain asymmetric price responses (Cabral-Fishman, 2006, Yang-Ye, 2007). Their common key assumption is that the marginal consumers are uninformed about the cost shocks faced by price setting firms. This assumption, however, is not applicable to our case of VAT changes, because these had easily measurable, widely publicized and uniform effects on all affected firms.

pricing effects of the observed tax shocks. Flexible price models would have serious difficulty in explaining the fact that at the months of tax changes only around 60% of the affected firms changed their prices. A model with Calvo pricing assuming exogenous probability of price change, on the other hand, would be unable to account for the fact that the frequency of price changing firms at the months of VAT shocks increased to 31.4% from the average 12.1% during normal times (see Figure 3 later).

The standard menu cost model with idiosyncratic technology shock as suggested by Golosov and Lucas (2007) is able to hit both the average frequency and the magnitude of the price changes in 'normal times' by reasonable parameter values. The model, however, can not quantitatively explain the asymmetric inflation effects of the tax shock for two main reasons.

First the standard model underestimates the fraction of price changing firms, and can not account for the observed *decreased* average size of the price changes at the months of price changes (see Figures 3 and 4 later). Standard channels in menu costs models can give some explanation for the sign of these effects, especially if we consider that VAT shocks are more persistent than other cost shocks justifying a lower threshold for price changes. Their magnitude, however, as we show in the paper, can only explained by a lower average menu cost at the months of VAT changes, which can reproduce both the frequency and the size of price changes observed in the data. The lower average menu cost can be justified by the existence of multi-product firms with decreasing marginal menu costs for changing the second-third-etc. prices as in Midrigan (2006).

Second, the aggregate model calibrated to hit the average inflation rate, the average frequency and size of price changes still underestimates the observed inflation asymmetry. The main reasons for this is related to the sectoral heterogeneity in menu costs and inflation rates. In our model, the interaction of trend inflation and menu costs imply a convex effect on the asymmetry: the asymmetry implied by an average inflation rate and menu cost is lower than the average asymmetry implied by the sectors with different menu costs and inflation. As the sectoral inflation rate and price stickiness is not independent in our sample because of the relative importance of sectors with both high inflation rate and substantial price stickiness, the sectoral heterogeneity can substantially bias the estimates of the aggregate model. To take this also in consideration, the paper calibrates a sectoral menu cost model. This choice, furthermore, allows us to control for the different sectoral composition of the observed tax changes. The VAT changes affected various sectors differently, the VAT increase hitting sectors with more flexible prices disproportionally, providing partial explanation for the observed asymmetry.

The paper calibrates a sectoral menu cost model with aggregate inflation uncertainty using the methodology of Krusell-Smith, 1998. The main finding of the paper is that the model calibrated to hit the mean sectoral inflation rate and the level of price stickiness observable in the data can successfully account for the sectoral asymmetry observed in the data. The calibrated sectoral model, then, is used to simulate the asymmetric effects of large monetary policy shocks. The paper finds, that this shock - hitting all sectors symmetrically - would have a somewhat smaller, but economically very significant asymmetric effects on the inflation rate: the inflation effect of a positive shock can be almost twice as large as the effect of a negative one.

There is a long line of research documenting asymmetric price developments to monetary and cost shocks using aggregate (see e.g. Cover, 1992, Ravn and Sola, 2004) and sectoral data

(Peltzman, 2000) in reduced form estimations. Our paper is the first we know of, however, which uses VAT shocks to analyze the effects of the asymmetry, which is arguably a more easily measurable and identifiable shock than those used by the previous papers. As standard frictionless, flexible price models would have difficulty in explaining these asymmetries, our paper is also a contribution to the growing literature using natural experiments – like the effect of euro introduction to restaurant prices in Hobijn, Ravenna and Tambalotti, 2006 – and special environments – like the high inflation episodes in Mexico in Gagnon, 2007 – to provide evidence to the validity of the sticky price assumptions in general and menu cost models in particular.

The paper is organized as follows. Section 2 presents the model, solves for the flexible price case and explains the numerical algorithm used for the solution in case of positive menu costs. Section 3 presents the data and the moments the paper is about to match, and presents some stylized facts the data suggests. Section 4 presents the results for the non-sectoral and the sectoral calibrations and section 5 concludes.

2 The Model

The paper uses a standard quantitative menu cost model able to reproduce important moments of the price developments including the frequency of price changing firms, the average size of price changes and the inflation rate. To reproduce the observed large size of price changes, we follow the now standard assumption of Golosov and Lucas, 2007 (see Gertler-Leahy, 2007, Midrigan, 2006, Nakamura-Steinsson, 2007) that firms are hit by large idiosyncratic shocks, and we model equilibrium trend inflation by assuming (exogenously given) nominal aggregate output (money supply) growth larger than the mean aggregate technology growth.

We introduce value added taxes (VAT) to the framework in a straightforward way; as we do not model explicitly the production process, VAT in our framework is equivalent to a standard sales tax. We assume - in line with the Hungarian legal rules - that firms set gross prices and they need to pay a fixed menu cost in case they decide to change these.

We are going to deviate from the standard model in two key respects to be able to reproduce the observed response to the tax shocks. The first is that we allow for menu costs being lower at the month of the tax change as a shortcut for multiproduct firms with decreasing marginal costs of price changes (Midrigan, 2006). The second is that we introduce sectoral heterogeneity into the model. As our main reason is to take differences in sectoral inflation rates and price stickiness into consideration, we assume no sectoral interactions similarly to Klenow and Willis, 2006. We do this by assuming Cobb-Douglas preferences for sectoral product baskets resulting in fixed sectoral nominal expenditure shares, and assuming sector specific labor markets.

The model assumes that the firms determine correct linear beliefs about the development of the endogenous sectoral state variables as is assumed in Krusell and Smith, 1998, and it is going to be solved numerically by value function iteration over a discretized state space.

2.1 The consumer

The representative consumer is assumed to consume a Dixit-Stiglitz aggregate (C) of individual goods i, hold real balances M/P and supply sector specific labor L^s to maximize the expected present value of his utility (The subscript indices denote time (t), superscripts the sector (s) and individual firms and products are in brackets (i))

$$\max_{\{C_t^s(i), L_t^s, M_t\}} E \sum_{t=0}^{\infty} \beta^t \left(\log \left[C_t \cdot \left(\frac{M_t}{P_t} \right)^{\nu} \right] - \sum_{s=1}^{S} \frac{\mu^s}{1+\psi^s} \left(L_t^s \right)^{1+\psi^s} \right), \tag{1}$$

where the aggregate C_t and sectoral consumptions C_t^s are determined by constant elasticity of substitution aggregators

$$C_t = \prod_{s=1}^{S} \left(\frac{C_t^s}{\alpha^s} \right)^{\alpha^s}, \quad C_t^s = \left(\sum_{i=1}^{n^s} (n^s)^{-\frac{1}{\theta}} C_t^s(i)^{\frac{\theta-1}{\theta}} \right)^{\frac{\theta}{\theta-1}},$$

with elasticity of substitution θ. Note that with $S = 1$ and $\alpha^1 = 1$ the model reduces to a standard non-sectoral model. We assume that n^s is large, implying that a consumption good i constitutes only an infinitesimal part of the consumer's utility.

The consumer's budget constraints for all time period t and history h^t (dependence on history is suppressed for notational convenience) are given by

$$\sum_{s=1}^{S} \sum_{i=1}^{n^s} P_t^s(i) C_t^s(i) + \sum_{h_{t+1}} B_{t+1}(h_{t+1}) + M_{t+1} = R_t B_t + M_t + \sum_{s=1}^{S} \tilde{w}_t^s L^s + \tilde{\Pi}_t + T_t, \tag{2}$$

where $P_t^s(i)$ is the gross price, $B_t(h_t)$ is a nominal Arrow-security with state dependent gross return R_t, M_t is the nominal money balance and T_t is a lump-sum transfer.

In order to express the consumer's optimality conditions in a convenient form, it is useful to define the aggregate (P_t) and the sectoral price levels (P_t^s) by

$$P_t = \prod_{s=1}^{S} (P_t^s)^{\alpha^s}, \quad P_t^s = \left(\sum_{i=1}^{n_s} \frac{P_t^s(i)^{1-\theta}}{n_s} \right)^{\frac{1}{1-\theta}},$$

which are standard definitions, implying that the aggregate and sectoral expenditures are given by $P_t C_t$ and $P_t^s C_t^s$ respectively.

The Cobb-Douglas formulation across sectors implies that the consumer will spend a constant α_s fraction of his nominal expenditures on the sectoral composite good. His real demand can be expressed as

$$C_t^s = \alpha^s \left(\frac{P_t^s}{P_t} \right)^{-1} C_t, \tag{3}$$

and his demand for individual good i from sector s, in turn, is given by

$$C_t^s(i) = \frac{1}{n^s} \left(\frac{P_t^s(i)}{P_t^s} \right)^{-\theta} C_t^s. \tag{4}$$

The Euler equation of the consumer implies that the stochastic discount factor $\frac{1}{R_{t+1}}$ is given by

$$\frac{1}{R_{t+1}} = \beta \frac{P_t C_t}{P_{t+1} C_{t+1}}, \tag{5}$$

the labor supply equation in each sector $s = 1, \ldots, S$ is given by

$$\mu^s \left(L_t^s\right)^{\psi_s} C_t = \frac{\tilde{w}_t}{P_t}, \tag{6}$$

and, finally, the money demand equation is going to be

$$\frac{M_t}{P_t} = \nu C_t \frac{i_t + 1}{i_t}, \tag{7}$$

where i_t is the nominal interest rate.

