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How Well Does a Small Structural Model with Sticky Prices and Wages Fit Postwar U.S. Data?[1]

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Résumé

Dans cet article, nous évaluons dans quelle mesure un modèle structurel de petite taille – incluant à la fois des prix et des salaires visqueux, ainsi que des hypothèses permettant d'augmenter la complémentarité stratégique entre les firmes ayant un pouvoir de monopole sur la fixation des prix – est capable de répliquer les données américaines d'après guerre. Pour répondre à ce problème, nous évaluons le modèle en deux étapes. Dans une première étape, nous estimons le modèle en minimisant la distance entre les autocovariances théoriques des variables macroéconomiques clefs et leurs contreparties empiriques obtenues à partir d'un modèle VAR. Dans une seconde étape, nous appliquons la procédure de Watson (1993) [Measures of fit for calibrated models. Journal of Political Economy 101 (6), 1011–1041] afin de quantifier la capacité du modèle à répliquer les données. Notre principal résultat est que la combinaison des prix et salaires visqueux est nécessaire pour obtenir une bonne adéquation entre le modèle et les données. Notre analyse révèle aussi que le modèle composé uniquement de salaires visqueux n'est pas satisfisant au regard des résultats proposés par la procédure de Watson (1993).

Mots-clefs : Prix et salaires visqueux, complémentarité stratégique, test de Watson.

Codes JEL : C52, E31, E32.

Abstract

In this paper, we ask whether a small structural model with sticky prices and wages, embedding various modelling devices designed to increase the degree of strategic complementarity between price-setters, can fit postwar US data. To answer this question, we resort to a two-step empirical evaluation of our model. In a first step, we estimate the model by minimizing the distance between theoretical autocovariances of key macroeconomic variables and their VAR-based empirical counterparts. In a second step, we resort to Watson's (1993) procedure [Measures of fit for calibrated models. Journal of Political Economy 101 (6), 1011–1041] to quantify the model's goodness-of-fit. Our main result is that the combination of sticky prices and sticky wages is central in order to obtain a good empirical fit. Our analysis also reveals that a model with only sticky wages does not perform well according to Watson's criterion (1993).

Keywords: Sticky prices, sticky wages, strategic complementarities, Watson's test.

JEL Codes: C52, E31, E32.

Résumé non technique :

Les modèles de type Nouveaux Keynésiens incorporant des rigidités nominales sur les prix et les salaires ont reçu une attention considérable ces dernières années et sont de plus en plus utilisés dans une optique d'analyse des politiques économiques. De plus, les développements théoriques récents associés à ce type de modèle ont déclenché un bon nombre de travaux de recherche portant sur l'évaluation empirique de ceux-ci. L'objectif de ce papier est de contribuer à cette littérature en effectuant une évaluation formelle d'une version améliorée du modèle Nouveau Keynésien qui est habituellement utilisé dans la littérature. Nous réalisons l'évaluation de ce modèle sur des données trimestrielles des Etats Unis sur la période 1965(1)-2002(4). Nous procédons en trois étapes.

Dans une première étape, nous développons un modèle structurel d'Equilibre Général Dynamique Stochastique de petite taille qui est composé à la fois de prix et de salaires visqueux. Par ailleurs, nous introduisons différents éléments de modélisation qui sont reconnus comme améliorant la capacité du modèle à reproduire les données. Tout d'abord, afin de reproduire au mieux les propriétés de persistance de l'inflation et de la production, nous supposons la formation d'habitudes sur la consommation et l'existence d'un processus d'indexation des prix et des salaires. Ensuite, nous supposons la présence de biens matériels comme facteur de production et d'une élasticité variable de la demande de biens afin de générer une persistance de l'inflation suffisamment élevée, sans pour autant avoir un degré de rigidité des prix trop élevé.

Dans une deuxième étape, nous estimons les paramètres du modèle en utilisant les Moindres Carrés Asymptotiques appliqués sur les autocovariances obtenues à partir d'un modèle Vectoriel AutoRegressif (VAR) canonique sur le taux de croissance de la production, l'inflation, l'inflation salariale, et le taux d'intérêt nominal de court terme. Cette méthode consiste à identifier les paramètres de notre modèle structurel qui minimisent la distance entre les autocovariances théoriques de ces variables et leurs contreparties empiriques obtenues à partir du modèle VAR. En répliquant aussi précisément que possible la version tronquée des autocovariances obtenues par le modèle VAR, nous nous assurons que le modèle étudié est capable de reproduire les caractéristiques clefs relatives à la persistance et aux co-mouvements des données américaines. Par ailleurs, l'introduction de viscosité des salaires dans ce type de modèle ayant été promu comme un élément essentiel dans les modèles Nouveaux Keynésiens, nous accordons une attention particulière à l'hypothèse de viscosité des prix et celle de viscosité des salaires. Plus précisément, nous estimons tout d'abord le modèle de référence qui est composé à la fois de prix et salaires visqueux, puis nous estimons ensuite un

modèle alternatif qui combine salaires visqueux et prix flexibles et un autre modèle composé de prix visqueux et salaires flexibles.

