FEDERAL RESERVE BANK OF SAN FRANCISCO

WORKING PAPER SERIES

Sticker Shocks: Using VAT Changes to Estimate Upper-Level Elasticities of Substitution

Bart Hobijn
Arizona State University

Fernanda Nechio
Federal Reserve Bank of San Francisco

October 2015

Working Paper 2015-17
http://www.frbsf.org/economic-research/publications/working-papers/wp2015-17.pdf

Suggested citation:

Hobijn, Bart, Fernanda Nechio. 2015. "Sticker Shocks: Using VAT Changes to Estimate Upper-Level Elasticities of Substitution." Federal Reserve Bank of San Francisco Working Paper 2015-17. http://www.frbsf.org/economic-research/publications/working-papers/wp2015-17.pdf

Sticker Shocks: Using VAT Changes to Estimate Upper-Level Elasticities of Substitution*

Bart Hobijn
Arizona State University

Fernanda Nechio
FRB San Francisco

October 26, 2015

Abstract

We estimate the upper-level elasticity of substitution between goods and services of a nested aggregate CES preference specification. We show how this elasticity can be derived from the long-run response of the relative price of a good to a change in its VAT rate. We estimate this elasticity using new data on changes in VAT rates across 74 goods and services for 25 E.U. countries from 1996 through 2015. Our results point to an upper-level elasticity of between 1, at a high level of aggregation that distinguishes 12 categories of goods and services, and 3, at the lowest level of aggregation with 74 categories.

JEL classification codes: E19, E21, D12.

Keywords: demand elasticities, multisector model, heterogeneity, aggregation, VAT rates.

*The views expressed in this paper are those of the authors and do not necessarily reflect the position of the Federal Reserve Bank of San Francisco or the Federal Reserve System. This paper has benefited from comments and suggestions by Carlos Carvalho, Adam Shapiro, and seminar participants at the EEA 2015, ES World Congress 2015, Federal Reserve Bank of San Francisco, SED 2015, and the Federal Reserve Macro System Meetings 2015. We would like to thank Eric Hsu and Ben Bradshaw for their excellent research assistance. E-mail: bhobijn@asu.edu, fernanda.nechio@sf.frb.org.

1 Introduction

Nested constant-elasticity of substitution (CES) preferences are the workhorse functional form used in multisector macroeconomic models.[1] This is because they allow for a parsimonious representation of consumers' willingness to substitute between varieties within a particular expenditure category, as well as their willingness to substitute across the broad classes of goods and services that make up expenditure categories.

The substitutability of varieties is determined by a lower-level elasticity of substitution while the substitutability across expenditure categories is parameterized by an upper-level elasticity of substitution. There are ample estimates of the elasticity of substitution between varieties of narrowly defined goods or services.[2] These are essentially estimates of lower-level elasticities of substitution.[3]

There are, however, very few estimates of an upper-level elasticity of substitution. Those that exist are based on two different empirical approaches. The first approach is to estimate elasticities of substitution using micro-level data for many goods or countries.[4] However, the expenditures covered in such analyses are still only a small subset of all goods and services over which aggregate preferences are commonly defined. The second approach is to estimate a classical demand system using macroeconomic data.[5] Studies based on this approach run into the usual endogeneity problems that arise when movements in prices, which are used as explanatory variables, are jointly determined with variation in the dependent variables which are either quantities or expenditure shares.[6]

The literature reflects the absence of a broader based estimate for this elasticity by resorting to analytically convenient calibrations for this parameter, such as choosing a unit elasticity or setting it equal to the substitution across varieties, which imply a wide range for this elasticity.[7]

[1] Among the many papers in which they are used are Aoki (2001), Bouakez *et al.* (2009), Carvalho (2006), Carvalho and Nechio (2011), Carvalho and Nechio (2014), Eusepi *et al.* (2011), Hobijn *et al.* (2006), Long and Plosser (1983), Long and Plosser (1983), Midrigan (2011), Nakamura and Steinsson (2010), Ngai and Pissarides (2007), Woodford (2003).

[2] See, for example, Nevo (2001) for the market for cereals, Manning *et al.* (1987) for the demand for medical services, Petrin (2002) for the market for cars and minivans, Leslie (2004) for Broadway plays, and Broda and Weinstein (2006), Imbs and Mejean (2015), Simonovska and Waugh (2014), and Feenstra *et al.* (2014) for tradable goods.

[3] Specially if one assumes that price movements in these narrow goods categories do not significantly affect the overall price level (Dixit and Stiglitz, 1983).

[4] See, for example, Gabriel and Reiff (2010), Broda and Weinstein (2006), and Feenstra *et al.* (2014).

[5] For example, the "Almost Ideal Demand System" introduced in Deaton and Muellbauer (1980).

[6] See Berry (1994), and Berry *et al.* (2004) for a thorough discussion of simultaneity bias in such models.

[7] Examples of papers that have followed one or the other calibration approach can be found in the macroeconomics New Keynesian literature, such as Carvalho (2006) and Carvalho and Nechio (2011), among others, and in the trade literature (see Costinot and Rodríguez-Clare (2014) for a summary). These two branches of the economics literature have showed calibrations of the upper-level elasticity ranging from 1 to as large as 11,

In this paper we provide an estimate of the upper-level elasticity of substitution that is different from the two aforementioned approaches for two reasons. First of all, it is comprehensive in its scope. It is based on data that cover *all* expenditures in the basket of goods and services that make up the Harmonised Index of Consumer Prices (HICP), which is the main inflation gauge in the European Union. Secondly, it does not suffer from endogeneity bias. This is because we identify the upper-level elasticity of substitution by estimating the *long-run* response of relative prices to changes in relative VAT rates. These VAT changes are the proverbial "sticker shocks" referred to in the title.

We are not the first to study the impact of VAT changes on inflation.[8] What sets our analysis apart from previous ones is that it covers *all* goods and services in the HICP as well as that we use data for a large sample of countries. Moreover, ours is the first paper to use VAT changes for the estimation of the elasticity of substitution.

In principle, one could use these VAT changes as conventional instruments that act as exogenous supply shocks and use Two-stage Least Squares (2SLS) to estimate the upper-level demand elasticity. The problem is that this would require high frequency data on expenditure shares or quantities consumed by expenditure category. Such data is not available for the countries in our sample.[9] Our approach is to estimate the equivalent of the first-stage regression of such a 2SLS approach, and then use a common model of production and price setting to parse out our estimate of the upper-level elasticity of demand.

In particular, we derive the *long-run* response of relative prices to changes in relative VAT rates in a model of price setting under monopolistic competition, where demand is determined by two-tier nested CES preferences and goods are produced using a Cobb-Douglas production technology. This framework is very general and nests those commonly applied in trade models, Real Business Cycle (RBC) models, and New-Keynesian models.[10] Thus, our estimate is applicable under a very general set of assumptions that cover many macroeconomic models.

We show that, within this framework, the reduced-form long-run response of relative prices to changes in relative VAT rates only depends on two parameters; (*i*) the labor elasticity of output and (*ii*) the upper-level elasticity of substitution that is the focus of our analysis.

a frequent choice in the New Keynesian literature for the elasticity of substitution across varieties (Basu and Fernald, 1997).

[8]Studies that also do so are, for example, Karadi and Reiff (2007), Gabriel and Reiff (2010), Gautier and Lalliard (2014), Carare and Danninger (2008).

[9]The countries in our sample are, by construction, those that have VAT rates. In such countries national accounts are generally constructed from the production side of the accounts and detailed expenditure data is only collected at a low frequency. Even in countries that do have high-frequency expenditure data the fluctuations in these data at low levels of aggregation are imprecisely measured (Wilcox, 1992).

[10]Three examples of trade, RBC, and New-Keynesian models that fit into the framework we use are those in Costinot and Rodríguez-Clare (2014), Long and Plosser (1983), and Aoki (2001), respectively.

Hence, for a given labor elasticity of output, we can back out the upper-level elasticity of substitution from the long-run effect of VAT changes on inflation.

In order to estimate the reduced-form elasticity of interest, we use the local projection method introduced in Jordà (2005). In the context of our problem, this local projection method boils down to running a simple panel-data regression of goods-specific changes in VAT rates on cumulative goods-specific inflation rates for different horizons. Based on these regression results, we use the effects of VAT rate changes on inflation in the long term to back out the upper-level elasticity of substitution. In principle, this panel-data model can be estimated on a country-by-country basis. However, since VAT changes occur rather infrequently, this results in imprecise estimates of the reduced-form elasticity. For this reason we pool our regression across countries.[11]

The construction of this evidence and the estimation of our panel-data model requires disaggregated data, by expenditure category for each country, on inflation as well as on VAT changes. We measure inflation using monthly inflation rates for all categories of expenditures that are included in the HICP. Monthly data on VAT rates by expenditure category and country are not available. We construct them from two administrative sources, namely European Commission (2015), and Eurostat (2015). The result is a dataset of monthly inflation and VAT rates for 74 expenditure categories and 25 countries, that covers the period from January 1996 to January 2015.

Because the degree of substitutability of expenditure categories depends on the level of granularity at which they are defined, we estimate the upper-level elasticity of substitution (separately) for the three levels of aggregation at which the data are provided. These correspond to 1-, 2-, and 3-digit expenditure classifications which are the basis for the HICP measure. Our estimate of the upper-level elasticity of substitution at the highest level of aggregation, at which expenditures are split up into 12 categories, is one. This supports the common choice of Cobb-Douglas preferences at this high level of aggregation. At the lowest level of aggregation, we find an elasticity of substitution of approximately 3. Given the standard error around this estimate, the upper bound of the 95 percent confidence interval on the latter estimate is about 5.

These results provide useful guidance for the choice of the value of the upper-level elasticity of substitution in multisector macroeconomic models. The value of the upper-level elasticity of substitution is important in many models. For example, it affects the magnitude of the

[11]For the estimates obtained using this local projection method to be consistent, the changes in VAT rates need to be exogenous, uncorrelated with future VAT changes, and expected to be permanent. Throughout the text, we present evidence that this is indeed the case.

gains from trade.[12] It is also important in New-Keynesian models with sticky prices, in which it influences the size of the distortion in relative prices due to nominal rigidities. This is the distortion that monetary policy (partially) offsets in these models.[13]

The bottom line is that, depending on the level of aggregation at which the upper-level expenditures are defined, a reasonable choice for the upper-level elasticity of substitution is between 1 and 3. For any choice higher than 5 there is little support in the data.

2 Model

We consider a partial equilibrium model of price setting where demand functions are determined by 2-tier nested CES preferences and production is done using a Cobb-Douglas technology. This setup nests both price setting under monopolistic competition with sticky prices, as discussed in Woodford (2003), as well as (in the limit) price setting under perfect competition, as for example in Long and Plosser (1983). We follow Karadi and Reiff (2007) and add a VAT rate that affects the firms' price setting decisions.

Given this setup, we take the following approach. We first derive the price setting decisions of firms and solve for the goods' relative price that will prevail in steady state. We then show that, for each good, the response of its relative price with respect to its VAT rate only depends on two parameters. The first is the curvature of the production function with respect to flexible inputs, which in our model is pinned down by the labor share. The second is the elasticity of substitution between goods, which is our parameter of interest.

2.1 Price setting with nested CES preferences and Cobb-Douglas technology

The economy that we consider is one in which consumers derive utility from the consumption of different types of goods, indexed by $j = 1 \dots J$. A continuum (of measure one) of varieties of each of these goods is supplied. Each variety $i \in (0, 1)$ is highly substitutable for the others within the goods category. These varieties constitute the lower level of the nested CES preferences that we analyze. The goods represent the higher level of these preferences. Our parameter of interest is the elasticity of substitution between goods at this higher level of CES preferences.

[12] See Costinot and Rodríguez-Clare (2014) for a discussion of the effect of the upper-level elasticity of substitution on the gains from trade.

[13] See, for example, Blanchard and Galí (2007).

2.1.1 Price setting decisions by producer of variety i

Each firm produces a variety i of good j in period t, Y_{ijt}, using a decreasing returns to scale Cobb-Douglas production technology, in which labor, L_{ijt}, is the sole input. That is, at a given total factor productivity level, A_t, output of the firm equals

$$Y_{ijt} = A_t L_{ijt}^{1-\alpha}. \tag{1}$$

This firm hires labor at the nominal wage rate, W_t, which it takes as given. Consequently, the marginal cost of producing an extra unit of output for the firm is

$$MC_{ijt} = \frac{1}{1-\alpha} W_t \left(Y_{ijt}^{\frac{\alpha}{1-\alpha}} \Big/ A_t^{\frac{1}{1-\alpha}} \right) \tag{2}$$

Since each of the varieties are close, but not necessarily perfect substitutes, the firm is a monopolistic competitor. This means that this firm is not a price taker but, instead, chooses a point on its variety-specific demand curve.