2.2 The government and the central bank

The value added tax rates ($\tau_t(i)$) are assumed to follow two state Markov processes with high persistence. In line with the products in our sample, we assume that a subset of firms ($i \leq n^{s,1}$) in each sectors face tax rate τ_t^1, while the remaining firms ($i > n^{s,1}$) in the same sector face tax rate τ_t^2. We assume that the revenues are redistributed in a lump sum fashion:[5]

$$\sum_{i=1}^{n} P_t(i) \frac{\tau_t(i)}{1 + \tau_t(i)} C_t(i) = T_t^g.$$

The central bank is assumed to follow a nominal income targeting rule by maintaining a predetermined growth rate (g_{PY}) of the nominal aggregate output ($P_t Y_t$), which we assumed to be equal to the money supply (M_t). The exogenous nominal growth assumption substantially simplifies the analysis, allowing the paper to focus on firm level and sectoral incentives for responding to tax changes. The resulting extra money supply M_t in the economy is redistributed in a lump sum way

$$M_t - M_{t-1} = T_t^m,$$

where $T_t^m + T_t^g = T_t$ is the total transfer to the consumers.

2.3 The firms

The firms are assumed to maximize the present value of their profits. Each firm i is assumed to produce product i monopolistically, post gross prices $P_t(i)$ and satisfy all demand given this price. They face a small menu cost $\tilde{\phi}_t$ if they choose to change their prices; these are assumed to be proportional to their revenues. We also allow the menu cost to be smaller for tax shocks as an implicit way of modeling multi-product firms. In what follows, we characterize the firms' problem in sector s, for notational convenience we suppress the reference to this.

[5] If the net prices P^n are taxed by τ VAT, then the gross price is $P = (1+\tau)P^n$, so the tax revenue equals $\tau P^n C = \frac{\tau}{1+\tau} PC$.

We distinguish two types of firms: there are n^1 number of firms with tax rate τ^1 and $n^2 = n - n^1$ number of firms with tax rate τ^2. By this, we are modeling the fact that firms face different VAT tax levels and the tax changes influenced only a subset of firms in each sector.

The firms' problem is to maximize the expected discounted present value of their profits

$$\max E \sum_{t=0}^{\infty} \frac{1}{\prod_{q=0}^{t} R(q)} \tilde{\Pi}_t(i), \tag{8}$$

where the periodic profit level is given by

$$\tilde{\Pi}_t(i) = \frac{1}{1+\tau_t(i)} P_t(i) Y_t(i) - \tilde{w}_t L_t(i). \tag{9}$$

We assume that the firms use a constant returns to scale technology function with only labor as a factor to produce their differentiated good i and face idiosyncratic $A_t(i)$ and sectoral technology shocks Z_t. The production functions of the firms are given by

$$Y_t(i) = Z_t A_t(i) L_t(i). \tag{10}$$

This production function (10) implies an individual labor demand

$$L_t(i) = \frac{Y_t(i)}{Z_t A_t(i)}, \tag{11}$$

which aggregates to a sectoral labor demand given by

$$L_t = \sum_{i=1}^{n} L_t(i). \tag{12}$$

Substituting the the individual demand (equation (4)) and the labor demand (equation (11)) into the periodic profit function (9) and using the equilibrium condition $Y_t(i) = C_t(i)$, we get that

$$\tilde{\Pi}_t(i) = \frac{1}{1+\tau_t} \frac{1}{n} P_t(i) \left(\frac{P_t(i)}{P_t}\right)^{-\theta} Y_t - \tilde{w}_t \frac{\frac{1}{n}\left(\frac{P_t(i)}{P_t}\right)^{-\theta} Y_t}{Z_t A_t(i)}.$$

As we have constant nominal growth in the model, we are going to normalize the profit level by the average sectoral revenues

$$\Pi_t(i) = \frac{\tilde{\Pi}_t(i) n}{\alpha P_t Y_t},$$

where we used equation (3) implying constant relative sectoral expenditures given by α. Let $p_t(i) = \frac{P_t(i)}{P_t}$ be the (sectoral) relative price, $w_t = \frac{\tilde{w}_t}{P_t Y_t}$ the normalized wage rate, and $\phi_t = \frac{\tilde{\phi}_t n}{\alpha P_t Y_t}$ the normalized menu cost. Let $\zeta(t) = w_t \frac{Y_t}{Z_t}$ be a sectoral cost factor. Substituting these variables into the normalized periodic profit function we get that

$$\Pi_t(p_t(i), A_t(i), \zeta_t, \tau_t(i)) = \frac{1}{1+\tau_t(i)} \left(p_t(i)\right)^{1-\theta} - \left(p_t(i)\right)^{-\theta} \zeta_t A_t(i)^{-1} - \phi_t, \tag{13}$$

where the firm needs to pay ϕ_t only if it chooses to change its price.

The exogenous state variables of the normalized problem are given by $(A_t(i), g_{Zt}, \tau_t^1, \tau_t^2, \phi_t)$ where g_{Zt} is the random growth rate of the sectoral technology process. The endogenous state variables of the problem are given by $(p_{t-1}(i), \pi_t, \zeta_t, \Gamma_t)$, where π_t is the sectoral inflation rate and Γ_t is the distribution of relative prices. To present the firms' Bellman equation, we express the set of state variables as $(p_{-1}(i), \Omega(i))$, where $\Omega(i) = (A(i), \pi, \zeta, g_Z, \tau^1, \tau^2, \phi, \Gamma)$.

Given these state variables the value of firm i is determined by the maximum it can get by changing (C) its nominal price or keeping it constant NC

$$V\left(p_{-1}(i), \Omega(i)\right) = \max_{\{C,NC\}} \left[V^{NC}\left(p_{-1}(i), \Omega(i)\right), V^{C}\left(p_{-1}(i), \Omega(i)\right)\right], \tag{14}$$

where the value function in case of no price change (NC) is given by

$$V^{NC}\left(p_{-1}(i), \Omega(i)\right) = \Pi(i)\left(\frac{p_{-1}(i)}{1+\pi}, A(i), \zeta, \tau(i)\right) + \beta EV\left(\frac{p_{-1}(i)}{1+\pi}, \Omega'(i)\right) \tag{15}$$

where we used the fact that if the firm decides to keep its nominal price constant, its relative price $p(i)$ is going to depreciate by the inflation rate. The value function of firm i in case of price change is given by

$$V^{C}\left(p_{-1}(i), \Omega(i)\right) = \max_{p(i)} \Pi(p(i), A(i), \zeta, \tau(i)) + \beta EV\left(p(i), \Omega'(i)\right). \tag{16}$$

The endogenous distribution of the relative prices Γ is, in general, a very complicated function of the last period price distribution Γ_{t-1}, the current distribution of the sectoral idiosyncratic technology distribution Λ_t and the development of the exogenous state variables $g_{Zt}, \tau_t^1, \tau_t^2, \phi_t$:

$$\Gamma_t = \Theta(\Gamma_t, \Lambda_t, g_{Zt}, \tau_t^1, \tau_t^2, \phi_t) \tag{17}$$

2.4 The equilibrium

We consider a closed economy stochastic dynamic general equilibrium with firms forming linear forecasts about the aggregate endogenous state variables like in Krusell and Smith, 1998. The equilibrium requires

1. The representative consumer chooses $C_t(i), L_t^s, M_t$ to maximize his utility function (1) given his budget constraint (2) taking goods prices $\{P_t(i)\}$, the interest rates R_t and the sectoral wages $\{w^s\}$ as given.

2. The firms are assumed to set prices $P_t(i)$ to maximize their value function (14), (15), (16) given the current exogenous state variables and having correct beliefs about the random process of the idiosyncratic shock, the sectoral technology growth shock (which both assumed to follow independent AR(1) processes with sector specific persistence parameters $\rho_A^s, \rho_{g_Z}^s$ and standard errors $\sigma_A^s, \sigma_{g_Z}^s$), the taxes and the menu costs.

   Following Krusell and Smith, 1998, we assume that the firms – instead of calculating the whole distribution of prices given by equation (17) predict only the current inflation rate – the aggregate moment they are interested in – using a linear equation:

$$\hat{\pi}_t^s = \gamma_1 + \gamma_2 \pi_{t-1}^s + \gamma_3 \zeta_{t-1}^s + \gamma_4 g_{Zt}^s + \gamma_5 g_{\overline{\tau}^s t}^{+} + \gamma_6 g_{\overline{\tau}^s t}^{-} + \epsilon_\pi^s,\ \epsilon_\pi^s \sim N(0, \sigma_{\epsilon_\pi}^s) \tag{18}$$

containing all the current exogenous aggregate state variables and the lagged values of the endogenous state variables. The firms take into account the potential asymmetry in the inflation effect of the average tax rates by incorporating the increase/decrease $g_{\bar{\tau}}^{\pm}$ separately in the equation. The prediction error ϵ_π – incorporating errors resulting from ignoring the whole price distribution – is assumed to be orthogonal to the regressors.

Given this predicted sectoral inflation rate, the forecast for the sectoral output growth $\hat{g}_{Y_t^s}$ is given by

$$\hat{g}_{Y_t^s} = g_{PY} - \hat{\pi}_t^s, \tag{19}$$

and the sectoral cost parameter $\hat{\zeta}_t^s$ is

$$\log \hat{\zeta}_t^s = \log \zeta_{t-1}^s + \hat{g}_{Y_t^s} - g_{Z_t^s} + \hat{g}_{w_t^s}, \quad \hat{g}_{w_t^s} = \psi \left(\hat{g}_{Y_t^s} - g_{Z_t^s} \right). \tag{20}$$

The estimate for the expected wage growth uses the approximate result $\log L_t^s \sim \psi(\log Y_t^s - \log Z_t^s)$ using the individual labor demand (11) and the labor supply (6) equations .