For our representation of this demand curve, we denote the retail price that the firm charges for its variety, including the VAT, by P_{ijt} and the price of good j across all varieties by P_{jt}.[14] The aggregate price and demand levels are P_t and Y_t respectively. The resulting demand curve for variety i is the one implied by the nested CES preferences and equals:

$$Y_{ijt} = \left(\frac{P_{ijt}}{P_{jt}} \right)^{-\eta_j} Y_{jt}, \text{ where } Y_{jt} = \left(\frac{P_{jt}}{P_t} \right)^{-\varepsilon} Y_t, \tag{3}$$

and where

$$P_t = \left[\sum_{j=1}^{J} P_{jt}^{1-\varepsilon} \right]^{\frac{1}{1-\varepsilon}} \text{ and } P_{jt} = \left[\int_0^1 P_{ijt}^{1-\eta_j} di \right]^{\frac{1}{1-\eta_j}}. \tag{4}$$

Here, $\eta_j > 1$ is the elasticity of substitution between varieties for good j and ε is the elasticity of substitution between goods.[15] The latter is the parameter we aim to estimate. Since the firm is of negligible size in terms of the supply of varieties of good j, its choice of P_{ijt} does not affect the price of good j (Dixit and Stiglitz, 1983).

The existence of a Value Added Tax means that the firm does not receive all the revenue

[14] In the European countries in our sample, prices are quoted including VAT charges, as opposed to the United States, where most price quotes are excluding sales tax. Therefore, P_{ijt} is the price the consumer pays for the variety, and thus, the price that determines the consumer's demand for the variety. Hence, P_{ijt} is the price that affects the household's cost of living. This is why prices in consumer price indices include VAT. In terms of such indices, P_{ijt} is referred to as the "purchaser price."(Eurostat, 2009)

[15] Throughout this derivation we abstract from different expenditure weights across goods j. We do so to simplify notation and the main reduced form equation we derive does not depend on this assumption.

generated at the price P_{ijt}. Instead, the VAT involves charging a tax, τ_j, as a fraction of the pretax price.[16] In terms of our notation, this means that, after the payment of the VAT, the firm receives $P_{ijt}/\left(1+\tau_j\right)$ in net revenue per unit sold. Consequently, the firm's per period flow profits are given by:

$$\frac{P_{ijt}}{1+\tau_j}Y_{ijt} - W_t L_{ijt}. \tag{5}$$

To show that our results do not depend on whether or not one assumes price stickiness, we solve the firm's price setting decision under sticky prices, in a similar way to Calvo (1983). We then show that the price-stickiness parameters do not affect the relevant reduced-form elasticity of a good's price with respect to its VAT rate.

We assume that in each period with probability ϕ_j the firm can adjust its price costlessly, while with probability $1-\phi_j$ it faces an infinite adjustment cost and will keeps its price P_{ijt} fixed. The flexible price case is simply nested in this model. It corresponds to the case where $\phi_j = 1$.

The solution to this problem yields that the fraction of firms, ϕ_j, that reset their price will all set the same price. This reset price, P^*_{jt}, is the product of three components. The first, and most important one, is that the reset price is proportional to a weighted average of the future marginal costs the firms face over the horizon that they have not adjusted their price and are still charging the reset price they currently choose. The proportionality factor is made up of the other two components. The first is the gross markup factor $\left(\frac{\eta_j}{\eta_j-1}\right)$. The second is the gross VAT rate, $(1+\tau_j)$. That is,

$$P^*_{jt} = \left[(1+\tau_j)\left(\frac{\eta_j}{\eta_j-1}\right)\sum_{s=0}^{\infty}\omega_{jt,s} MC^*_{jt+s}\right], \tag{6}$$

where MC^*_{jt+s} is the marginal cost at time $t+s$, given in equation (2), evaluated at the reset price P^*_{jt} and the weight, $\omega_{ijt,s}$, is given by

$$\omega_{jt,s} = \left[(1-\phi_j)^s\left(\prod_{j=0}^{s}\frac{1}{1+r_{t+j}}\right)P^{\eta_j}_{jt+s}Y_{jt+s}\Bigg/\sum_{q=0}^{\infty}(1-\phi_j)^q\left(\prod_{j=0}^{q}\frac{1}{1+r_{t+j}}\right)P^{\eta_j}_{jt+q}Y_{jt+q}\right]. \tag{7}$$

Because MC^*_{jt+s} is itself a function of the reset price, P^*_{jt}, equation (6) needs to be solved

[16] Because we focus on changes in steady-state relative prices, throughout we assume that the VAT rates are constant at their steady-state values. Results with time-varying VAT rates are algebraically more cumbersome but yield the same elasticity as we derive here.

for P^*_{jt} to obtain the reset price. Doing so yields:

$$P^*_{jt} = \left[(1+\tau_j) \left(\frac{\eta_j}{\eta_j - 1} \right) \left(\frac{1}{1-\alpha} \right) \sum_{s=0}^{\infty} \omega_{jt,s} W_{t+s} P_{jt+s}^{\frac{\alpha}{1-\alpha}\eta_j} \left(\frac{Y_{jt+s}^{\frac{\alpha}{1-\alpha}}}{A_{t+s}^{\frac{1}{1-\alpha}}} \right)^{\frac{1}{1-\alpha}} \right]^{\frac{1}{1+\frac{\alpha}{1-\alpha}\eta_j}}. \tag{8}$$

Based on this result, it is tempting to conclude that because for all producers of varieties of good j it is the case that

$$\frac{\partial \ln P^*_{ijt}}{\partial \ln(1+\tau_j)} = \frac{1}{1+\frac{\alpha}{1-\alpha}\eta_j}, \tag{9}$$

the elasticity of the price of good j, P_{jt}, after all firms adjust their price with respect to the value added tax rate is equal to the right-hand-side of the above equation. However, this ignores the fact that, everything else equal, consumers will substitute away from goods whose value added taxes increase more than others. Therefore, in order to fully understand the effect of the VAT change on the price of a good, we have to solve for this substitution in demand. We do so below, under the assumption that the economy is in steady state, or rather on a balanced growth path.

2.2 Relative prices in steady state

The balanced growth path is characterized by the following four properties: (i) The aggregate output Y_t grows at a constant rate, g, which is equal to the steady-state level of productivity growth $\frac{A_{t+1}}{A_t} = (1+g)$. (ii) Inflation is constant, such that the aggregate price level, P_t, as well as the prices of each of the goods, P_{jt}, grow at rate π.[17] (iii) The real interest rate, r_t, is constant and equal to r. (iv) Nominal wage growth is constant and equal to productivity growth plus inflation, i.e., $\frac{W_{t+1}}{W_t} = (1+g)(1+\pi)$.

On this balanced growth path, the forward-looking components of the reset price, P^*_{jt} defined in equation (8), can be solved to obtain:

$$P^*_{jt} = \left[(1+\tau_j)\, s_j W_t P_{jt}^{\frac{\alpha}{1-\alpha}\eta_j} \left(\frac{Y_{jt}^{\frac{\alpha}{1-\alpha}}}{A_t^{\frac{1}{1-\alpha}}} \right) \right]^{\frac{1}{1+\frac{\alpha}{1-\alpha}\eta_j}}, \tag{10}$$

[17] If there are trends in relative prices, then there is neither a balanced growth path nor steady state when $\varepsilon \neq 1$. The lack of a steady state and balanced growth path in this case is the main topic of studies of long-run structural transformation (see Herrendorf *et al.*, 2014, for example). Because we are interested in shorter horizons, we abstract from such trends in relative prices in our derivations, which assures the existence of a balanced growth path. We do allow for such trends in our empirical analysis, however.

where the goods-specific constant, s_j, equals:[18]

$$s_j = \frac{\eta_j}{(\eta_j - 1)} \frac{1}{1-\alpha} \left(\frac{1 - \left(\frac{1-\phi_j}{1+r}\right)(1+\pi)^{\eta_j}(1+g)}{1 - \left(\frac{1-\phi_j}{1+r}\right)(1+\pi)^{1+\frac{\eta_j}{1-\alpha}}(1+g)} \right). \tag{11}$$

Given the Calvo-type price setting, the law of motion of the price level of good j, P_{jt}, as a function of the reset price, P^*_{jt}, and the previous period's price, P_{jt}, is

$$P_{jt} = \left[(1-\phi_j)(P_{jt-1})^{1-\eta_j} + \phi_j \left(P^*_{jt}\right)^{1-\eta_j}\right]^{\frac{1}{1-\eta_j}}. \tag{12}$$

On the balanced growth path, good j's inflation rate equals π. In combination with the law of motion of the price level above, this allows us to solve for the level of the reset price set by the producers of varieties of good j that change their price, P^*_{jt} relative to the overall price level, P_{jt}. This yields:

$$P_{jt} = P^*_{jt} F_j, \text{ where } F_j = \left[\frac{1-(1-\phi_j)(1+\pi)^{\eta_j - 1}}{\phi_j}\right]^{\frac{1}{1-\eta_j}}. \tag{13}$$

Combining this with the solution of the reset price from equation (10), we obtain that the relative price of good j along the balanced growth path, $\frac{P_{jt}}{P_t}$, is given by:

$$\frac{P_{jt}}{P_t} = (1+\tau_j) F_j^{1+\frac{\alpha}{1-\alpha}\eta_j} s_j \frac{W_t}{P_t} \left(\frac{Y_{jt}^{\frac{\alpha}{1-\alpha}}}{A_t^{\frac{1}{1-\alpha}}}\right). \tag{14}$$

The final step in our derivation of the main equation for the steady-state relative price level is to substitute in the demand function for good j, from equation (3), to take into account that shifts in the relative price, $\frac{P_{jt}}{P_t}$, affect the level of demand for good j, Y_{jt}. Doing so, yields that, in steady state, the relative price of good j equals:

$$\frac{P_{jt}}{P_t} = (1+\tau_j)^{\frac{1}{1+\frac{\alpha}{1-\alpha}\varepsilon}} F_j^{\frac{1+\frac{\alpha}{1-\alpha}\eta_j}{1+\frac{\alpha}{1-\alpha}\varepsilon}} s_j^{\frac{1}{1+\frac{\alpha}{1-\alpha}\varepsilon}} \left[\frac{W_t}{P_t}\left(\frac{Y_t^{\frac{\alpha}{1-\alpha}}}{A_t^{\frac{1}{1-\alpha}}}\right)\right]^{\frac{1}{1+\frac{\alpha}{1-\alpha}\varepsilon}}. \tag{15}$$

Therefore, conditional on the real wage, $\frac{W_t}{P_t}$, the aggregate productivity level, A_t, and the level of output, Y_t, the elasticity of the relative price level of good j with respect to the VAT

[18] Our solution is derived under the assumption that $\left(\frac{1-\phi_j}{1+r}\right)(1+\pi)^{1+\frac{\eta_j}{1-\alpha}}(1+g) < 1$, which is true for common calibrations of r, π, g, η_j, and the price-stickiness parameter, ϕ_j.

rate, τ_j, which we denote by β, is equal to

$$\beta = \frac{1}{1 + \frac{\alpha}{1-\alpha}\varepsilon}. \tag{16}$$

Of course, this does not take into account that a VAT change potentially also has an effect on the overall economy, and thus on the real wage and output. However, the effect of the changes in these aggregate variables are the same across all goods j, which is what we exploit in our construction of the reduced-form equation that is at the heart of our empirical analysis.

The relationship between β and ε implied by equation (16) hinges on the assumption that firms face a decreasing returns to scale production technology, i.e. $\alpha \in (0, 1)$. If $\alpha = 0$ then the production function, given in equation (1), has constant returns to scale and marginal costs do not vary by the level of output of the firm. Consequently, in this case, equation (6) implies that changes in VAT rates are fully passed through in prices and $\beta = 1$. If firms face decreasing returns to scale, their level of marginal costs depends on output. Hence, a change in the relative VAT rate of a good results in a relative price change that affects relative demand, and because of the decreasing returns to scale with respect to the flexible factors, a change in the marginal cost of production. The equilibrium outcome is the fixed point in which the relative price change is consistent with the change in the marginal cost. This results in $\beta < 1$, which turns out to be what we find in the data.

In addition, the CES preferences and Dixit and Stiglitz (1983) assumption about monopolistic competition imply that our model is derived under the assumption of constant gross markups over marginal costs (equation (6)). Long-run movements in gross markups in response to change in relative VAT rates, for example because of entry and exit as in Jaimovich (2007), would affect the elasticity β. For example, if gross markups permanently decline in response to an increase in the relative VAT rate of a good, then this would bias our estimate of β down and our estimate of ε upwards.