3. The central bank sets i_t, M_t to keep the nominal output growth g_{PY} constant, and the fiscal transfers are set in a way to imply balanced budget.

4. Market clearing in all goods markets $C_t(i) = Y_t(i)$.

5. Assets in zero net supply in nominal Arrow securities: $B_t = 0$.

6. Equilibrium in the sectoral labor markets implying sectoral wages w_t^s equating sectoral labor demand (11) and labor supply (6).

2.5 Flexible-price equilibrium

If the menu cost is zero, then nothing prevents stores from re-optimizing their price each month. In this case, there is an analytical solution of the model, which serves as a benchmark for the menu cost case. The derivation for the flexible price solution is in the appendix, where it is shown that the expected inflation rate can be expressed as:

$$E\left(\pi_t\right) = g_{PY} - E\left(g_{Yt}\right) = g_{PY} - E\left(g_{Zt}\right) + \frac{1}{1+\psi} E\left(g_{1+\bar{\tau},t}\right), \tag{21}$$

where $g_{1+\bar{\tau},t}$ is the shock to the average tax rate. The main finding is that in the absence of menu costs, there is no asymmetry in the pass-through after tax increases and decreases. The pass-through of tax changes into inflation is a decreasing function of ψ which is the inverse of the Frisch elasticity of labor supply. The reason is that higher labor supply elasticity leads to larger drop in equilibrium working hours and output, so the inflation effect will be higher assuming constant nominal GDP-growth. Also, these equations imply that without tax changes, real GDP-growth and inflation are determined by the growth rate of the aggregate technology shock $Z(t)$.

2.6 Model solution with menu costs

The model with menu cost does not have a closed form solution, so we need to solve it numerically. Following the standard Krusell-Smith, 1998 methodology, the algorithm is looking for a fixed point of the inflation forecasting equation parameters γ, such that the equation provides the best linear forecast under the optimal choices of the individual firms (using the same linear forecasting rule). The policy functions of the firms are obtained by value function iteration over a discretized state space. The main steps of the algorithm are[6]

1. Guess the parameters of the inflation forecasting equation $(\gamma_1, \gamma_2, \gamma_3, \gamma_4, \gamma_5, \gamma_6)$
2. Given this guess obtain the value function and the policy function of the firms.
3. Use the policy functions to simulate the behavior of the economy.
4. Obtain new estimates for the $\hat{\gamma}$ parameters from the aggregates of the simulated economy.
5. Do it until the guessed parameters γ and the obtained parameters $\hat{\gamma}$ are sufficiently close to each other.

3 Data

We estimate the model and the effect of various value-added tax changes on a data set containing store-level price quotes. These data are originally used to the monthly calculation of the Consumer Price Index in Hungary.

The data set contains price quotes between December 2001 and December 2006, which enables us to observe the frequency and magnitude of price *changes* in 60 consecutive months. In terms of product categories, we have price information about 770 different representative items; the total CPI-weight of these items is 70.12% in 2006. The missing representative items are mainly regulated prices, or in some cases methodological problems make it impossible to collect data from different stores (e.g. used cars, computers).

After an initial data analysis, we dropped another 220 representative items, so finally we ended up with 550 representative items with a total CPI-weight of approximately 45.3%. Among these excluded items there are the fuels, alcoholic beverages and tobacco, where frequent changes in oil prices, and/or frequent indirect tax changes make it difficult to estimate the effect of value-added tax changes. Another reason of these exclusions was that for these representative items the maximum length of price spells were constrained. This could be because the Central Statistical Office began data collection about these products at a later date. (Examples are LCD TV-s, memory cards, MP3 players etc.) The other typical reason of exclusion was seasonal data collection: for some products (cherries, gloves, skis etc) the statistical office collects price quotes only in certain, pre-specified months of the year. All in all, this way we ensured that the maximum length of the observed price spells exceeds 36 months (3 years) for each representative item, which we regard long enough to get reliable estimates. The sample coverage (in 2006) by main CPI-categories is illustrated in Table 1[7].

[6] The details of the algorithm are found in the appendix. The Matlab code is available on request.

[7] The single 'Energy' item (propan-butan gas) remaining after the exclusions is included in the 'Other goods' category

Table 1: Coverage of the data set by CPI-categories

	CPI basket		Original sample		Final sample	
CPI category	**Weight**	**Items**	**Weight**	**Items**	**Weight**	**Items**
Food, alcohol, tobacco	31.842	222	31.322	220	20.272	162
Unprocessed food	5.665	53	5.665	53	4.151	34
Processed food	26.177	169	25.657	167	16.121	128
Proc. food excl. alc, tob	17.427	139	16.907	137	16.121	128
Clothing	5.305	171	5.305	171	3.147	101
Durable goods	9.240	112	4.976	73	3.562	49
Other goods	15.277	214	12.979	192	7.852	159
Energy	13.203	16	6.350	8	0.723	1
Services	25.134	161	14.679	106	9.789	78
TOTAL	**100.000**	**896**	**70.122**	**770**	**45.346**	**550**

These 550 representative items in the data set can be regarded as 550 mini panels, containing time series of price quotes from different outlets. As an example, consider item 10001 "Bony pork rib with tenderloin": the data set contains 7,922 observations from 162 different outlets, i.e. 48.9 price quotes per outlet. Moreover, for 96 of the 162 stores we have data for each month. As it is true for most of the representative items in the data set that the list of observed outlets is typically unchanged, the data is appropriate to investigate store-level developments in the prices, and also the pricing behavior of different stores.

On average, there are approximately 6,566 observations per representative item in the data set, which means that the total number of observations exceeds 3.6 million (3,611,335).

Our analysis will focus on regular prices, rather than sales prices. The price collectors of the Central Statistical Office use a sales flag to identify sales prices (i.e. prices that are temporarily low, and have a "sales" label), and we use these flags to filter out sales prices in the first round. After this we also filter out any remaining price changes that are (1) at least 10 %, (2) and are completely reversed within 2 months.

Individual products in Hungary are categorized into 3 distinct groups and face 3 distinct VAT tax rates. There are only a few products with extra subsidy (drugs, school books) facing the lowest tax rates. As these products only constitute 2.1% in CPI-weights of our sample, we will ignore these. The two other categories constitute the 46.8% and 51% of our sample. Table 2 presents the changes in the various tax rates. In January 2004 the middle rate was increased from 12% to 15%, and this same rate was increased again in September 2006 from 15% to 20%. The top rate, meanwhile was decreased in January 2006 from 25% to 20%. Ex post, the tax changes can be seen as a stepwise convergence of the middle and the top tax rates, though this outcome was not explicitly expressed ex ante by the fiscal authorities.

It also should be noted that the monetary authority has expressed in advance that it is not planning to react to direct effects of the VAT shocks, and it is prepared to fully accommodate their full effects. Looking at nominal consumption growth (though the available data, shown in figure 3 is quarterly only), we see that the growth rate was volatile, but relatively flat throughout 2006, and if anything it dropped more at the month of the tax decrease (with lower inflation effect), *decreasing* the observed inflation asymmetry.

Table 2: VAT rates in Hungary

VAT-rates in Hungary	Lower	Middle	Top
– Dec 31, 2003	0%	12%	25%
Jan 1, 2004 – Dec 31, 2005	5%	**15%**	25%
Jan 1, 2006 – Aug 31, 2006	5%	15%	**20%**
Sep 1, 2006 –	5%	**20%**	20%

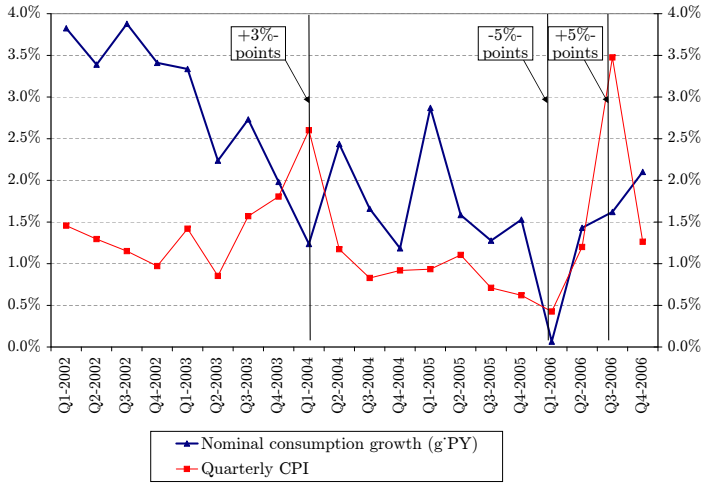

Figure 2: Nominal consumption growth in Hungary, 2002-2006, quarterly

3.1 Data moments

We estimate model parameters by matching some data moments. This means that for an arbitrary combination of model parameters, we solve the model, simulate hypothetical data, calculate the "theoretical" moments, and compare them with the same moments estimated from data.

At the heart of this procedure is the choice of moments, upon which the matching is based. Our choice is similar – though not identical – to the one by Klenow-Willis (2006), as we also use some extra moments to account for the tax changes:

- mean sectoral monthly inflation rate ($\bar{\pi}$);
- (time-series) standard deviation of sectoral monthly inflation rate ($\sigma_{\bar{\pi}}$);
- autocorrelation of the sectoral monthly inflation rate ($\rho_{\bar{\pi}}$)
- frequency of price changes ($\bar{I}$);
- average size of non-zero price changes ($\bar{\Delta P}$);

- autocorrelation of new relative prices (ρ_p);
- inflation effect of unit value-added tax increases and decreases (γ^+, γ^-).

The calculation of these data moments is described in the appendix, here we only show how the immediate VAT-effects are calculated. To be consistent with the model simulations, we calculate these effects from time-series data.[8] We estimate the following time-series regression for each representative item (j):

$$\pi_t^j = \beta_0 + \sum_{k=1}^{11} \beta_k^j (MONTH = k)_t + \beta_{12}^j VAT04J_t + \beta_{13}^j VAT06J_t + \beta_{14}^j VAT06S_t + \varepsilon_t, \quad (22)$$

where the explanatory variables are month dummies, and other dummies corresponding to the 2004:01, 2006:01 and 2006:09 value-added tax changes. So the inflation effect of the 2006:01 VAT decrease and the 2006:09 VAT increase are estimated by $\left(\widehat{\beta}_{13}^j, \widehat{\beta}_{14}^j\right)$ respectively, and the overall inflation effects are given by

$$\gamma^+ = \sum_j w_j \widehat{\beta}_{14}^j, \quad \gamma^- = \sum_j w_j \widehat{\beta}_{13}^j, \quad (23)$$

with w_j being the relative consumer expenditure weights of specific representative items.

3.2 Stylized Facts

This section presents some qualitative statements obtained from observation of the data moments. The first couple of stylized facts are about the development of the moments during 'normal' times. It claims that the prices in terms of their major moments behave very similarly as was documented by numerous studies using CPI data. It also emphasizes the heterogeneity across sectors.

The second part of the section presents the stylized facts about the development of major moments as a response to the tax changes. Other than emphasizing the asymmetry of inflation effects, the section shows that the fraction of the price changing firms clearly increased as a result of the tax shocks supporting 'state-dependent' pricing models, and is clearly smaller than 1 even in the subsample of directly affected products contrary to predictions of a frictionless, flexible price model. It also presents the result that the absolute magnitude of average price-changes decreases as a result of the tax shocks.

3.2.1 Stable frequency and large average absolute size of price changes.

Figure 3 and 4 show the fraction of firms changing prices and the magnitude of average price changes respectively in our sample. The average frequency of price change is fairly stable around the average (12%) level during 'normal' times in line with findings of previous studies (see e.g. Klenow and Krystov, 2008). Its relatively low level can be explained by the fact that some sectors with very flexible prices (e.g. fuel) were excluded from the sample (with

[8] The VAT effects calculated by Gabriel-Reiff, 2007 using panel estimations and the (averaged) time-series method provided very similar results.

these sectors the average frequency is 18.5%). The average absolute size of the price changes during 'normal' times – also in line with other studies – is high: it fluctuates around its average of 11.5%.

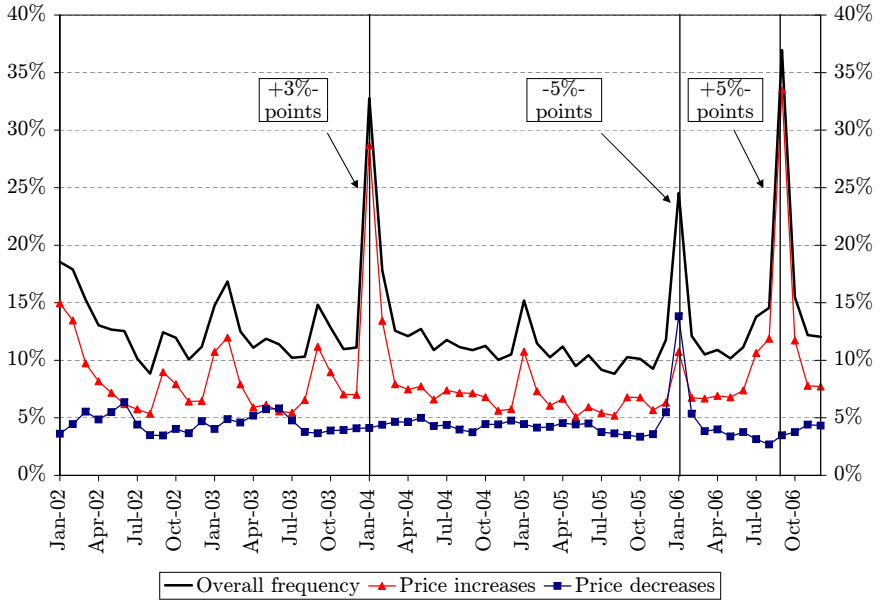

Figure 3: Frequency of price changes

3.2.2 Large sectoral heterogeneity

Table 3 reports the calculated moments for the main CPI-categories for each examined sectors, with frequency and sizes calculated for 'normal' non-tax change months and tax change months as well. We can interpret these figures as moments calculated from a "typical" representative item in the respective CPI-categories. The model parameters are about to be calibrated to for each CPI-category to match these typical representative items.

Table 3: Estimated data moments

CPI category	$\overline{\pi}$	$\overline{\sigma}$	$\overline{I}_{NT}$	$\overline{I}_{T}$	$\overline{\Delta P}_{NT}$	$\overline{\Delta P}_{T}$	ρ_p
Proc. food	0.429%	0.91%	0.134	0.529	0.099	0.088	0.010
Unproc. food	0.282%	2.64%	0.322	0.640	0.116	0.111	0.230
Clothing	-0.119%	0.63%	0.065	0.089	0.157	0.137	-0.123
Durable goods	-0.260%	0.51%	0.088	0.130	0.100	0.093	-0.100
Other goods	0.125%	0.62%	0.094	0.2	0.110	0.105	-0.076
Services	0.699%	0.67%	0.063	0.252	0.138	0.111	-0.031
All	**0.324%**	**0.91%**	**0.120**	**0.356**	**0.115**	**0.102**	**-0.013**

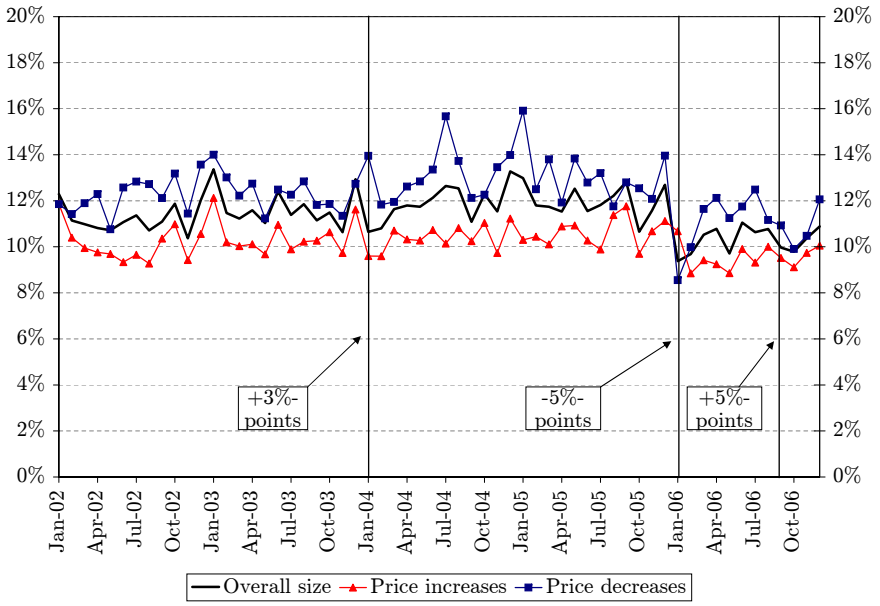

Figure 4: Average absolute size of price changes

Table 3 shows that both in terms of the sectoral inflation rates and price rigidities the sectors show sizable heterogeneity. The two largest sectors affected by all of the tax changes (the processed food and the services sectors) are also different: the services sector facing higher inflation rate and higher price rigidity than the processed food sector.

3.2.3 Asymmetric inflation effect

Table 4 presents the estimated VAT effects for each sectors. As the relative weights of products facing the middle and top tax rates are different across sectors (see the columns 'middle' and 'top'), the overall inflation effects calculated from the regressions 50 are not directly comparable. We can, however, compute the inflation effect of a unit tax increase which is directly comparable (presented in the last three columns of Table 4) and calculated by dividing the overall effect by change in the average sectoral tax rate.

3.2.4 Frequency increases, size decreases during VAT shocks

The fraction of firms changing prices clearly increases (31.4% from 13%) in the months of the tax changes, as it can be seen in Figure 3, in line with international evidence (see Gagnon, 2007 and the references there.) It should be noted, however, that the frequency of price changing firms is strictly less than 1 even in the subsamples of directly affected firms (around 60%) providing some evidence against fully flexible price setting.

The average size of price changes during the months of tax changes decreases (from 11.5% to 10.2%) as it can be seen in Figure 4. The systematic nature of this outcome is underlined

Table 4: Estimated VAT effects

CPI category	**CPI weights**		**Overall**		**Unit**	
	middle	top	VAT06j	VAT06s	VAT06j	VAT06s
Proc. food	78.8%	21.2%	-0.770	3.471	0.839	1.044
Unproc. food	100%	0%	-0.137	4.724	NA	1.11
Clothing	1%	99%	-1.085	0.212	0.268	NA
Durable goods	0%	100%	-1.661	0.602	0.407	NA
Other goods	17.8%	82.1%	-1.032	0.881	0.303	1.203
Services	32.1%	67.9%	-0.848	1.413	0.299	1.06
All	**47.9%**	**52.1%**	**-0.870**	**2.200**	**0.395**	**1.104**

by the fact that it is also true in all of the sectors, as Table 3 shows.