2.3 Reduced-form equation

To construct the reduced-form equation implied by the expression for the steady-state levels of the relative prices, (equation (15)), we define the log of the price level of good j and the average log price across all goods as

$$p_{jt} = \ln P_{jt} \text{ and } \bar{p}_t = \frac{1}{n}\sum_{j=1}^{n} p_{jt}, \tag{17}$$

respectively. Using this notation, and the approximation $\ln(1+\tau_{jt}) \approx \tau_{jt}$, we can rewrite equation (15) to obtain:

$$p_{jt} \approx \beta\left(\tau_{jt} - \bar{\tau}_t\right) + \delta_j + \xi_t + \bar{p}_t, \tag{18}$$

where

$$\bar{\tau}_t = \frac{1}{n}\sum_{j=1}^{n} \tau_{jt}, \tag{19}$$

and

$$\delta_j = \beta\left[\left(1 + \frac{\alpha}{1-\alpha}\eta_j\right)\left(\ln F_j - \frac{1}{n}\sum_{k=1}^{n}\ln F_k\right) + \left(\ln s_j - \frac{1}{n}\sum_{k=1}^{n}\ln s_k\right)\right]. \tag{20}$$

In practice, however, we do not have data for the log of the price *levels*, p_{jt}. So, empirically implementing the above as a reduced form equation is not feasible. Instead, we have data on log price *indices*, changes in which are constructed to be proportional to changes in the log of the price levels.

Hence, to operationalize equation (18) as a reduced-form equation, we focus on the *change* in the log of the steady-state price levels, Δp_{jt} in response to a *change* in the VAT rate, $\Delta\tau_{jt}$ compared to the change in the average VAT across goods, $\Delta\bar{\tau}_t$. That is, the reduced-form equation that forms the basis of our empirical analysis is:

$$\Delta p_{jt} \approx \beta\left(\Delta\tau_{jt} - \Delta\bar{\tau}_t\right) + \Delta\xi_t + \Delta\bar{p}_t. \tag{21}$$

Note that in this specific equation, the reduced form parameter, β, represents the elasticity of the response of relative prices to changes in relative VAT rates, and it only depends on two parameters: (i) the output elasticity of labor, $(1-\alpha)$, and (ii) the between-goods elasticity of substitution, ε. The latter is the parameter we aim to estimate.

More important is the list of other parameters that it does *not* depend on. First of all, because we focus on steady-state levels of relative prices, the elasticity, β, does not depend on the frequency of price adjustment. It is the same, no matter whether prices are flexible (i.e. $\phi_j \to 1$) or sticky.[19] Neither heterogeneity in the degree of price stickiness, by assuming ϕ_j is different across goods (as in Carvalho, 2006), nor in markups, by assuming η_j varies across goods (as in Eusepi *et al.*, 2011), affect the elasticity of relative prices with respect to VAT rate changes, β. One potential source of heterogeneity across goods that would affect β is heterogeneity in the output elasticity of labor, $(1-\alpha)$. Such heterogeneity would result in

[19] We have derived our results under Calvo-style nominal rigidities. Our steady-state results are also valid under state-dependent pricing. See Klenow and Kryvtsov (2008) for a detailed comparison of models under these two types of price setting.

different labor shares in the production of different consumption goods. In fact, Fisher (1969) shows that such heterogeneity would prevent us from finding a simple closed-form solution. Moreover, Eusepi *et al.* (2011) show that, in the U.S., labor shares do not vary much across the production of consumption goods at the level of aggregation that we consider in this paper. Therefore, we abstract from this source of cross-good heterogeneity.[20]

2.4 Why not a general equilibrium analysis?

As discussed above, the reduced-form parameter, β in equation (21), is the same under a broad set of underlying assumptions. As a result, an analysis based on the estimation of (21) should be robust to a large set of assumptions.

An alternative approach would be to estimate the parameters of the model, including ε, using a dynamic stochastic general equilibrium model. This is what Karadi and Reiff (2007) do. Such an approach would allow for the estimation of all the parameters, and not only ε, underlying the general equilibrium structure of the model.

Though such an approach allows one to focus on a broad set of parameters, it does require one to make specific assumptions about the sources of heterogeneity that our reduced-form parameter, β, does not depend on. Moreover, to close the model one also has to make specific assumptions about household preferences. In particular, about the intertemporal elasticity of substitution and Frisch elasticity of the labor supply. In order to fit the path of aggregate inflation, one also has to add a monetary policy rule.

Our approach, which identifies the between-goods elasticity of substitution, ε, from the correlation between long-run changes in relative prices and changes in relative VAT rates, is valid for any type of aggregate household preferences and monetary policy rule. It is derived solely based on assumptions about the household's intratemporal utility maximization problem, which is driven by nested CES preferences, and firms' price-setting decisions when using a Cobb-Douglas production technology where labor flows freely between producers of different varieties and goods.

3 Empirical implementation

Exploiting the insight about the relationship between the upper-level elasticity of substitution and the long-run response of relative prices to relative VAT changes in practice requires map-

[20]In addition, with identical growth rates of total factor productivity, g, a balanced growth path does not exist when the output elasticity of labor, α, varies across goods and $\varepsilon \neq 1$.

ping equation (21) into existing data. In this section we describe how equation (21) can be estimated using a relatively simple panel data regression that implements a local projection method (Jordà, 2005). This allows us to estimate the long-run response of relative prices to VAT changes.

3.1 Data

In principle, the elasticity, β, in equation (21) could be estimated solely based on cross-good variation in changes in relative prices. However, because changes in VAT rates in a country are relatively infrequent, it is useful to pool the regression across countries.

Thus, our panel data analysis uses three sources of variation for the estimation of β; goods, j, countries, c, and time, t. Prices, p_{jct}, are measured using the logarithm of the monthly Harmonised Index of Consumer Prices (HICP), and used to calculate inflation at the good-specific level for 25 European Union (E.U.) countries in our sample. Our data cover the time period January 1996 through January 2015.[21]

Expenditures included in HICPs are classified in categories/goods, j, called COICOPs.[22] The COICOP classification system consists of three levels of aggregation. In our sample, the highest level of aggregation (one-digit level) consists of 12 *divisions*, such as food and non-alcoholic beverages, communication, restaurants and hotels, etc.[23] The next level of aggregation (two-digit level) is called a *group*. As an example, accommodation services is a group within the division of restaurants and hotels. Our sample includes 36 groups. The lowest level of aggregation (three-digit level) is a *class*. The group of alcoholic beverages consists of classes that cover spirits, wine, and beer, separately. Our sample includes 74 classes.

The level of granularity at which goods and services are defined matters for their substitutability. Hence, we report our estimate of β for each of these different levels of aggregation at which we have data.

Data on VAT rates by COICOP for the 25 countries in our sample are not readily avail-

[21]The Appendix Table A1 provides a complete list of countries in our sample. In particular, our sample includes class-level HICP price indices and VAT rates for Austria, Belgium, Cyprus, Czech Republic, Denmark, Germany, Estonia, Greece, Spain, Finland, France, Hungary, Ireland, Italy, Lithuania, Latvia, Luxembourg, Malta, Netherlands, Poland, Portugal, Slovakia, Slovenia, Sweden, and the United Kingdom. Austria, Denmark and Sweden have no documented VAT rate changes between 1996 and 2015. In addition, our data originally included Bulgaria and Romania, which we dropped because those countries faced periods of hyperinflation.

[22]COICOP is an acronym for "Classification of Individual Consumption According to Purpose."

[23]The complete list of COICOPs is provided in the Appendix Tables A2 to A4. Our data include 12 divisions, from which we dropped all classes (and groups) pertaining divisions 6 (Health) and 10 (Education) because of the non-market nature of price setting in these sectors. In addition, our data do not include two additional divisions that cover spending by non-profit institutions serving households (NPISH) and government consumption for which price data is imputed rather than directly measured.

able. We construct them from two administrative sources: European Commission (2015), and Eurostat (2015). These give us information about which VAT rates are applicable to which goods and services in a country over time.

VAT rates are not the same for all goods within a country. Most countries have four different VAT rates that apply to different goods: super-reduced, reduced, standard, and parking rate. In addition, many countries have goods and services, such as education, for example, that are exempt from value added taxes.

E.U. law requires that the standard VAT rate is at least 15% and the reduced rate at least 5%. Actual rates applied vary across countries.[24] For example, in 2015 goods facing a standard rate in the United Kingdom were charged a 20% VAT rate, while in Luxembourg the standard rate was 15%. In addition, the four broad rate categories can also include a range of levels of VAT rates, and the level applied to a certain category may change over time.[25] Finally, the same categories may face different VAT rates in different countries. The result is that VAT rates for the same goods and services vary across countries and over time. For example, as of January 2015, Spain charges a 10% VAT rate on hotel accommodation services while Portugal charges a 6% rate. Restaurants face a 7% VAT rate in France and a 19% rate in Germany.

The most important variation in VAT rates for the estimation of β, however, is changes in VAT rates on specific goods and services over time. Most of these changes are because the rate associated with the VAT category that a good or service is classified in changes. For example, between June 2010 and September 2012, Spain increased its standard VAT rate from 16% to 21%. The majority of these VAT rate changes in our sample occur either on January 1st or July 1st.

Sometimes a good or service gets reclassified into a different VAT rate category. For example, before September 2012 cosmetic surgery in Spain was charged a reduced VAT rate while afterward it fell under the standard VAT rate category.

Our matching of the administrative data on VAT rates with COICOPs yields good-country specific time series for the applicable VAT rate, τ_{jct}, where j is defined, alternatively, at the COICOP class, group, or division levels. VAT rates for the group- and division-level expenditure categories are constructed as weighted averages of the VAT rates in the underlying classes.[26]

[24] These simple rules are, however, complicated by a multitude of derogations granted to certain European Union Member States, and in some instances, to a majority of Member States.

[25] For example, in 2010, France applied two different super reduced VAT rates, 2.1% and 5.5%. In 2012, a third 7% super reduced rate was introduced.

[26] Due to lack of data on consumer expenditures at the class level, to aggregate from the class up to the group level, we use an equally weighted average of VAT rates across classes within each group. We use consumer expenditure shares to aggregate from group up to division level. The construction of our data is described in

The HICP price data, p_{jct}, and VAT data, τ_{jct}, are the left- and right-hand side variables of our reduced-form equation (21). What is left is to map equation (21) into a specification that allows for the identification and estimation of β for the country-COICOP-time panel structure of our data.

3.2 Model specification and identification

Equation (21) describes log changes in relative price levels, for good j, between two steady states that differ in the VAT rates charged. In practice, of course, log changes in observed prices reflect more than only shifts between steady states. Throughout, we interpret the steady-state response of relative prices as *long-run* movements in the data.

In particular, we estimate β as the effect of cumulative VAT rate changes from t to $t+l$, which we denote by $\Delta_l \tau_{jct+l} = \tau_{jct+l} - \tau_{jct}$, on the cumulative log change of prices from t to $t+h$, where $h \gg l$. We denote this cumulative log change in prices by $\Delta_h p_{jct+h} = \ln p_{jct+h} - \ln p_{jct}$.

For each choice of the length of period over which we accumulate VAT changes, l, and the horizon over which we consider the long-run response of log prices, h, we obtain an estimate $\beta_{l,h}$. This parameter is estimated using the local projection method introduced in Jordà (2005).

For the particular estimation problem at hand, this amounts to running a panel data regression of the form:

$$\Delta_h p_{jct+h} = \beta_{l,h} \Delta_l \tau_{jct+l} + \gamma_{ct} + \alpha_{jcm} + u_{jct}. \tag{22}$$

In addition to the changes in the log prices and VAT rates, this equation contains a country-time fixed effect, γ_{ct}, and a COICOP-country-month fixed effect, α_{jcm}. u_{jct} is the residual.

The country-time fixed effect absorbs both the effect of the average change in VAT rates, $\Delta\bar{\tau}$, the average log change in prices, $\Delta\bar{p}_t$, as well as the change in country-wide economic conditions, $\Delta\xi_t$, from equation (21). Including this fixed effect means that we do not have to specify our regression in terms of the logs of relative prices and relative VAT changes. The country-time fixed effect captures the country-specific changes in the log of the overall price level and the average VAT rate change across goods in a country. We also include a country-COICOP-month fixed effect, α_{jcm}, to allow for potential trends in relative prices across goods and countries, as well as seasonal effects in country-good-specific inflation rates.

Figure 1 helps illustrate our identification strategy. We estimate the effect of tax changes between t and $t+l$, i.e. $(\tau_{t+l} - \tau_t)$, on the cumulative log price change between $t+h$ and t,

more detail in Appendix A.

$(p_{t+h} - p_t)$, where $h \gg l$.

For our estimate of $\beta_{l,h}$ to be consistent, the VAT changes, $\Delta_l \tau_{jct+l}$, need to have three properties. First, they need to be uncorrelated with the grayed out future tax changes, between $t+l$ and $t+h$, in Figure 1. If this is not the case, then our estimate of $\beta_{l,h}$ suffers from omitted variable bias, since it will partly include the effect of tax changes between $t+l$ and $t+h$, rather than between t and $t+l$, on the log change of the price level. Secondly, since we focus on the long-run response of prices to the VAT changes, the VAT changes, $\Delta_l \tau_{jct+l}$, need to be (expected to be) permanent. If this is not the case then the long-run response to the VAT changes we identify will be muted in the data compared to our model. Finally, they need to be exogenous with respect to the residual u_{jct}. We revisit these three properties in Subsection 4.1, where we provide evidence in support of them.

In our derivation of equation (21) we use labor as the only adjustable factor of production. The parameter α reflects the degree of decreasing returns to scale of the production function with respect to this factor. Because labor is mobile across firms, all firms pay the same real wage and face the same marginal cost schedule.

Our empirical approach in equation (22) allows for marginal cost schedules to vary across goods. As long as changes in these schedules are uncorrelated with VAT changes, $\Delta_l \tau_{jct+l}$, our estimate of $\beta_{l,h}$ remains consistent. In that case, these orthogonal changes in marginal costs generate unexplained changes in log prices, $\Delta_h p_{jct+h}$, that are either absorbed by the fixed effect, α_{jcm}, or are part of the residual, u_{jct}.

What remains, for the practical implementation of equation (22), is to choose l, and the value of h that we interpret as the "long-run." As for the value of l, we go with the natural choice of $l = 12$. That means we consider the long-run effect of 12-month changes in VAT rates, $\Delta_{12}\tau_{t+12}$, on log price levels.

What exactly we mean by "long-run" is less clear-cut. The obvious choice seems to associate long run with h as large as possible. In practice, however, increasing h reduces the effective sample size and thus the degrees of freedom and the precision of the parameter estimate we are interested in. Moreover, choosing h too large for the estimation of equation (22) poses a theoretical challenge to our approach.

This challenge is that the production function in equation (1) has decreasing returns to scale in the adjustable production factor, which is labor in this case. Therefore, implicitly, it assumes that capital inputs are fixed. The decreasing returns to scale with respect to the adjustable inputs are captured by the term $\frac{\alpha}{1-\alpha}$ in the expression for the reduced form parameter β in equation (16). Therefore, in the longer run, the appropriate standard production function would also include capital as an adjustable factor, making returns to scale constant. However, solving the model under constant returns to scale to adjustable inputs results in $\beta = 1$, which case it does not depend on ε. As we show later, however, $\beta < 1$ in the data.

Though it might seem contradictory to estimate "long-run" effects assuming the capital inputs are fixed, the relevant "long-run" for our purpose is the duration of the transitional price setting dynamics in response to a change in VAT rates. Such transitional dynamics tend to die out after about 40 months in common multisector New Keynesian models.[27] In the short-run, the transitional price-setting dynamics in these models are affected by the degree of price stickiness across goods and services, which is something that the steady-state effects that we aim to estimate does not depend on. Thus, choosing h too small would render our estimates inconsistent. With this in mind, we focus on $h = 48$ in our baseline set of results and discuss how they change when we vary h between 12 to 62.

4 Empirical results

We present our empirical results in four parts. First, we document the amount of variation in VAT rates across countries and COICOPs. This is the variation in the right-hand side of equation (22), i.e. in $\Delta\tau_{jct}$, that we use to identify the parameter $\beta_{l,h}$. Next, we present the results for our baseline specification, i.e. equation (22), for the lowest level of aggregation of COICOPs, where goods and services, j, are defined at the class level. Third, we show that these results are robust to various different model-specifications and sample choices. Finally, we present the results for the two other, higher, levels of aggregation, namely at the group and division levels.

4.1 Changes in VAT rates

Of course, as with any regression, ours, based on equation (22), needs substantial variation in the explanatory variable of interest to reliably estimate the associated coefficient. So, before we present our estimation results, we consider the variation in changes in VAT-rates in our sample. In addition, we provide evidence to show that these changes are (*i*) uncorrelated with future changes in VAT rates, (*ii*) are expected to be permanent, and (*iii*) the variation in these changes that use for identification of β is exogenous with respect a wide range of factors affecting long-run price changes.

[27]See Bouakez *et al.* (2009), Carvalho (2006), and Carvalho and Nechio (2014) for a detailed analysis of such transitional dynamics.

Variation in VAT changes

Figure 2 provides two measures of the variation in VAT rate changes across countries and COICOPs at the class level of aggregation. The first is the monthly count of the number of non-zero 12-month changes in VAT rates across COICOPs and countries. This is depicted by the dark-shaded area in the figure. Changes in VAT rates have occurred during the whole sample period, although they are concentrated in the post-2008 part of the sample. This means that the bulk of the variation that identifies our parameter of interest, $\beta_{l,h}$, is in the last seven years of the sample.

Note, however, that while the right-hand-side variable of equation (22) corresponds to a simple 12-month change in the VAT rate by country and COICOP, our estimates reflect the effects of VAT rate changes on prices after accounting for COICOP-country-month (α_{jcm}) and country-time (γ_{ct}) fixed effects. Therefore, the variation that we actually exploit in our empirical approach involves *demeaned* changes in VAT rates that result from the absorption of these fixed effects. This demeaned variable captures deviations of changes in VAT rates from the average change across COICOPS in a country.

The second measure of variation in VAT rate changes in Figure 2, depicted by the light-shaded area, is the monthly standard deviation of this demeaned variable across countries and COICOPs. This measure shows that, while the concentration of changes in VAT rates is in the latter part of the sample, the effective variation in the explanatory variable, after taking into account the fixed effects, is more evenly distributed across the sample.

Hence, our sample to estimate $\beta_{l,h}$ includes not only a large number of VAT changes, but also substantial variation in changes in VAT rates across COICOPs, countries and time.[28]

Exogeneity and expected permanence of relative changes in VAT rates

When we introduced our identification strategy in Section 3.2, we emphasized that, for us to obtain a consistent estimate of β, the VAT changes, $\Delta_l \tau_{jct+l}$, need to have three properties: (i) They need to be uncorrelated $\Delta_{h-l} \tau_{jct+h}$, (ii) they need to be (expected to be) permanent, and (iii) they need to be exogenous with respect to the residual u_{jct}.

To test for the first property, i.e. whether future changes in tax rates are uncorrelated with the changes we include as explanatory variables, we estimate the effects of $\tau_{t+h} - \tau_{t+l}$ on $\tau_{t+l} - \tau_t$ by performing simple OLS regressions for $l = 12$ and $h = 12, \ldots, 62$. We find that changes in tax rates between $t + l$ and $t + h$ are only very weakly correlated with changes in

[28] More detailed summary statistics about the incidence of and variation in VAT changes across countries and COICOPs can be found in Appendix A.

taxes between t and $t+l$, with all regressions yielding a $R^2 < 0.002$.[29]

Moreover, in line with the second property, the changes in VAT rates included in our data set are persistent. A simple OLS regression of VAT rates on their lags, controlling for the same fixed effects included in equation (22), shows that the coefficient on the first lag is near one (specifically, 0.997), and it is the only statistically significant coefficient.[30]

As for the exogeneity of the relative VAT changes in equation (22), it is important to realize that most factors with which VAT rate changes are potentially correlated are captured by the fixed effects. Inclusion of the country-time fixed effects, γ_{ct}, means that our estimates are not affected by any country-specific factors that vary over time, like country-specific business cycles, legislation, or inflationary effects. In addition, the COICOP-country-calendar-month fixed effect, α_{jcm}, filters out seasonal fluctuations by COICOP and country. This leaves only the case in which the incidence of legislated relative VAT changes across goods is correlated with factors that affect long-run relative price changes across goods beyond the VAT rate changes themselves.

4.2 Results at class level

The dark line in Figure 3 depicts the estimates $\hat{\beta}_{l,h}$ for $l = 12$ and $h = 12, \ldots, 62$. The shaded area is the associated 95 percent confidence interval. The estimates of β initially decline as a function of h, from 0.30 at $h = 12$ to 0.24 at $h = 25$. After that, $\hat{\beta}_{l,h}$ steadily increases and peaks at 0.43 at $h = 52$.

At our choice for the long run, $h = 48$, the estimated effect of changes in VAT rates on long-term changes in relative prices equals 0.40. Column (I) of Table 1 provides the detailed regression results for this baseline specification. The estimate of β is relatively precise with a standard error of 0.04.

To map the estimated coefficient $\hat{\beta}_{l,h}$ into the implied elasticity of substitution, $\hat{\varepsilon}$, using equation (16), one needs to take a stand on the value of α. The value of α is commonly picked based on evidence on the average labor share, which equals $(1-\alpha)$ in the class of models we consider. European data on GDP and labor compensation suggest that, despite some variability across countries, the average labor share for the European Union is approximately 2/3 (European Commission, 2007, Chapter 5). Our model is not about the overall economy, however, but rather about different goods and services that make up consumer spending. U.S. evidence (Eusepi *et al.*, 2011, Table 1) suggests that there is little variation in the labor

[29]This also holds when we consider other levels of aggregation.

[30]In particular we consider the regression: $\tau_{jct} = \sum_{k=1}^{24} \lambda_k \tau_{jct-k} + \gamma_{ct} + \alpha_{jcm} + \varepsilon_{jct}$. Estimates of this equation by country yield qualitatively similar results.

share across consumer goods. Because of this, we calculate our results for the case where $(1-\alpha) = 2/3$. In Section 4.3 we discuss the sensitivity of our results to our choice of α.

As can be seen from the bottom two rows of Column (I) of Table 1, the elasticity of substitution, $\hat{\varepsilon}$, implied by $(1-\alpha) = 2/3$ and our parameter estimate $\hat{\beta}_{l,h} = 0.40$, is 2.98. Taking into account the standard error of the reduced form parameter, this suggests that, with a 95% probability, the implied elasticity of substitution is between 2.1 and 4.3. Thus, our evidence at the class level of aggregation of COICOPs is that goods and services bought by consumers at this low level of aggregation are substantially more substitutable than in case of Cobb-Douglas preferences. However, the elasticity of substitution is much lower than 11, which is the value of the between-varieties elasticity of substitution, η_j, that most commonly gets used.[31]

4.3 Robustness of class-level results

The value of the implied elasticity is robust with respect to many perturbations of our sample and model. To illustrate this, in this section we present different estimates of β for a shortened sample period, for a different specification that includes lagged inflation, and for different choices of h and l. In addition, we also discuss how the choice of α affects our results, why it is important to pool the model across many countries and COICOPs, and why we conclude there are little or no anticipatory price increases in our data.

Sample period and lagged inflation

Table 1 reports estimates for an alternative sample period, Column (II), and for a specification that includes lagged inflation, Column (III). The estimated reveal that the estimated reduced-form parameter, $\hat{\beta}_{12,48}$, as well as the implied elasticity of substitution, does not vary very much when we change the sample period or add lagged inflation terms to the model.

In particular, Column (II) of the table lists the results obtained by estimating equation (22) for the time period that *excludes* the Great Recession and its aftermath. The point estimate of the reduced form parameter, β, is almost the same if one excludes the VAT changes in the latter seven years of the sample. However, because, as we showed in Figure 2, most variation in VAT rates occurs in these years, this widens the confidence interval for the implied elasticity.

Column (III) of Table 1 contains the results for an extension of equation (22) that includes 12 lags of COICOP-country-specific inflation. We include these lags of inflation to capture

[31] This common calibration of η_j is based on evidence in Basu and Fernald (1997) that suggests that markups of price over marginal cost are in the order of 10 percent.

potential inflation dynamics at work at the time of the VAT rate changes that might influence our estimated β because of omitted variable bias. Thus, the generalized specification estimated reads

$$\Delta_h p_{jct+h} = \beta_{l,h} \Delta_l \tau_{jct+l} + \sum_{k=1}^{12} \delta_k \Delta p_{jct-k} + \gamma_{ct} + \alpha_{jcm} + u_{jct}, \tag{23}$$

where δ_k for $k = 1, \dots, 12$ are the added parameters.

It turns out that adding these lagged inflation terms does not substantially alter our estimates. The implied elasticity in this generalized specification is 2.3 with an associated 95% confidence interval ranging from 1.6 to 3.2. Again, this suggest that at this low level of aggregation consumer goods and services are more substitutable than Cobb-Douglas but a lot less than the commonly assumed substitutability between varieties.

Sensitivity to choice of h and l

We already discussed how theoretical results guided us to choose $h = 48$ as our benchmark and showed, in Figure 3, that, over the range $h = 36, \dots, 60$ our estimate $\hat{\beta}_{12,h}$ is very similar.

The blue lines in the same figure provide more evidence in support of our choice of $h = 48$. The estimation of equation (22) focuses on the effects changes of in VAT rates accumulated over a 12-month period. Alternatively, one could slice this 12-month change up into 12 one-month changes. This results in an unrestricted version of regression (22), in which, for each horizon h, we quantify the effect of all one-period changes in VAT between t and $t + 12$. The resulting regression specification is

$$\Delta_h p_{jct+h} = \sum_{m=1}^{12} \beta_{12,h,m} \Delta \tau_{jct+m} + \gamma_{ct} + \alpha_{jcm} + u_{jct}, \tag{24}$$

where $\Delta \tau_{jct+m} = \tau_{jct+m} - \tau_{jct+m-1}$.

Our theory suggests that β is the same no matter what the choice of the parameter l over which we measure VAT rate changes, as long as we focus on their long-run effects. So, splitting up our baseline choice of $l = 12$ into its 12 sub-periods should yield that $\beta_{l,h,m} = \beta_{l,h}$ for all $k = 1, \dots, 12$.