4 Results

We calibrate parameters to hit some important moments of the data, and our main interest is whether the model is able to explain the (asymmetric) response of the inflation rate to the tax changes. After presenting the parameterization strategy, this section is going to present calibrations using the standard model and then it introduces heterogeneous menu costs and sectoral calibrations in steps. The 2 major sectoral calibrations we present are those of the processed food and the services sectors. These two sectors are the largest in our sample – with 16.1% and 9.8% overall CPI weights – both including products facing both tax rates. The services sector has both higher trend inflation and stickier prices and, thus, our model predicts higher inflation asymmetry. We show that this is the case in the data and our model is able to quantitatively reproduce these facts.

4.1 Parameterization

We fix some parameters exogenously. We calibrate $\beta = 0.96^{1/12}$ implying 4% yearly real rate, and the mean aggregate nominal growth rate to $g_{PY} = 0.0934 \cdot (1/12)$, which is the average monthly nominal consumption growth in Hungary over the period 2002:01-2006:12. We set the value of θ determining the level of competition within a sector to 5, which is a usual number used in the industrial organization literature. There is no agreement about the right value of θ in the menu cost literature, the values range from 3 (Midrigan, 2006) to 11 (Gertler-Leahy, 2007). The choice of θ though influences our estimates of menu costs and the standard deviation of idiosyncratic shocks, has no observable effect on our estimates on the estimated inflation asymmetry within this range.

The other parameters of the model are calibrated to match some important sectoral pricing characteristics. The mean of the sectoral technology growth μ_{g_Z} is calibrated to make the simulated inflation rate equal to the mean sectoral inflation $\bar{\pi}$. The persistence of the aggregate technology shock, similarly, is set to be equal to the persistence of the (seasonally and VAT adjusted) inflation rate $\rho_{g_Z} = \rho_{\bar{\pi}}$. Other parameters do not have a clear one-to-one relationship between one moments, but we have a good idea how they influence the moments we would like to hit. The standard deviation of the sectoral technology

growth σ_{g_Z} increases the estimated average standard deviation of the inflation rate $\sigma_{\bar{\pi}}$, the frequency and marginally the size of the price changes. The persistence of the idiosyncratic technology shock ρ_A increases the persistence of the relative price developments ρ_p and its standard deviation σ_A increases the frequency, the size of the price changes. The menu costs ϕ decreases the frequency and increases the size of the price changes. The labor supply elasticity $1/\psi$ influences the inflation effects of the tax changes, as it influences how much of its effect is buffered by the relative wage and thereby the cost adjustment. Higher labor supply elasticity (lower ψ) implies lower wage response, thereby higher inflation effects of the tax change.

So to sum up, we have 8 'moving' parameters $(\mu_{g_Z}, \rho_{g_Z}, \sigma_{g_Z}, \rho_A, \sigma_A, \phi_{NT}, \phi_T, \psi)$ and we use them to hit 9 major moments (we call them 'matched moments') of the data $(\bar{\pi}, \rho_{\bar{\pi}}, \sigma_{\bar{\pi}}, \rho_p, \bar{I}_{NT}, \bar{I}_T, \Delta P_{NT}, \Delta P_T, \bar{\pi}_T)$, where the subscripts NT and T refer to averages during 'normal' times and tax shocks respectively, and the $\bar{\pi}_T$ is the average inflation effects of the tax changes. The 'unmatched moments' for which we have no independent parameters to hit are the effects of the positive and negative unit tax changes $(\gamma^{\pm})$, which are used to evaluate the model's performance.

4.2 Calibration results

Table 5 presents the parameters for each calibrations (for the aggregate sample: homogeneous and heterogeneous menu costs; for the sectoral samples: processed food and services) and the estimated forecasting equations with the resulting goodness of fit parameters. Table 6 presents the values of the moments which were directly used for the calibration ('matched moments') and the resulting values of the 'unmatched' moments including the asymmetry estimates.

4.2.1 Standard model

In this section, we discuss the calibration results of the standard model calibrated on the aggregate sample with homogeneous menu costs (first column in tables 5 and 6). The model is able to hit the moments in 'normal' times with reasonable parameter values, but the model predicts essentially no asymmetry and misses the frequency and the size of price changes at the months of the tax changes.

The overall frequency of price change during 'normal times' is 12.0% and the size of the price changes is 11.5% which are somewhat different from those found in the CPI data in other countries[9] showing that our sample overrepresents sectors with less price flexibility (we have excluded energy, alcohol and tobacco for example, because of their different tax development). Similarly to previous menu cost models with idiosyncratic shocks, the model needs persistent idiosyncratic shocks ($\rho_A = 0.76$) with large standard deviation ($\sigma_A = 5.3\%$) to be able to hit the large average size and frequency of the price changes.

The menu cost is estimated to be 3.0% when paid, but note that it is only paid in case of price change which – under no tax change – happens with 12% probability. It means that the yearly menu cost proportional to the firms' revenue is estimated to be 0.36%, which is within

[9] Nakamura and Steinsson, 2008, for example, reports median frequency of 21.1% and size at 8.5% for the US data

the range of estimated menu cost levels in previous studies (Klenow-Willis, 2006 estimates a yearly cost of 1.4%, while Nakamura-Steinsson, 2007 finds this measure to be 0.2%).

The forecasting equations show that the estimated parameters of most of the state variables in equation (18) are significant and are able to explain 97.1% of the variance of the inflation rate implying that assuming linear forecasting equation is a good approximation to the rational expectations equilibrium.

The standard model, however, is unable to explain the inflation asymmetry observed in the data. While the inflation effect of a unit VAT increase and decrease are estimated to be 1.09 and 0.40 respectively, the model only predicts 0.49 and 0.46. Furthermore, the frequency of price increases at the months of VAT changes (28.7%) is substantially underestimated (15.2%) and the size of price changes (8.9%) is overestimated (12.0%).

4.2.2 Introducing heterogeneous menu costs

As a shortcut to assuming multiproduct firms with decreasing marginal menu costs as in Midrigan (2006), we assume that the average menu costs are smaller at the months of the VAT changes. This assumption would explain both the increased frequency and the decreased size of the price changes observed at the months of tax changes.

Note that other explanations in the literature of monetary policy shocks proposing potential channels why prices move together would not be able to explain the underestimated inflation effect of the VAT shock. Assuming decreasing returns to scale or other real rigidities in the firm level (Klenow-Willis, 2007), for example, would further *reduce* the inflation effects of the VAT shocks. Similarly, as VAT shocks are aggregate supply shocks, regional labor markets as in Gertler-Leahy, 2007 would not help either, because the wage effect of a VAT shock can be expected to be negative, *reducing* the inflation effects. Intermediate products, as in Carvalho, 2006 and Nakamura-Steinsson, 2007 would not help either, because as firms only need to pay taxes for the value added they produce, it is the net inflation rate that influence their costs. The net inflation, however, is *dropping* at the month of VAT increase, thereby would reduce the inflation effect.

The results of the heterogeneous menu cost estimation is shown in the second column of tables 5 and 6. The results show that the best calibration is indeed one with lower menu costs during the VAT shock: the 2.4% menu cost during normal times is estimated to drop approximately to its half: 1.3%. With this calibration, the model is still able to hit the moments during 'normal' times with reasonably high goodness of fit of the forecasting equation (with R^2 over 90%), and – though somewhat overestimates both – it is able to reproduce the higher frequency and the lower size of the price change at the month of the tax changes. The results also show, however, that though the model predicts somewhat higher asymmetry (0.66-0.61 positive and negative inflation effects respectively) than in the homogeneous menu cost case, it still significantly underestimates the asymmetry observed in the data.

4.2.3 Sectoral calibrations

In this section, we show that by calibrating the model to sectoral moments, we are able to reproduce the asymmetry observed in the data. We argue that the reason for it is that the asymmetry in the model is a convex function of the inflation rate and the menu costs so a

Table 5: Calibrations

	Parameters	**Aggregate**		**Sectoral**	
		Hom.	**Het.**	**Food**	**Services**
σ_{g_Z}	Std. dev. of sectoral technology growth	8.5%	9.0%	8.4%	7.6%
ρ_A	AR. parameter of the idiosyncratic shocks	0.76	0.73	0.70	0.0
σ_A	Std. dev. of idiosyncratic shocks	5.3%	5.1%	4.4%	7.1%
ϕ_{NT}	Menu costs during 'normal' times	3.0%	2.4%	1.6%	4.5%
ϕ_T	Menu costs during tax changes	3.0%	1.3%	0.85%	2.2%
ψ	Inverse of the Frisch elasticity of labor supply	0	0	0	0
	Forecast of π^s_{t+1} as a function of	**Aggregate**		**Sectoral**	
		Hom.	**Het.**	**Food**	**Services**
γ_1	Constant	0.0041	0.013	0.014	0.028
		(0.001)	(0.001)	(0.001)	(0.002)
γ_2	Past inflation π^s_{t-1}	0.41	0.30	0.35	0.29
		(0.004)	(0.007)	(0.007)	(0.01)
γ_3	Cost parameter ξ^s_{t-1}	-.001	0.02	0.03	0.05
		(0.002)	(0.003)	(0.003)	(0.004)
γ_4	Technology growth g^s_{Zt}	-.57	-.58	-.49	-.35
		(0.004)	(0.006)	(0.005)	(0.01)
γ_5	Positive average tax growth $g^+_{\bar{\tau},t}$	0.49	0.69	0.72	0.84
		(0.006)	(0.01)	(0.01)	(0.01)
γ_6	Negative average tax growth $g^-_{\bar{\tau},t}$	0.48	0.61	0.62	0.27
		(0.006)	(0.01)	(0.01)	(0.01)
R^2	Goodness of fit	97.1%	92.5%	91.4%	75.3%

standard Jensen's inequality applies. We argue that because of this bias, sectoral estimation is necessary to obtain valid estimates for the aggregate asymmetry of monetary policy shocks.