The blue lines in Figure 3 show the estimates $\hat{\beta}_{12,h,m}$ where $m = 1, \dots, 12$. As can be seen from the figure, there is noticeably more variation across m in the $\hat{\beta}_{12,h,m}$'s for $h < 24$ than for $24 \leq h \leq 50$. In fact, a formal F-test of $H_0 : \beta_{l,h,m} = \beta_{l,h}$ for all $k = 1, \dots, 12$ rejects this null at a 10 percent significance level for all $h < 22$. For all $h \geq 22$ this null is not rejected. However, the variation in $\hat{\beta}_{12,h,m}$'s does increase a lot after $h = 50$. At our choice of $h = 48$

the variation in the point estimates $\hat{\beta}_{12,h,m}$ across m is the smallest in our sample.

This result provides corroborating evidence in support of our choice of $h = 48$. Furthermore, because the estimated $\hat{\beta}_{12,h,m}$'s are so similar for the different m, the result also shows that our baseline result in column (I) of Table 1 is not driven by our choice of $l = 12$.

Sensitivity to the choice of α

Our mapping between the long-term estimate of $\beta_{12,48}$ and the upper-level elasticity of substitution is directly dependent on our assumed value of the labor share. Our choice for the benchmark level of the parameter α was guided by the empirical evidence on the average labor share across goods in the U.S. and across E.U. countries. However, given the estimate of β reported in Table 1, a different choice of α implies a different value of ε.

To consider how the implied elasticity is affected by α, the blue line in Figure 4 reports how the estimated ε varies as the labor share $(1 - \alpha)$ increases from 0.4 to 0.85.[32] The shaded area is the associated 95 percent confidence interval for the class level estimates. We discuss the results for the other levels of aggregation plotted in the figure in Section 4.4.

Figure 4 shows that, as the labor share increases, the point estimates for the elasticity ε increases. As the labor share increases from 66%, our benchmark level, to 85%, for example, point estimates for the upper-level elasticity of substitution increase from 3 to 8 at the class level. The larger the labor share, however, the smaller the precision of our estimates, as can be seen in the widening of the 95% confidence interval for the class level. Taking into account the uncertainty around our estimates, as for a labor share between 60% and 75%, the range of estimates for the upper-level elasticity at the class level still includes our benchmark point estimate of 3.

The virtue of pooling across countries

Thus far, we focused on results obtained from the estimation of equation (22) with data pooled across both countries and COICOPs. Of course, the expression (21), on which equation (22) is based, is only for one particular country.

The reason we pool our results across countries is that VAT rate changes are infrequent enough that there is not enough variation in the right-hand side variable to reliably estimate the reduced-form parameter, β, separately for all countries.

[32]For the class of goods in our data, U.S. evidence (Eusepi *et al.*, 2011, Table 1) suggests the labor share varying between 52% and 77%.

To illustrate this, we estimate the following country-specific regressions:

$$\Delta_h p_{jt+h}^{(c)} = \beta_{l,h}^{(c)} \Delta_l \tau_{jt+l}^{(c)} + \gamma_t^{(c)} + \alpha_{jm}^{(c)} + u_{jt}^{(c)}, \;\; \forall c, \tag{25}$$

where γ_t is a time fixed effect and α_{jm} is a COICOP-month fixed effect.

Figure 5 shows the distribution of the country-specific estimators of $\hat{\beta}_{12,48}^{(c)}$, as well as the asymptotic distribution of the pooled estimate, $\hat{\beta}_{12,48}$, reported in Table 1.[33]

The comparison between the two distributions highlights the gains from pooling across countries. First of all, many of the country-specific point estimates are negative which implies an implausible elasticity of substitution, ε, smaller than -0.5. Moreover the standard errors of the country-specific coefficients are rather large.

The pooled Ordinary Least Squares (OLS) estimate, $\hat{\beta}_{12,48}$, is a complicated weighted average of the country-specific ones where those country-specific coefficients with the smallest estimated standard errors get the highest weights.[34] The pooled parameter estimate coincides with the only mode of the density of the country-specific $\hat{\beta}_{12,48}^{(c)}$'s in positive territory. Moreover, because it is based on many more VAT changes it is estimated with a much smaller standard error than the country-specific parameters.[35]

Possible anticipatory effects

Our estimate $\hat{\beta}_{l,h}$ implies a less than one-to-one pass-through from changes in VAT rates to prices. While this finding is not at odds with some of the literature that explored the effects of changes in VAT on inflation in specific countries, it is possible that our estimates do not capture anticipatory effects of changes in VAT rates.[36] Such anticipatory effects occur when relative prices move in response to the announcement of VAT changes rather than in response to the actual VAT change that we measure in our data.

Changes in VAT rates are frequently announced in advance of their effective implementation. For example, on November 2008, the U.K. government announced a reduction in its standard VAT rate from 17.5% to 15% to become effective by December 2008 (see Seely

[33]The distribution of the $\hat{\beta}_{12,48}^{(c)}$'s depicted in the figure is a kernel density estimate using a normal kernel and setting the bandwidth for each $\hat{\beta}_{12,48}^{(c)}$ equal to its estimated standard error.

[34]See Wooldridge (2001), page 150, for a derivation of this result.

[35]Though, a formal F-test for the null $H_0 : \hat{\beta}_{12,48}^{(c)} = \hat{\beta}_{12,48}$ for all c is rejected at a 1% significance level.

[36]In particular, Andrade *et al.* (2010) use firm-level French trade data to show that the degree of pass-through of changes in VAT rates to prices can be smaller than one-to-one, as it depends on both the degree of competition on a given market and each firm's market share. Using Hungarian micro-level data, Karadi and Reiff (2014) also find that the degree of pass-through can be smaller than one-to-one depending on the size and the sign of changes in VAT shocks.

(2013)). In addition, in a case study of a preannounced tax hike in Germany, Carare and Danninger (2008) find some of the effects in inflation to occur during the announcement period. Therefore, one possible concern is that our estimates of $\beta_{l,h}$ do not account for the whole effect of VAT rate changes on prices.

Because we have no comprehensive data on announcement dates for all VAT rate changes included in our sample, we test whether prices react to changes in VAT rates before their effective implementation by estimating our main regression (equation (22)) for h ranging between 1 and 12. This captures announcement effects in the following way. Suppose that VAT rates changed in July 2005. Then, our regression for $h = 1, \ldots, 12$ considers the relative price change in August 2004 to be affected by the change in VAT rates over the subsequent 12 months that include July 2005. Hence, this regression captures anticipatory effects up to a year before the actual VAT change. Interestingly, our results suggest little evidence of anticipatory effects of changes in VAR rates on prices. In particular, the estimated $\beta_{l,h}$ for $h = 1, \ldots, 12$ roughly increases linearly up to its value at $h = 13$, before attaining the dynamics portrayed in Figure 3. This pattern is consistent with no anticipatory effects at all.

4.4 Results at higher levels of aggregation

The results above pertain to the *class* level of aggregation of COICOPs, which is the lowest level of aggregation in the data. However, the degree of substitutability of goods and services depends on the granularity with which they are defined. For this reason, we present results for the two higher levels of aggregation, the *group* and *division* classifications, in Table 2.

For reference purposes, column (I) of the table contains the benchmark results at the class level of aggregation from Table 1. Columns (II) and (III) contain the results at the group and division levels respectively. Note that the number of COICOPs decreases from 74 at the class level to 10 at the division level.[37]

The first thing to notice in this table is that our point estimate of $\beta_{12,48}$ is increasing in the level of aggregation. This implies a lower elasticity of substitution at higher levels of aggregation. In fact, our point estimate of the implied elasticity for the division level is one, which coincides with Cobb-Douglas preferences.

At higher levels of aggregation the size of the cross-section in the data is, by definition, smaller. Thus the number of observations, listed in Table 2, declines as we move from column (I) to (III). As a result, the precision with which β is estimated decreases due to a reduction in the degrees of freedom.

[37]Tables A2 through A4 contain a detailed list of all COICOPs in our dataset.

The lower precision of the estimates at the higher levels of aggregation is reflected in the wider confidence intervals for the implied elasticity ε. Note, however, that none of the intervals includes elasticities that are substantially higher than 5.

Finally, as for our mapping between $\beta_{12,48}$ and ε at the class level, estimates at higher levels of aggregation also depend on our assumption for α. The remaining two lines depicted in Figure 4, show that as the labor share increases, the point estimates for the elasticity at at the group and division levels of aggregation also increase. As the labor share increases from 2/3, our benchmark level, to 85%, for example, point estimates for the upper-level elasticity of substitution increase from 2.6 to 7.4 at the group level, and from 1 to 2.8 at the division level.[38]

The bottom line of Table 2 is that when one has to choose ε in practice, the appropriate choice depends on the level of aggregation of the consumption aggregate considered. For a high level of aggregation, Cobb-Douglas preferences are consistent with our point estimates. For lower levels of aggregation our results point to an elasticity of substitution of about 3. Taking into account the uncertainty around our estimates we obtain an upper-bound of approximately 5.

5 Implications

The empirical results above imply that, depending on the level of aggregation considered, a reasonable choice of elasticity of substitution for the upper level of nested CES preferences ranges from 1 to 5. Here we discuss two strands of the literature for which this elasticity is relevant and its value has important quantitative implications.

Throughout, it is important to realize that the larger ε, the quicker consumers substitute towards goods with the lowest relative price. In the limit, when $\varepsilon \to \infty$, the only thing that matters for utility is the lowest price charged across goods. In this case, if relative prices change, but the lowest price remains the same, nothing happens to demand.

Our first example, related to gains from trade, is a case in which relative prices are set efficiently. Our second example is that of the distortion to relative prices due nominal rigidities. These distortions result in a welfare loss that monetary policy could possibly reduce.

Gains from trade The elasticity of substitution between goods and services is of first-order importance for the magnitude of the gains from trade. As Costinot and Rodríguez-Clare (2014) point out most analyses of the gains from trade in multisector models follow Dawkins *et al.*'s

[38]Unreported 95% confidence intervals for the group and division levels show similar patterns as those at the class level, that is, as the labor share increases, precision of our estimates declines.

(2001) "idiot's law of elasticities" which mandates assuming that the upper-level elasticity of substitution is one (aggregate preferences are Cobb-Douglas), until shown to be otherwise. Our estimates, especially those in column (III) of Table 2, show that this might not be such an "idiotic" choice after all when one defines the upper-level at a high level of aggregation of goods and services.

In multisector models, the magnitude of the gains from trade depend a lot on the value of the upper-level elasticity of substitution, ε. This is because trade results in relative price changes of goods. Costinot and Rodríguez-Clare (2014) show that, in a relatively standard trade model, gains from trade decline in ε. This is because, when there is very little substitution across goods (small ε), autarky implies the absence of certain goods, which, by assumption, cannot easily be substituted. Consequentially, the immediate gain from opening to trade, and the availability of a wider set of goods, are larger the smaller is ε. In particular, Costinot and Rodríguez-Clare (2014) show that gains from opening to trade for $\varepsilon = 1$ tend to be more than double those at $\varepsilon = 5$. However, when processes of globalization continuously lower relative prices of tradable goods and services, then this process in the long-run will yield higher welfare gains in the case in which consumers can substitute towards the goods that globalization renders cheaper over time. In that case the benefits from globalization are increasing in the upper-level of substitution, ε. Moreover, if globalization results in gains from varieties, as in Broda and Weinstein (2006), then those gains from varieties effectively result in relative price declines of goods where these gains are largest. Again, in such a case, the benefits of such gains are higher in case ε is higher.

Distortions from price rigidities Another example where the value of the upper-level elasticity of substitution between expenditure categories matters is in models with nominal rigidities due to price stickiness.[39] In those models, the inability of producers to adjust their prices distorts relative prices across goods. Such distortions occur because of sector-specific shocks or due to different degrees of price rigidities across goods. This distortion is the one that monetary policy can, partly or completely depending on the model, offset (Blanchard and Galí, 2007).

The higher the upper-level of substitution, the more demand shifts to the goods and services with distorted low relative prices. Thus, the higher is the upper-level elasticity of substitution the more the allocation of resources in such models is distorted due to nominal rigidities. Everything else equal, when aggregate shocks hit the (multisector) economy and relative prices are distorted, a larger ε results in quicker substitution to relatively cheaper goods, which entails

[39]For example, in models such as those of Aoki (2001), Bouakez *et al.* (2009), Carvalho (2006), Carvalho and Nechio (2011), and Nakamura and Steinsson (2010), among others.

a larger deviation from the flexible price (undistorted) case.

6 Conclusion

Even though nested CES preferences are the most commonly used functional form in multisector macroeconomic models, there are very few estimates of the upper-level elasticity of substitution for this preference specification. Those that do exist either do not cover all expenditure categories over which aggregate preferences are defined or are based on estimation methods that inherently suffer from endogeneity bias.

In this paper we estimated this upper-level elasticity of substitution using data on all expenditure categories included in the calculation of the HICP. Moreover, our estimate of the elasticity is identified from the long-run response of relative prices to changes in relative VAT rates. Since these VAT rate changes are essentially exogenous, our estimate is not susceptible to endogeneity bias.

Using a newly constructed data set on monthly inflation and VAT rates for 74 expenditure categories in 25 E.U. countries from 1996 to 2015, we estimate the upper-level elasticity of substitution to be between 1 for the 1-digit level of aggregation of expenditures and 3 for the 3-digit level. Taking into account the uncertainty around these estimates, a value of 5 for the elasticity at the low (3-digit) level of aggregation is the upperbound for the elasticity for which we find support in the data.

Our estimates are meant to provide more informed guidance for the choice of this parameter in quantitative macroeconomic studies. In many applications it has been chosen to equal 1, such that upper-level preferences are Cobb-Douglas. Our results suggest that this is not an unreasonable choice, as long as this upper-level is defined over 1-digit level expenditure aggregates. At lower levels of aggregation an upper-level elasticity of substitution of 3 is consistent with our estimates.

References

Andrade, Philippe, Bénassy-Quéré, Agnès, and Carré, Martine. 2010. *Competition and pass-through on international markets: Firm level evidence from VAT shocks*. mimeo.

Aoki, Kosuke. 2001. Optimal monetary policy responses to relative-price changes. *Journal of Monetary Economics*, **48**(1), 55 – 80.

Basu, Susanto, and Fernald, John G. 1997. Returns to Scale in U.S. Production: Estimates and Implications. *Journal of Political Economy*, **105**(2), 249–283.

Berry, Steven T. 1994. Estimating Discrete-Choice Models of Product Differentiation. *The RAND Journal of Economics*, **25**(2), pp. 242–262.

Berry, Steven T., Levinsohn, James, and Pakes, Ariel. 2004. Differentiated Products Demand Systems from a Combination of Micro and Macro Data: The New Car Market. *Journal of Political Economy*, **112**(1), pp. 68–105.

Blanchard, Olivier, and Galí, Jordi. 2007. Real Wage Rigidities and the New Keynesian Model. *Journal of Money, Credit and Banking*, **39**, 35–65.

Bouakez, Hafedh, Cardia, Emanuela, and Ruge-Murcia, Francisco. 2009. The Transmission of Monetary Policy in a Multi-sector Economy. *International Economic Review*, **50**, 1243–1266.

Broda, Christian, and Weinstein, David E. 2006. Globalization and the Gains From Variety. *The Quarterly Journal of Economics*, **121**(2), 541–585.

Calvo, Guillermo. 1983. Staggered Prices in a Utility Maximizing Framework. *Journal of Monetary Economics*, **12**, 383–398.

Carare, Alina, and Danninger, Stephan. 2008. *Inflation Smoothing and the Modest Effect of VAT in Germany*.

Carvalho, Carlos. 2006. Heterogeneity in Price Stickiness and the Real Effects of Monetary Shocks. *Frontiers of Macroeconomics*, **2**(1).

Carvalho, Carlos, and Nechio, Fernanda. 2011. Aggregation and the PPP Puzzle in a Sticky-Price Model. *The American Economic Review*, **101**(6), pp. 2391–2424.

Carvalho, Carlos, and Nechio, Fernanda. 2014. *Factor Specificities and Real Rigidities*.

Costinot, Arnaud, and Rodríguez-Clare, Andrés. 2014. Chapter 4 - Trade Theory with Numbers: Quantifying the Consequences of Globalization. *Pages 197 – 261 of:* Gopinath, Gita, Helpman, Elhanan, and Rogoff, Kenneth (eds), *Handbook of International Economics*. Handbook of International Economics, vol. 4. Elsevier.

Dawkins, Christina, Srinivasan, T.N., and Whalley, John. 2001. Chapter 58 - Calibration. *Pages 3653 – 3703 of:* Heckman, James J., and Leamer, Edward (eds), *Handbook of Econometrics*. Handbook of Econometrics, vol. 5. Elsevier.

Deaton, Angus, and Muellbauer, John. 1980. An Almost Ideal Demand System. *The American Economic Review*, **70**(3), pp. 312–326.

Dixit, Avinash K., and Stiglitz, Joseph E. 1983. Monopolistic Competition and Optimum Product Diversity. *American Economic Review*, **67**, 297–308.

European Commission. 2007. *Employment in Europe*. European Commission.

European Commission. 2015. *VAT Rates Applied in the Member States of the European Union*. European Commission.

Eurostat. 2009. *HICP CT manual*. European Commission.

Eurostat. 2015. *HICP at constant tax rates*. European Commission.

Eusepi, Stefano, Hobijn, Bart, and Tambalotti, Andrea. 2011. CONDI: A Cost-of-Nominal-Distortions Index. *American Economic Journal: Macroeconomics*, **3**(3), 53–91.

Feenstra, Robert C., Luck, Philip A., Obstfeld, Maurice, and Russ, Katheryn N. 2014. *In Search of the Armington Elasticity*. NBER Working Papers 20063. National Bureau of Economic Research, Inc.

Fisher, Franklin M. 1969. The Existence of Aggregate Production Functions. *Econometrica*, **37**(4), pp. 553–577.

Gabriel, Peter, and Reiff, Adam. 2010. Price setting in Hungary: A store-level analysis. *Managerial and Decision Economics*, **31**, 161–176.

Gautier, Erwan, and Lalliard, Antoine. 2014. How do VAT changes affect inflation in France? *Banque de France Quarterly Selection of Articles*, **32**.

Herrendorf, Berthold, Rogerson, Richard, and Ákos Valentinyi. 2014. Chapter 6 - Growth and Structural Transformation. *Pages 855 – 941 of:* Aghion, Philippe, and Durlauf, Steven N. (eds), *Handbook of Economic Growth*. Handbook of Economic Growth, vol. 2. Elsevier.

Hobijn, Bart, Ravenna, Federico, and Tambalotti, Andrea. 2006. Menu Costs at Work: Restaurant Prices and the Introduction of the Euro. *The Quarterly Journal of Economics*, **121**(3), pp. 1103–1131.

Imbs, Jean, and Mejean, Isabelle. 2015. Elasticity Optimism. *American Economic Journal: Macroeconomics*, **7**(3), 43–83.

Jaimovich, Nir. 2007. Firm dynamics and markup variations: Implications for sunspot equilibria and endogenous economic fluctuations. *Journal of Economic Theory*, **137**(1), 300–325.

Jordà, Òscar. 2005. Estimation and Inference of Impulse Responses by Local Projections. *The American Economic Review*, **95**(1), pp. 161–182.

Karadi, Peter, and Reiff, Adam. 2007. *Menu Costs and Inflation Asymmetries Some Micro Data Evidence*. IEHAS Discussion Papers 0706. Institute of Economics, Centre for Economic and Regional Studies, Hungarian Academy of Sciences.

Karadi, Peter, and Reiff, Adam. 2014. *Menu Costs, Aggregate Fluctuations, and Large Shocks*. CEPR Discussion Paper 10138.

Klenow, Peter J., and Kryvtsov, Oleksiy. 2008. State-Dependent or Time-Dependent Pricing: Does It Matter for Recent U.S. Inflation? *The Quarterly Journal of Economics*, **123**(3), pp. 863–904.

Leslie, Phillip. 2004. Price Discrimination in Broadway Theater. *The RAND Journal of Economics*, **35**(3), pp. 520–541.

Long, John B., and Plosser, Charles I. 1983. Real Business Cycles. *Journal of Political Economy*, **91**(1), pp. 39–69.

Manning, Willard G., Newhouse, Joseph P., Duan, Naihua, Keeler, Emmett B., and Leibowitz, Arleen. 1987. Health Insurance and the Demand for Medical Care: Evidence from a Randomized Experiment. *The American Economic Review*, **77**(3), pp. 251–277.

Midrigan, Virgiliu. 2011. Menu Costs, Multiproduct Firms, and Aggregate Fluctuations. *Econometrica*, **79**(4), pp. 1139–1180.

Nakamura, Emi, and Steinsson, Jón. 2010. Monetary Non-neutrality in a Multisector Menu Cost Model. *The Quarterly Journal of Economics*, **125**(3), pp. 961–1013.

Nevo, Aviv. 2001. Measuring Market Power in the Ready-to-Eat Cereal Industry. *Econometrica*, **69**(2), pp. 307–342.

Ngai, L. Rachel, and Pissarides, Christopher A. 2007. Structural Change in a Multisector Model of Growth. *American Economic Review*, **97**(1), 429–443.

Petrin, Amil. 2002. Quantifying the Benefits of New Products: The Case of the Minivan. *Journal of Political Economy*, **110**(4), pp. 705–729.

Seely, Antony. 2013. *VAT : the temporary cut in the standard rate.* House of Commons: Business and Transportation Section.

Simonovska, Ina, and Waugh, Michael E. 2014. The elasticity of trade: Estimates and evidence. *Journal of International Economics*, **92**(1), 34 – 50.

Wilcox, David W. 1992. The Construction of U.S. Consumption Data: Some Facts and Their Implications for Empirical Work. *The American Economic Review*, **82**(4), pp. 922–941.

Woodford, Michael. 2003. *Interest and Prices: Foundations of a Theory of Monetary Policy.* Princeton University Press.

Wooldridge, Jeffrey M. 2001. *Econometric Analysis of Cross Section and Panel Data.* MIT Press Books, vol. 1. The MIT Press.

Table 1: Estimates for $h = 48$ for various model specifications

	(I)	(II)	(III)
β	0.40	0.38	0.47
	(0.04)	(0.06)	(0.04)
No. of obs.	293,625	148,270	254,222
Degrees of freedom	268,650	125,702	229,954
Sample:			
Period	1996-2015	1996-2007	1996-2015
No. of COICOPs	74	74	74
Lag inflation	No	No	Yes
Implied elasticity:			
ε	2.98	3.22	2.27
95% conf. interval	[2.1,4.3]	[1.9,5.7]	[1.7,3.1]

Note: Standard errors in parentheses. Implied elasticities calculated using $\alpha = 1/3$. 95% confidence interval constructed by calculating implied elasticities for boundaries of symmetric confidence interval for estimated elasticity, β.

Table 2: Estimates for $h = 48$ for various levels of aggregation

	(*I*) *Class*	(*II*) *Group*	(*III*) *Division*
β	0.40	0.43	0.66
	(0.04)	(0.08)	(0.17)
No. of obs.	293,635	128,171	16,863
Degrees of freedom	268,666	115,061	11,466
Sample:			
Period	1996-2015	1996-2015	1996-2015
No. of COICOPs	74	36	10
Lag inflation	No	No	No
Implied elasticity:			
ε	2.98	2.61	1.01
95% conf. interval	[2.1,4.3]	[1.4,5.1]	[0.0,4.2]

Note: Standard errors in parentheses. Implied elasticities calculated using $\alpha = 1/3$. 95% confidence interval constructed by calculating implied elasticities for boundaries of symmetric confidence interval for estimated elasticity, β.