The sectoral estimations can also be considered successful at hitting the major moments. The major difference between the two sectors is that the services sector faces both higher yearly trend inflation (8.7% compared to 4.9%) and higher estimated menu costs (4.5% compared to 1.6%[10]) than the processed food sector. These differences can fully explain both the substantially lower fraction of price changing firms in the services sector (6.3% compared to 13.3%) and the substantially higher average absolute size of price changes in the services sector (13.8% compared to 9.9%) than in the processed food sector. The calibrated idiosyncratic technology processes are somewhat different: the process is estimated to be a white noise with relatively high standard errors (7%) at the services sector, while, in the processed food sector, the process is more persistent (0.7) with lower standard error (even

[10]Implying reasonable 0.28% and 0.21% yearly expected menu costs

Table 6: Moments

Matched Moments			**Aggregate**		**Sectoral**	
			Hom.	**Het.**	**Food**	**Services**
$\bar{\pi}$	Average inflation rate	Data	0.33%	0.33%	0.43%	0.7%
		Estimate	0.33%	0.33%	0.43%	0.7%
$\sigma_{\bar{\pi}}$	Std. dev. of inflation rate	Data	0.55%	0.55%	0.56%	0.36%
		Estimate	0.56%	0.55%	0.56%	0.37%
ρ_p	Autocorrelation of new relative prices	Data	-0.013	-.013	0.01	-.03
		Estimate	-0.013	-.013	0.01	0.03
$\bar{I}_{NT}$	Frequency during 'normal' times	Data	12.1%	12.1%	13.4%	6.8%
		Estimate	12.1%	12.1%	13.4%	6.8%
ΔP_{NT}	Size during 'normal' times	Data	11.5%	11.5%	9.8%	13.5%
		Estimate	11.5%	11.5%	9.8%	13.5%
$\bar{I}_T$	Frequency during tax changes	Data	28.6%	28.6%	34.2%	18.8%
		Estimate	15.2%	30.3%	33.2%	22.6%
ΔP_T	Size during tax changes	Data	8.9%	8.9%	8.9%	9.3%
		Estimate	12.0%	9.9%	8.5%	11.8%
Unmatched Moments			**Aggregate**		**Sectoral**	
			Hom.	**Het.**	**Food**	**Services**
γ^+	Inflation effect of a unit tax increase	Data	1.09	1.09	1.04	1.05
		Estimate	0.49	0.66	0.72	0.94
γ^-	Inflation effect of a unit tax decrease	Data	0.40	0.40	0.85	0.30
		Estimate	0.46	0.61	0.61	0.35

its unconditional standard error is somewhat lower than that of the service sector with σ_A^u being 6.2%).

The menu costs at the month of the VAT shocks are calibrated to reduce approximately to half of its normal value: it is estimated to reduce to 0.85% (from 1.6%) for the processed food sector and to 2.2% (from 4.49%) for the services sector. The frequency and the size moments at the month of tax changes are somewhat underestimated in case of the processed food sector and overestimated in case of the services sector, though the directions and the magnitudes are hit. In the services sector estimation, the R^2 of the inflation forecasting equation is 75%, which suggests that the linear equation in this case might be improved by including higher order moments. However, as the marginal effect of better prediction of the inflation rate can be considered very limited.[11]

The sectoral calibrations, in line with the stylized facts, predict much lower asymmetry in the processed food sector than in the services sector. The model is also successful in quantitatively predicting the asymmetric inflation effects of the tax changes. In the processed food sector the inflation effects of unit tax increases are estimated to be 1.04 and 0.84 respectively, which are somewhat underestimated by the model which predicts coefficients of 0.72 and 0.61[12]. In the services sector, the coefficients of the unit tax changes are 1.06 and

[11] In a closely related model with rational expectations but no aggregate uncertainty, we have obtained very similar asymmetry predictions, which strongly supports the above conjecture.

[12] A possible reason for this is that the firms might have considered the tax changes more persistent ex ante

0.299 respectively, and the model is very successful in hitting these parameters by predicting 0.94 and 0.35 as coefficients for the tax increase and decrease respectively.

The main reason why the aggregate calibration underestimates the asymmetry is the result of a standard Jensen's inequality resulting from a sectoral estimation. Figure (5) presents the asymmetry (as the log difference between the unit effects of tax increase and decrease) predicted by the model for different values of trend inflation and menu costs. Disregarding the simulation errors, the model implies increasing asymmetry in both variables, and the fact that the constant slopes with respect to each variable as a function of the other are increasing suggests that the relationship can be approximated by a linear equation with a cross term in the form:

$$u = \log \gamma^+ - \log \gamma^- = f(\pi, \phi) \approx \delta_0 + \delta_1 \pi + \delta_2 \phi + \delta_3 \pi \phi,$$

where u stands for the asymmetry. Estimating this equation on a sample of 121 points simulated for figure (5) we find that while $\delta_0, \delta_1, \delta_2$ are all estimated to be zero, the estimate of the cross term is positive, large ($\hat{\delta}_3 = 0.35$) and highly significant (with a standard error of 0.017), and the equation explains 95.2% of the variance in the simulated asymmetry.

In this case, unless the sectoral inflation rates and the menu costs are uncorrelated, a convex relationship like this implies that the aggregate model with average inflation rate and menu costs will underestimate the asymmetry, and the Hungarian case shows that this bias can be substantial.

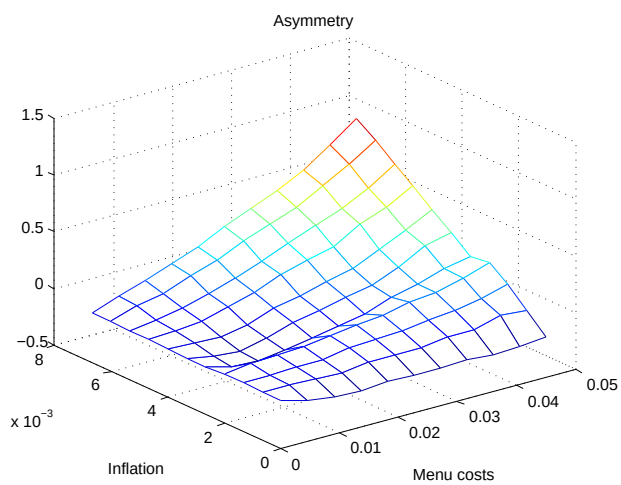

Figure 5: Simulated asymmetry as a function of inflation and menu costs

The positive relationship between the inflation rate and the price stickiness in our sample is because of the relative importance of sectors with both high inflation and high price stickiness. Figure 6 plots a histogram of the distribution of (123) sectors over their trend inflation between 2002 and 2006 and their average frequency of price change during 'normal' times, each sector weighted by their relative consumption expenditure. Considering the

than in our current simulations based on the ex post tax changing frequencies.

frequency of price change as a proxy for price stickiness, we see that the distribution is not symmetric: there are more mass at the high inflation-low frequency quadrant. As frequency is increasing with the inflation rate, the figure can be expected to underestimate the asymmetry of an inflation-menu cost distribution.

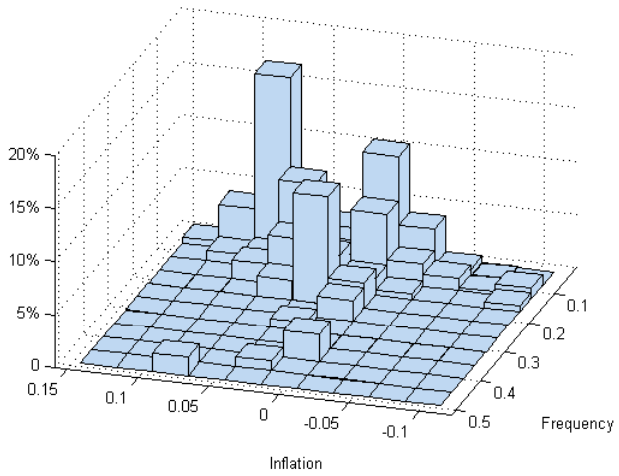

Figure 6: Weighted histogram of sectoral trend-inflation and frequency

4.2.4 Inflation is the major reason of the sectoral asymmetry

For our parametrization, the sole economically significant reason of the observed sectoral asymmetry is the trend inflation. In a similar framework as ours, Devereux and Siu, 2007 provided a different argument for asymmetry by showing that individual firms' strategic incentives are asymmetric: while prices are strategic complements in case of positive aggregate shocks to the (nominal) marginal cost, they are strategic substitutes in case of negative shocks. Running a counterfactual experiment resulting in 0 average inflation rate, we have found no significant asymmetry in any of the sectors, and in some cases we even found *negative* asymmetry (larger negative than positive effects).

The channel responsible for the negative asymmetry at zero inflation results from the asymmetric profit function, the same characteristic emphasized by Devereux and Siu, 2007. In our model, this implies asymmetric steady state relative price distribution with higher median than mean, because more firms are willing to have nominal prices above the aggregate price level than below it. But this asymmetry implies that in case of a negative shock (with no inflation channel) more firms are going to found themselves outside their inactivity range and decrease their prices, than in case of a positive shock. This distribution effect can be expected to be much lower in the model of Devereux and Siu, 2007, where the maximum length of a price contract is limited to 2 periods.

Devereux and Siu, 2007 found that this strategic asymmetry is larger with more intense competition (higher elasticity of substitution parameter θ). But our negative asymmetry

results were true even if we increased the level of competition substantially ($\theta = 11$), implying that the 5%-points tax changes were not large enough in our model to make the strategic asymmetry overcome the effects of the distributional asymmetry.

4.2.5 Monetary shock

The calibrated model can provide estimates to the asymmetric inflation effects of symmetric large monetary policy shocks. Other than providing a contribution to an important policy question, the exercise is necessary to assess what percentage the observed inflation asymmetry was the result of the asymmetric distribution of the specific VAT shocks.

The model without real rigidities (for most of our calibrations, $\psi = 0$ implying constant (normalized) real wages) implies similar inflation effects for a similar size VAT and monetary policy shocks. In a monetary policy simulation using the aggregate parameter estimates of the model with heterogeneous menu costs, the asymmetry of a large permanent increase in the level of nominal output ($P_t Y_t$) would predict a limited asymmetry (with unit effects of 0.83-0.76 for the positive and negative shocks respectively) similarly to the aggregate VAT estimation.

To get a sense what the sectoral estimation would imply, the first step is to calculate the weighted average of the inflation effects of unit VAT effects. Averaging these, instead of the inflation rates, we implicitly control for the asymmetric distribution of the VAT shocks across sectors. Doing this, we find that the average inflation effect of a unit VAT increase would be 0.92, while the effect of a VAT decrease would be 0.47. These estimates, though imply lower asymmetry than the headline numbers of 1.1-0.4 (which do not take sectoral differences into consideration) still imply that the inflation effect of a positive shock can be almost twice as large as the inflation effect of a negative shock.

Though the sectors with both positive and negative VAT changes cover 76% of our sample, there are 3 sectors (unprocessed food, clothes and durables) which were only hit by one of the tax changes. The model can be used to simulate the effect of the missing tax (or monetary policy) shocks in these sectors, and obtain an estimate for the sectoral asymmetry. As all of these sectors had relatively low inflation rates (2.3%, -1.2%, -2.7%), and – except for the clothes sector with the lowest inflation rate – low menu costs (1.2%,6.5%,1.8%), we have found limited asymmetry in these sectors. Taking them into consideration, we have found that the inflation effect of a positive monetary shock would be approximately 75% larger than the inflation effect of a negative monetary shock.

5 Conclusion

The paper presented a sectoral menu cost model calibrated to fit some key moments of the sectoral price development in Hungary between 2002-2006 in order to explain the observed asymmetric inflation response to major VAT changes of 2006. The paper deviated from the standard framework by introducing heterogeneous menu costs and calibrated the model on sectoral level. The paper found that (sectoral) trend inflation can successfully account for the observed asymmetry, thereby it provided direct evidence to the argument of Ball and Mankiw, 1994. Simulations suggested that a positive monetary policy shock would have almost twice as large inflation effect as a symmetric negative one.

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6 Appendix

6.1 Calculation of the flexible price equilibrium

This section presents the derivations for obtaining the expected inflation pass through in our model under flexible prices. [13]

Solving the firms' profit maximization problem, it is easy to derive that the optimal relative price is

$$p_i^*(t) = \frac{\theta\zeta_t}{\theta - 1}\frac{1+\tau_t(i)}{A_t(i)}, \tag{24}$$

with $\zeta_t = w_t \frac{C_t}{Z_t}$, and $w_t = \frac{\widetilde{w}_t}{P_t C_t}$ being the normalized nominal wage.

Then the optimal relative consumptions are $\frac{C_t^*}{C_t/n} = p_t^*(i)^{-\theta}$, which implies

$$C_t^*(i) = \left(\frac{C_t}{n}\right)^{1-\theta} Z_t^{\theta}\left(\frac{\theta n w_t}{(\theta-1)}\right)^{-\theta}\left(\frac{1+\tau_t(i)}{A_t(i)}\right)^{-\theta}. \tag{25}$$

Aggregating these with the CES-aggregator $C_t = \left[\sum_i n^{-\frac{1}{\theta}} C_t(i)^{\frac{\theta-1}{\theta}}\right]^{\frac{\theta}{\theta-1}}$, we can derive that

$$\frac{(\theta-1)\,\eta}{\theta\zeta_t} = \left[\sum_{i=1}^{n} n^{-1}\left(\frac{1+\tau_t(i)}{A_t(i)}\right)^{1-\theta}\right]^{\frac{1}{1-\theta}} \equiv 1+\overline{\tau}_t, \tag{26}$$

where the summation is a CES-aggregate of individual "effective" tax rates $\frac{1+\tau_i(t)}{A_i(t)}$, denoted as an average tax rate $1+\overline{\tau}(t)$.

[13] Sectoral subscripts s are suppressed for notational convenience.

With this average tax rate we can write the optimal individual relative prices as

$$p_t^*(i) = \frac{(1+\tau_t(i))/A_t(i)}{1+\overline{\tau}_t}, \tag{27}$$

and relative outputs as

$$\frac{C_t^*(i)}{C_t^*/n} = \left[\frac{(1+\tau_t(i))/A_t(i)}{1+\overline{\tau}_t}\right]^{-\theta} \tag{28}$$

which says that the optimal relative prices and relative outputs are determined by the relative effective tax rates (i.e. the ratio of the individual effective tax rates $\frac{1+\tau_t(i)}{A_t(i)}$ and the average tax rate $1+\overline{\tau}_t$).

The wage rate will be determined on the labor market by making labor demand and supply equal. Labor supply can be derived from the consumers' maximization problem. Rewriting equation (6) leads to

$$L_t = \left(\frac{\widetilde{w}_t}{\mu P_t C_t}\right)^{\frac{1}{\psi}} = \left(\frac{w_t}{\mu}\right)^{\frac{1}{\psi}}, \tag{29}$$

while the labor demand equation (11) can be written as:

$$L_t(i) = \frac{C_t^*(i)}{Z_t A_t(i)} = \frac{(\theta-1)}{\theta n w_t\left[1+\overline{\tau}_t\right]A_t(i)}\left[\frac{(1+\tau_t(i))/A_t(i)}{1+\overline{\tau}_t}\right]^{-\theta}. \tag{30}$$

Aggregate labor demand is the sum of individual demands. A little algebra shows that the equilibrium wage rate is

$$w_t = \mu^{\frac{1}{1+\psi}}\left(\frac{(\theta-1)}{\theta}\right)^{\frac{\psi}{1+\psi}}\left[\sum_{i=1}^{n}\frac{1}{1+\tau_t(i)}\frac{\left[(1+\tau_t(i))/A_t(i)\right]^{1-\theta}}{n\left[1+\overline{\tau}_t\right]^{1-\theta}}\right]^{\frac{\psi}{1+\psi}}, \tag{31}$$

where the last term is a weighted average of $\frac{1}{1+\tau_t(i)}$-s (the weights sum to 1 by the definition of $1+\overline{\tau}_t$ in equation (26)), and can therefore be written as another average of individual tax rates: $\frac{1}{1+\widetilde{\tau}_t}$. Therefore the equilibrium wage rate is simply

$$w_t = \mu^{\frac{1}{1+\psi}}\left(\frac{(\theta-1)}{\theta\left(1+\widetilde{\tau}_t\right)}\right)^{\frac{\psi}{1+\psi}}, \tag{32}$$

a function of deep parameters and individual tax rates.

With this equilibrium wage we can derive the level of individual outputs and prices. Rearranging the aggregation equation (26), we can write $w(t)\frac{C_t}{Z_t} = \zeta_t = \frac{(\theta-1)}{\theta[1+\overline{\tau}_t]}$, which implies that the real GDP path is

$$C_t^* = nZ_t\frac{(\theta-1)}{\theta n w_t\left[1+\overline{\tau}_t\right]} \approx Z_t\left(\frac{(\theta-1)}{\theta\mu\left[1+\overline{\tau}_t\right]}\right)^{\frac{1}{1+\psi}}, \tag{33}$$

(where we used the approximation $1+\overline{\tau}_t \approx 1+\widetilde{\tau}_t$).

The aggregate price level is the ratio of the nominal GDP (which is exogenous) and the real GDP:

$$P_t^* = GDP_t/C_t^* \approx \frac{1}{Z_t}\left(\frac{\theta\mu\left[1+\overline{\tau}_t\right]}{(\theta-1)}\right)^{\frac{1}{1+\psi}}. \tag{34}$$

The expected growth rates can be calculated easily. From the wage equation (32), we have

$$E\left(g_{w,t}\right) = -\frac{\psi}{1+\psi}E\left(g_{1+\widetilde{\tau},t}\right) \approx -\frac{\psi}{1+\psi}E\left(g_{1+\overline{\tau},t}\right). \tag{35}$$

From the real GDP-equation (33), it follows that

$$E\left(g_{Yt}\right) = E\left(g_{Ct}\right) \approx E\left(g_{Zt}\right) - \frac{1}{1+\psi}E\left(g_{1+\overline{\tau},t}\right). \tag{36}$$

Finally, from the price level equation (34), inflation is the difference between the nominal GDP-growth (g_{PY}, given exogeneously) and real GDP-growth:

$$E\left(\pi_t\right) = E\left(g_{Pt}\right) = g_{PY} - E\left(g_{Yt}\right) = g_{PY} - E\left(g_{Zt}\right) + \frac{1}{1+\psi}E\left(g_{1+\overline{\tau},t}\right). \tag{37}$$

6.2 Numerical algorithm of the menu cost model

The state variables of the model are $(p_t(i), A_t(i), \pi_t, \zeta_t, g_{Zt}, \tau_t^1, \tau_t^2, \phi_t)$[14], where the first 2 are idiosyncratic variables, while the remaining 6 are aggregate variables. The forecasts of the endogenous aggregate state variables (π_t, ζ_t) are obtained by the linear inflation forecasting equation 18 and 20.