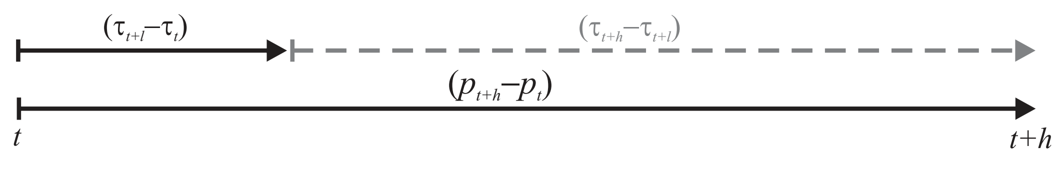

Figure 1: Identification strategy

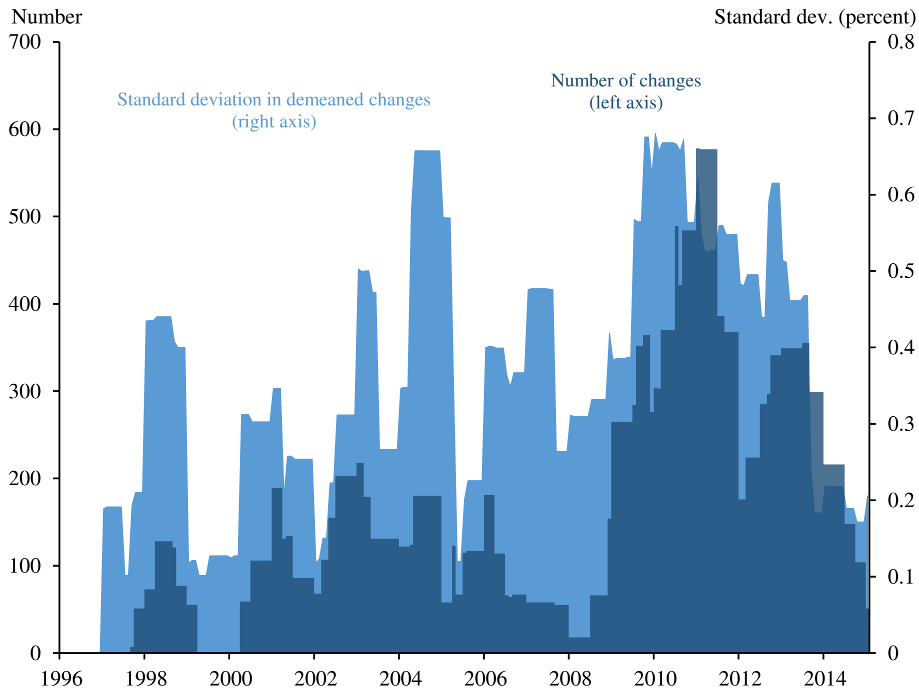

Figure 2: Time series of variation in VAT rate changes

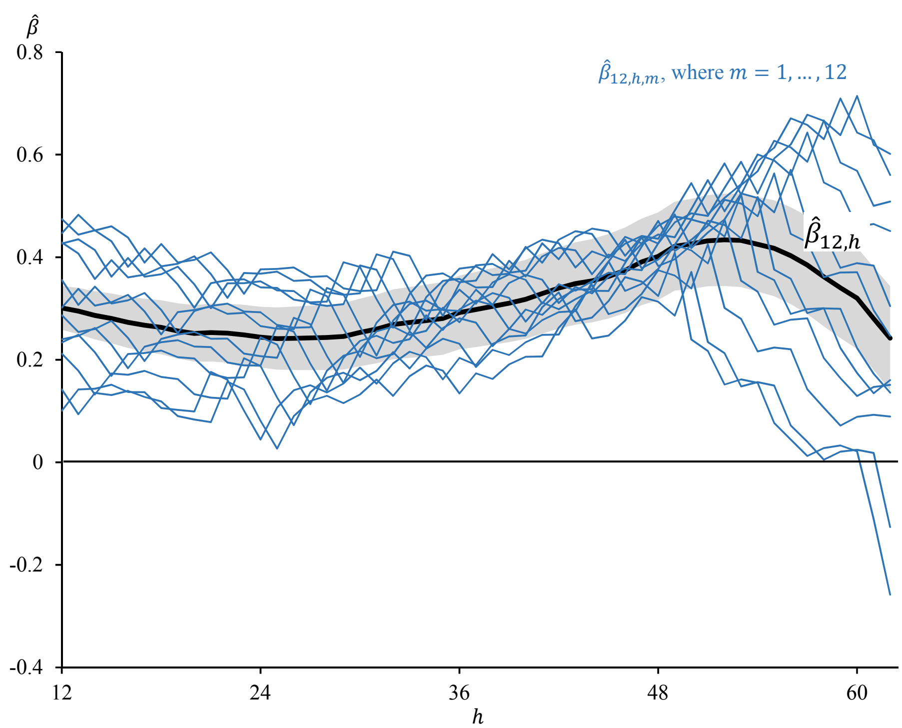

Figure 3: Estimated reduced-form parameter, β, for baseline and unrestricted specifications.

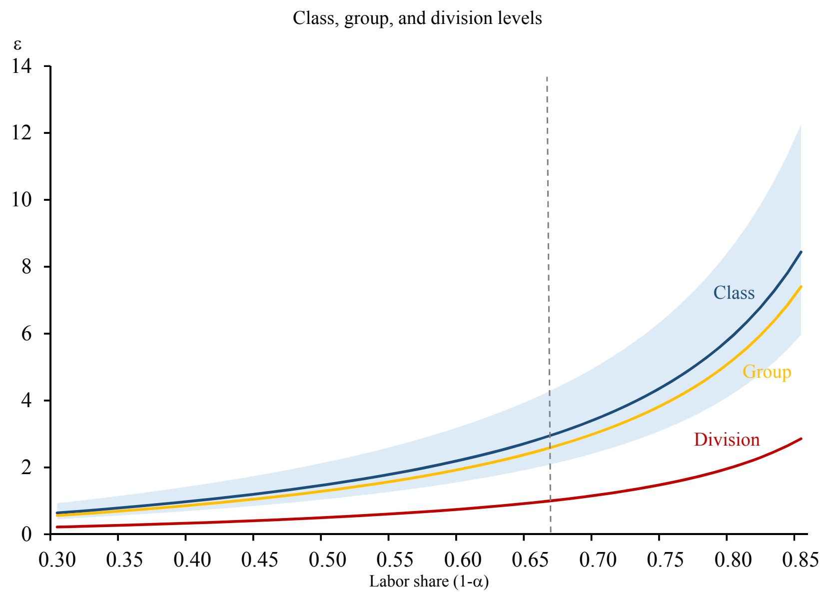

Figure 4: Estimated elasticity ε for varying labor share $(1 - \alpha)$

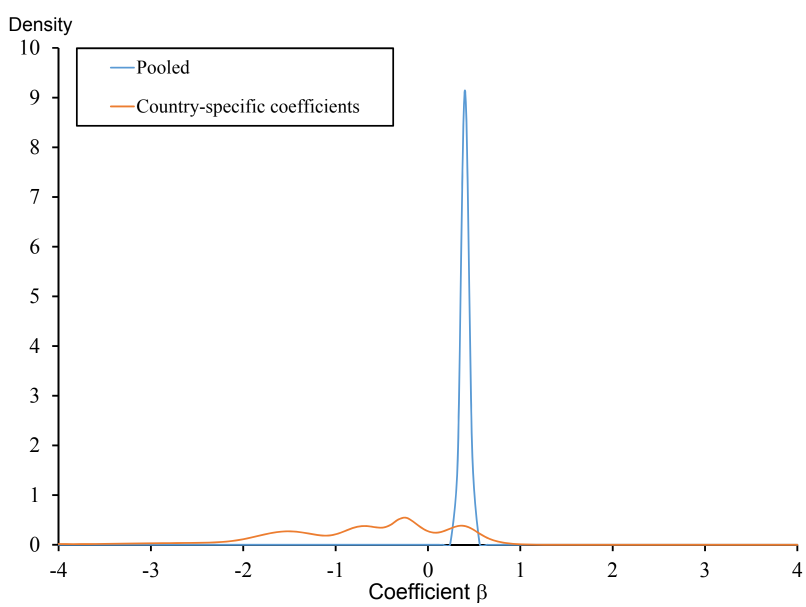

Figure 5: Distributions of pooled and non-pooled coefficients

A Data appendix

We measure inflation using the change in the logarithm of harmonized consumer price indices (HICP) by consumer expenditure category. Expenditures covered by HICPs are classified in categories/goods called COICOPs. Our data covers the time period January 1996 through January 2015.

The COICOP classification system consists of three levels of aggregation of categories. In our sample, the top level consists of 10 *divisions*. Our data originally included 12 divisions, from which we dropped all classes (and groups) pertaining divisions 6 (Health) and 10 (Education) because of the non-market price setting in these two sectors. In addition, we do not include two additional divisions that cover spending by non-profit institutions serving households (NPISH) and government consumption. The next level of aggregation is called a *group*. Our sample includes 36 groups. The lowest level of aggregation is a *class*. Our sample includes 74 classes.

Data on VAT rates by COICOP for the 25 European Union countries in our sample are not readily available. We construct them from two administrative sources: European Commission (2015), and Eurostat (2015). These give us information about which VAT rates are applicable to goods and services in a country at a point in time.

To construct our data set, we match all COICOPs for which we have price data for to their corresponding VAT rates at the lowest level of aggregation – the class level. VAT rates for the group- and division-level categories are (for the most part) constructed as weighted averages of the VAT rates in the underlying classes.

To aggregate VAT rates up from the class to the group level, whenever the applied VAT is not directly listed on either European Commission (2015) or Eurostat (2015), we apply a simple average of VAT rates across all classes within each group.

To aggregate up from the group level to the division level, we make use of expenditure data by group and division. In particular, we use Eurostat data on consumer expenditures by COICOP.[40] For each country and COICOP, consumer expenditures are available at an annual frequency up to 2013. We match our monthly data set of VAT rates and prices to yearly consumer expenditure shares assuming the latter are unchanged across all months of each year, and using the closest year to fill in missing expenditures data.[41] We then compute the share of expenditures in each group over total division expenditures. Finally, whenever

[40] We download the series named "nama-co3-k." For all countries and categories, units are in millions of 2005 euros.

[41] When matching COICOPs to yearly data on expenditures, for each division at a certain year, we either match expenditure values for all the groups or for none of them (assigning missing values in the latter case). We do so to avoid mixing and matching expenditure values from different years.

the applied VAT is not directly listed on either European Commission (2015) or Eurostat (2015), we aggregate VAT rates up from the group to the division level by calculating the expenditure-share weighted averages of the VAT rates across groups within each division.

Because this aggregation process potentially introduces noise to our measures of VAT rate, for both the group-level and the division-level estimations, we, respectively, exclude categories in which the within-group and the within-division dispersions of VAT rates is larger than one.[42]

Tables A1 to A4 report details of the data the we actually include in our sample after performing the adjustments just described.

Table A1 lists all countries included in the sample. For each country and level of aggregation, it reports the number of 12-month VAT rate changes (between 1996 to 2015), the number of categories (COICOPs) with non-zero VAT rate changes, and the number of non-zero *demeaned* changes in VAT rates that result from the absorption of COICOP-country-month (α_{jcm}) and country-time (γ_{ct}) fixed effects. The latter variation is the one we actually explore in our empirical approach (see equation (22) in the main document).

Tables A2 to A4 provide analogous information by expenditure category. In particular, they report descriptive statistics on VAT variation for all categories at all levels of aggregation. The first column reports the number of nonzero changes in VAT rates within 12-month for each category. The second column shows the number of countries with nonzero VAT rate changes (out of our sample of 25 countries) for each COICOP. These two columns reports statistics based on simple 12-month changes in VAT rates. The last column of Tables A2 to A4 reports the number of nonzero demeaned-VAT rate changes.

In conjunction with Figure 2 reported in the main text, Tables A1 to A4 suggest that the source of variability in our data does not rely on any specific time, country or category.

B Additional empirical results

For completeness, in this section we report the results from Figure 3 and Table 1 the group and division levels. In particular, Figures 6 and 7 report the estimated group- and division-levels $\beta_{l,h}$ for $l = 12$ and $h = 12, \ldots, 62$, respectively. The figures show that the patterns observed at the class level are similar at the group and division levels, despite the much wider 95% confidence bands.

Table A5 reports the group-level estimated β and the implied elasticity ε for the full sample, for the 1996-2007 period, and for the version that includes lagged inflation as an additional

[42] Estimated coefficients at the group and division levels that include all categories irrespective of the within-level VAT rate dispersion suffer from attenuation bias or are not statistically significant.

control variable. Table A6 performs the analogous exercise at the division level. Column *(I)* of Tables A5 and A6 replicate the results reported in Table 2 on the main text document. Columns *(II)* and *(III)* in both Tables A5 and A6 show that, at these levels of aggregation, because the sample size is much more limited, the precision of our estimates is much reduced, particularly when the sample is further limited by dropping the later part of the time sample (Column *(II)*). For both estimates at the group and the division levels, further limiting the sample makes the implied elasticity (ε) estimates not statistically different from zero.