Using these forecasting equations, we obtain a transition matrix P_{aggr} over the aggregate state variables $(\pi_t, \zeta_t, g_{Zt}, \tau_t^1, \tau_t^2, \phi_t)$ – determining the probabilities of all the possible states next period as a function of the current state – by building a VAR system. The VAR is of the form:

$$\begin{pmatrix} \hat{\pi}_t \\ \log\hat{\zeta}_t \\ g_{Zt} \\ \log\tau_t^1 \\ \log\tau_t^2 \\ \phi_t \end{pmatrix} = A_0 + A_1 \cdot \begin{pmatrix} \pi_{t-1} \\ \log\zeta_{t-1} \\ g_{Zt-1} \\ \log\tau_{t-1}^1 \\ \log\tau_{t-1}^2 \\ \phi_{t-1} \end{pmatrix} + \Xi \cdot \begin{pmatrix} \epsilon_{\pi,t} \\ \epsilon_{g_Z,t} \\ g_{\tau^1,t} \\ g_{\tau^2,t} \\ g_{\overline{\tau},t}^{+} \\ g_{\overline{\tau},t}^{-} \\ \Delta\phi_t \end{pmatrix} \tag{38}$$

where straightforward algebra shows that

$$A_0 = \begin{bmatrix} \gamma_1 + \gamma_4\mu_{g_Z} \\ (1+\psi)(\mu_{g_{PY}} - \mu_{g_Z} - \gamma_1) \\ \mu_{g_Z} \\ 0 \\ 0 \\ 0 \end{bmatrix},$$

[14] With $(100, 29, 7, 7, 3, 2, 2, 2)$ grids respectively.

$$A_1 = \begin{bmatrix} \gamma_2 & \gamma_3 & \rho_{gZ}\gamma_4 & 0 & 0 & 0 \\ (1+\psi)\gamma_2 & 1-(1+\psi)\gamma_3 & -(1+\psi)\rho_{gZ}(1+\gamma_4) & 0 & 0 & 0 \\ 0 & 0 & \rho_{gZ} & 0 & 0 & 0 \\ 0 & 0 & 0 & 1 & 0 & 0 \\ 0 & 0 & 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 0 & 0 & 1 \end{bmatrix},$$

and

$$\Xi = \begin{bmatrix} 1 & \gamma_4 & 0 & 0 & \gamma_5 & \gamma_6 & 0 \\ -(1+\psi) & -(1+\psi)(1+\gamma_4) & (1+\psi) & (1+\psi) & -(1+\psi)\gamma_5 & -(1+\psi)\gamma_6 & 0 \\ 0 & 1 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 1 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 1 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 1 \end{bmatrix},$$

To obtain the parameters for this VAR system, the algorithm guesses inital parameters $(\gamma_1, \gamma_2, \gamma_3, \gamma_4\, \gamma_5, \gamma_6)$ of the inflation estimating equation (18) using the flexible price solution. From this, it obtains an estimate for the $\hat{\zeta}_t$ using equation (20). Forecast for g_{Zt} is determined by $\rho_{gZ} g_{Z,t-1}$ and the development of the tax rates (τ^1, τ^2) is simulated exogenously and from this $g_{\tau^1}, g_{\tau^2}, g_{\bar{\tau}}^{+}, g_{\bar{\tau}}^{-}$ are obtained. The transition matrix is obtained by simulating 5000 shocks for each element of the current aggregate state space and obtaining the percentages of getting into next period states. The transition matrix for the idiosyncratic technology level P_A is similarly obtained by the one variable method suggested by Tauchen (1986).

The initial guess for the value function is obtained using the flexible price equilibrium, and then it is iterated using the transition matrices P_{aggr}, P_A until convergence. From the value functions, we obtain the policy functions determining the states the firm is willing to change its price $P^{C/NC}$ and the level of new relative price in case of price change P^C. Using the linear interpolations of the policy functions, we simulate price developments of 5000 firms for 2000 periods. The firms within a sector are partitioned between those facing τ^1 and τ^2 tax rates according to the sectoral CPI weights.

We obtain the aggregate state variables from this simulated sample. The wage rate required to calculate the current sectoral cost factor ζ is obtained by equating the simulated labor demand to the simulated labor supply (for most of our calibrations, we set $\psi = 0$, when the model reduces to a linear labor supply model with constant (normalized) wage rate). Using these aggregate variables, we run an OLS regression of the forecasting equation (18) obtaining new estimates for γ. We are running the algorithm until the guessed and obtained parameters in the forecasting equation are sufficiently close to each other.

6.3 Calculating data moments

To describe the calculation of the *mean sectoral monthly inflation rate*, let us introduce some notation. We index time by t, representative items by s, and stores by i. Then the mean sectoral (i.e. representative item-level) inflation rate is

$$\pi_{st} = \sum_i \frac{\log P_{sit} - \log P_{si,t-1}}{N_i}, \tag{39}$$

where N_i is the number of stores observed both at time t and time $t-1$ in sector (representative item) s. From these we calculate average monthly inflation rates for the representative items by time aggregation:

$$\pi_s = \sum_t \frac{\pi_{st}}{T}, \tag{40}$$

and finally the mean monthly inflation rate for the whole economy (or broader CPI-categories) is obtained by aggregating over representative items:

$$\overline{\pi} = \sum_s w_s \pi_s, \tag{41}$$

where w_s are CPI-weights (we use the CPI-weights in 2006). Note that seasonal variation in monthly inflation rates does not affect our estimates as we are using price changes between January 2002 - December 2006 to calculate mean representative item-level inflation rates.

The time-series *standard deviation of sectoral monthly inflation rates* is calculated similarly. First we calculate the time-series standard deviation of π_{st} for each representative item:

$$\sigma_s^2 = \sum_t \frac{(\pi_{st} - \pi_s)^2}{T-1}, \tag{42}$$

and then calculate the weighted average of these across representative items:

$$\overline{\sigma} = \sum_s w_s \sigma_s. \tag{43}$$

As our theoretical model does not contain any seasonal variation, we calculate these measures on the seasonally adjusted π_{st} series (i.e. we subtract the estimated seasonal dummies).

The third moment that we use for matching is the *frequency of price changes*. These are again calculated at the representative item-level, and then aggregated across representative items:

$$\overline{I}_s = \sum_t \sum_i \frac{I\left(\Delta P_{sit} \neq 0\right)}{N_s}, \tag{44}$$

where $I\left(\Delta P_{sit} \neq 0\right)$ is a dummy for price changes, and N_s is the total number of observations for representative item s. The overall average frequency is then

$$\overline{I} = \sum_s w_s \overline{I}_s. \tag{45}$$

Again, seasonal variation in frequencies does not bias our frequency estimates as we use price change data between January 2002 - December 2006 for the frequency calculations.

The *average size of price changes* is calculated first at the representative item level:

$$\Delta P_s = \sum_{I(\Delta P_{sit} \neq 0)} \frac{|\Delta P_{sit}|}{N_{I_s}}, \tag{46}$$

where N_{I_s} is the total number of price changes for representative item s: $\sum_t \sum_i I\left(\Delta P_{sit} \neq 0\right)$. Then the average size across representative items is

$$\overline{\Delta P} = \sum_s w_s \Delta P_s. \tag{47}$$

Our fifth moment, the *autocorrelation of new relative prices* is taken from Klenow-Willis (2006) to calibrate the persistence of idiosyncratic shocks the hit the stores. To calculate this, we first obtain relative prices. Firm i's relative price in sector s is $p_{sit} = \log P_{sit} - \log \overline{P}_{st}$, where $\overline{P}_{st} = \sum_i P_{sit}/N_i$ is the average price at time t. We consider all relative prices that are newly set, and calculate the autocorrelation between these newly set relative prices at the store level:

$$\rho_{p,s,i} = \frac{\sum_{I(\Delta P_{sit} \neq 0)} (\log p_{sit} - \log \overline{p_{sit}}) (\log p_{si,t-\tau_{sit}} - \log \overline{p_{sit}})}{\sum_{I(\Delta P_{sit} \neq 0)} (\log p_{sit} - \log \overline{p_{sit}})^2}, \tag{48}$$

where $\overline{p_{sit}}$ is the average of newly set relative prices, and τ_{sit} is the time (in months) between the previous and current price change. The autocorrelation at the representative item level is the average of $\rho_{p,s,i}$-s across stores: $\rho_{p,s} = \sum_i \rho_{p,s,i}/N_i$, while the overall autocorrelation of newly set relative prices is

$$\rho_p = \sum_s w_s \rho_{p,s}. \tag{49}$$

Finally, we also control for the *inflation effect of the value-added tax increases and decreases*. To be consistent with the model simulations, where we calculate these VAT-effects from time-series data, we also calculated these inflation effects from time-series data.[15] Specifically, we estimated the following time-series regression for each representative item:

$$\pi_{st} = \beta_0 + \sum_{k=1}^{11} \beta_k (MONTH = k)_t + \beta_{12} VAT04J_t + \beta_{13} VAT06J_t + \beta_{14} VAT06S_t + \varepsilon_t, \tag{50}$$

where the explanatory variables are month dummies, and other dummies corresponding to value-added tax changes. So the inflation effect of the various value-added tax changes are estimated by $\left(\widehat{\beta}_{12}, \widehat{\beta}_{13}, \widehat{\beta}_{14}\right)$, and the overall inflation effects are

$$\sum_s w_s \widehat{\beta}_{12,s}, \sum_s w_s \widehat{\beta}_{13,s}, \sum_s w_s \widehat{\beta}_{14,s}. \tag{51}$$

[15] The VAT effects calculated by Gabriel-Reiff, 2007 using panel estimations and the (averaged) time-series method we are using do not necessarily imply the same results, but as it can be seen by comparing Tables **??** and 3 they are sufficiently close to each other