Table A1: Variation in VAT rate changes by country

	Class			Group			Division		
	No. of $\Delta\tau_t \neq 0$	COICOPs with $\Delta\tau_t \neq 0$	Demeaned $\Delta\tau_t \neq 0$	No. of $\Delta\tau_t \neq 0$	COICOPs with $\Delta\tau_t \neq 0$	Demeaned $\Delta\tau_t \neq 0$	No. of $\Delta\tau_t \neq 0$	COICOPs with $\Delta\tau_t \neq 0$	Demeaned $\Delta\tau_t \neq 0$
Austria	0	0	0	0	0	0	192	1	868
Belgium	24	2	15841	12	1	7161	0	0	0
Cyprus	3120	46	14544	1496	24	6348	299	3	963
Czech Republic	2700	67	15841	1020	27	6727	156	4	868
Denmark	0	0	0	0	0	0	0	0	0
Estonia	828	65	15745	372	28	7547	60	5	0
Finland	2052	65	16058	654	25	6293	93	3	651
France	1296	53	15679	552	23	6510	60	3	855
Germany	1320	54	15841	624	26	6944	96	4	813
Greece	2014	64	15841	690	24	6293	90	3	606
Hungary	2438	66	15841	972	27	6998	187	5	936
Ireland	2568	53	15841	948	17	4774	184	2	534
Italy	1584	42	15841	792	22	6293	108	3	633
Latvia	2381	61	15275	996	27	7049	144	4	868
Lithuania	1512	65	15902	696	29	7161	120	5	765
Luxembourg	51	50	1387	18	18	475	193	1	434
Malta	588	36	15624	264	20	6076	12	1	0
Netherlands	1056	39	15624	504	19	6293	48	2	651
Poland	948	69	15841	300	24	6076	153	3	627
Portugal	2862	67	16058	1332	27	6727	178	3	709
Slovakia	2472	65	15793	1080	30	7595	228	6	1338
Slovenia	1644	65	12985	540	22	5049	72	3	525
Spain	1344	49	15841	576	24	6293	284	4	1056
Sweden	0	0	0	0	0	0	0	0	0
United Kingdom	1848	47	15841	948	25	6727	144	4	204

Table A2: Variation in VAT rate changes by expenditure category

	Expenditure Categories:	No. of $\Delta\tau_t \neq 0$ within 12-months	No. of countries with $\Delta\tau_t \neq 0$	No. of demeaned $\Delta\tau_t \neq 0$ changes
1	Food and non-alcoholic beverages	618	11	2835
1.1	Food	295	10	4263
1.11	Bread and cereals	331	11	4480
1.12	Meat	331	11	4480
1.13	Fish and seafood	331	11	4480
1.14	Milk, cheese and eggs	331	11	4480
1.15	Oils and fats	331	11	4480
1.16	Fruit	331	11	4480
1.17	Vegetables	331	11	4480
1.18	Sugar, jam, honey, chocolate and confectionery	331	11	4480
1.19	Food products n.e.c.	331	11	4480
1.2	Non-alcoholic beverages	319	11	2961
1.21	Coffee, tea and cocoa	343	12	4480
1.22	Mineral waters, soft drinks, fruit and vegetable juices	523	18	4480
2	Alcoholic beverages, tobacco and narcotics	823	20	3909
2.1	Alcoholic beverages	572	19	4298
2.11	Spirits	627	21	4534
2.12	Wine	591	21	4534
2.13	Beer	627	21	4534
2.2	Tobacco	627	21	4534
3	Clothing and footwear	542	13	2691
3.1	Clothing	442	13	3038
3.11	Clothing materials	627	17	4534
3.12	Garments	627	21	4534
3.13	Other articles of clothing and clothing accessories	627	21	4534
3.14	Cleaning, repair and hire of clothing	561	19	4534
3.2	Footwear	627	21	4534
4	Housing, water, electricity, gas and other fuels	0		0
4.1	Actual rentals for housing	0	0	4576
4.3	Maintenance and repair of the dwelling	317	13	2840
4.31	Materials for the maintenance and repair of the dwelling	627	21	4534
4.32	Services for the maintenance and repair of the dwelling	425	18	4480
4.4	Water supply and miscellaneous services relating to the dwelling	168	6	1302
4.41	Water supply	288	11	4480
4.42	Refuse collection	396	14	4480
4.43	Sewerage collection	520	18	4480
4.44	Other services relating to the dwelling n.e.c.	627	19	4534
4.5	Electricity, gas and other fuels	346	12	2779
4.51	Electricity	537	20	4450
4.52	Gas	537	19	4450
4.53	Liquid fuels	555	16	4534
4.54	Solid fuels	572	16	4388
4.55	Heat energy	553	12	4482

Table A3: Variation in VAT rate changes by expenditure category (cont.)

	Expenditure Categories:	No. of $\Delta\tau_t \neq 0$ within 12-months	No. of countries with $\Delta\tau_t \neq 0$	No. of demeaned $\Delta\tau_t \neq 0$ changes
5	Furnishings, household equipment and routine household maintenance	470	15	3007
5.1	Furniture and furnishings, carpets and other floor coverings	615	20	4317
5.11	Furniture and furnishings	615	20	4534
5.12	Carpets and other floor coverings	627	21	4534
5.13	Repair of furniture, furnishings and floor coverings	627	10	4534
5.2	Household textiles	627	21	4534
5.3	Household appliances	627	21	4534
5.31	Major and small household appliances	627	21	4534
5.33	Repair of household appliances	627	21	4534
5.4	Glassware, tableware and household utensils	627	21	4534
5.5	Tools and equipment for house and garden	627	21	4534
5.6	Goods and services for routine household maintenance	482	16	3689
5.61	Non-durable household goods	627	21	4534
5.62	Domestic services and household services	591	21	4534
7	Transport	443	6	1179
7.1	Purchase of vehicles	627	21	4534
7.11	Motor cars	627	21	4534
7.12	Motor cycles, bicycles and animal drawn vehicles	627	20	4534
7.2	Operation of personal transport equipment	442	13	3038
7.21	Spare parts and accessories for personal transport equipment	627	21	4534
7.22	Fuels and lubricants for personal transport equipment	627	21	4534
7.23	Maintenance and repair of personal transport equipment	561	19	4534
7.24	Other services in respect of personal transport equipment	627	21	4534
7.3	Transport services	144	5	1085
7.31	Passenger transport by railway	347	14	4016
7.32	Passenger transport by road	347	15	4480
7.33	Passenger transport by air	375	14	4576
7.34	Passenger transport by sea and inland waterway	375	9	4534
7.35	Combined passenger transport	48	1	651
7.36	Other purchased transport services	603	13	4317
8	Communications	0		0
8.1	Postal services	0	0	4359
8.2	Telephone and telefax equipment	567	16	4317
8.3	Telephone and telefax services	567	16	4317

Table A4: Variation in VAT rate changes by expenditure category (cont2.)

	Expenditure Categories:	No. of $\Delta\tau_t \neq 0$ within 12-months	No. of countries with $\Delta\tau_t \neq 0$	No. of demeaned $\Delta\tau_t \neq 0$ changes
9	Recreation and culture	12	0	36
9.1	Audio-visual, photographic and information processing equipment	627	21	4534
9.11	Equipment for sound and picture	627	21	4534
9.12	Photographic and cinematographic equipment and optical instruments	627	20	4534
9.13	Information processing equipment	627	21	4534
9.14	Recording media	627	21	4534
9.15	Repair of audio-visual, photographic and information processing equipment	627	20	4534
9.2	Other major durables for recreation and culture	627	19	4534
9.21	Major durables for indoor and outdoor recreation	627	19	4534
9.23	Maintenance and repair of other major durables for recreation and culture	627	4	4534
9.3	Other recreational items and equipment, gardens and pets	603	20	4317
9.31	Games, toys and hobbies	627	21	4534
9.32	Equipment for sport, camping and open-air recreation	627	21	4534
9.33	Gardens, plants and flowers	627	21	4534
9.34	Pets and related products; veterinary and other services for pets	627	21	4534
9.4	Recreational and cultural services	382	11	2821
9.41	Recreational and sporting services	590	19	4534
9.42	Cultural services	573	18	4534
9.5	Newspapers, books and stationery	12	1	217
9.51	Books	258	11	4248
9.52	Newspapers and periodicals	306	13	4296
9.53	Miscellaneous printed matter; stationery and drawing materials	627	21	4534
9.6	Package holidays	627	21	4534
11	Restaurants and hotels	193	7	1247
11.1	Catering services	418	15	3430
11.11	Restaurants, cafs and the like	490	18	4318
11.12	Canteens	552	18	4534
11.2	Accommodation services	406	16	4157
12	Miscellaneous goods and services	0		0
12.1	Personal care	406	13	3152
12.11	Hairdressing salons and personal grooming establishments	527	19	4480
12.12	Electrical appliances and articles for personal care	627	21	4534
12.3	Personal effects n.e.c.	627	21	4534
12.31	Jewellery, clocks and watches	627	21	4534
12.32	Other personal effects	627	21	4534
12.4	Social protection	347	12	4576
12.5	Insurance	12	1	4576
12.52	Insurance connected with the dwelling	12	1	4576
12.53	Insurance connected with health	12	0	4576
12.54	Insurance connected with transport	12	1	4576
12.55	Other insurance	12	0	4576
12.6	Financial services n.e.c.	0	0	4576
12.7	Other services n.e.c.	627	21	4534

Table A5: Estimates for $h = 48$ for various model specifications at the group level

	(I)	(II)	(III)
β	0.43	-0.15	0.66
	(0.08)	(0.11)	(0.08)
No. of obs.	128,171	64,703	110,858
Degrees of freedom	115,061	53,828	98,336
Sample:			
Period	1996-2015	1996-2007	1996-2015
No. of COICOPs	36	36	36
Lag inflation	No	No	Yes
Implied elasticity:			
ε	2.61	-15.62	1.03
95% conf. interval	[1.4,5.1]	[31.2,-7.7]	[0.5,1.9]

Note: Standard errors in parentheses. Implied elasticities calculated using $\alpha = 1/3$. 95% confidence interval constructed by calculating implied elasticities for boundaries of symmetric confidence interval for estimated elasticity, β.

Table A6: Estimates for $h = 48$ for various model specifications at the division level

	(I)	(II)	(III)
β	0.66	0.36	0.78
	(0.17)	(0.21)	(0.20)
No. of obs.	16,863	8,911	14,515
Degrees of freedom	11,466	5,664	9,718
Sample:			
Period	1996-2015	1996-2007	1996-2015
No. of COICOPs	36	36	36
Lag inflation	No	No	Yes
Implied elasticity:			
ε	1.01	3.56	0.56
95% conf. interval	[0.0,4.2]	[-45.7,0.6]	[-0.3,3.0]

Note: Standard errors in parentheses. Implied elasticities calculated using $\alpha = 1/3$. 95% confidence interval constructed by calculating implied elasticities for boundaries of symmetric confidence interval for estimated elasticity, β.

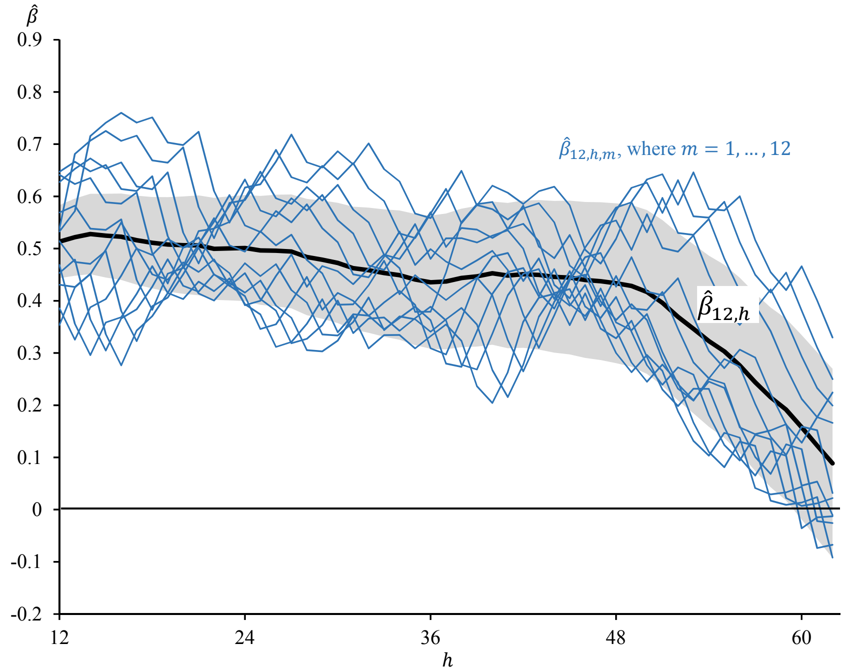

Figure 6: Estimated reduced-form parameter, β, for baseline and unrestricted specifications at the group level.

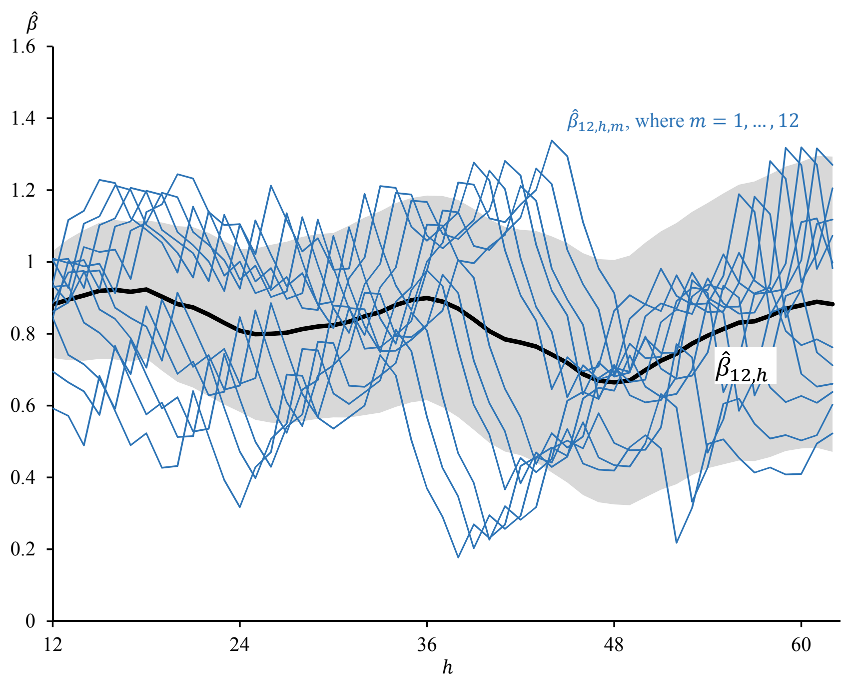

Figure 7: Estimated reduced-form parameter, β, for baseline and unrestricted specifications at the division level.