Dans la dernière étape, nous évaluons la capacité du modèle à reproduire les mouvements de court et long terme de ces variables macroéconomiques en implémentant la procédure de Watson (1993). Celle ci consiste à décomposer, dans le domaine des fréquences, l'erreur nécessaire pour réconcilier le spectre du VAR et celui du modèle. Un avantage important de cette procédure est qu'elle offre la possibilité de se concentrer directement sur les fréquences du cycle des affaires.

Nos principaux résultats sont les suivants. Tout d'abord, nous montrons que la combinaison des prix et salaires visqueux est nécessaire pour obtenir une bonne adéquation entre le modèle et les données, puisque le modèle de référence est capable de répliquer les principales caractéristiques des données américaines et les résultats obtenus par la procédure de Watson (1993) aux fréquences du cycle des affaires sont satisfaisants. Par ailleurs, notre analyse montre que le modèle avec seulement des salaires visqueux n'est pas capable de répliquer les spectres empiriques des variables au regard des résultats de proposés par la procédure de Watson (1993) ; tandis que le modèle avec seulement des prix visqueux, bien que moins performant que le modèle de référence, est meilleur que le modèle avec seulement des salaires visqueux. Ce résultat contraste avec ceux qui avaient été obtenus par ailleurs dans la littérature, lesquels suggèrent que la viscosité des salaires impliquent des effets de persistance plus importants que la viscosité des prix (Andersen, 1998; Huang and Liu, 2002; Christiano et al., 2005, par exemple). Cependant, contrairement à ces auteurs, nous nous concentrons sur la persistance globale plutôt que sur la seule persistance conditionnelle aux chocs monétaires.

Non-technical summary:

New Keynesian models embedding nominal rigidities in prices and wages have received considerable attention over the recent years and are being more and more used for policy analysis purposes. The theoretical developments associated with New Keynesian models have also sparked an important set of studies that have proposed to assess the empirical fit of these models. The objective of the present paper is to contribute to this literature by providing a formal assessment of the goodness-of-fit of an augmented version of the basic New Keynesian model that is currently in use in the literature. We implement the assessment of this model on quarterly U.S. data over the sample 1965(1)-2002(4). We approach the things in three stages.

In a first step, we develop a small structural Dynamic Stochastic General Equilibrium (DSGE) model which is featuring stickiness of prices and wages. In addition, we include various additional modelling elements, which are known to enhance the ability of this model to fit the data. Firstly, we assume habit formations on consumption and price and wage indexation schemes, in order to promote the reproduction of the persistence properties of inflation and output. Secondly, we assume the presence of material goods as a production input and a variable elasticity of demand for goods so as to generate a large amount of inflation persistence without relying on an unrealistically high degree of nominal rigidities.

In a second step, we estimate the model's parameters using the Asymptotic Least Square method applied on the autocovariances implied by a canonical VAR on output growth, inflation, wage inflation and the Federal Funds rate. The latter consists in pinning down the structural parameters of our DSGE model so as to minimize the distance between the theoretical autocovariances of these variables and their VAR-based counterparts. By replicating as closely as possible a truncated version of the autocovariances as implied by a canonical VAR, we make sure that the investigated model is able to reproduce key persistence and co-movement characteristics of U.S. data. In addition, since the inclusion of sticky wages assumption has been promoted as an essential feature of the New Keynesian framework, we pay particular attention to the relative merits of sticky prices versus sticky wages in obtaining a good fit. In particular, we first estimate the benchmark model which is composed by prices and wages both sticky, and we then estimate a alternative model which combines sticky wages and flexible prices, and an other with sticky prices and flexible wages.

In the last step, we assess the ability of our model to reproduce short- and long-run movements of these macroeconomic aggregates. To do so, we resort to Watson's (1993) procedure. The latter consists in decomposing in the frequency domain the error necessary to reconcile the VAR-based and the model-based spectra. Thus, it gives us a complete assessment of the model, based on the full set of second order moments, further decomposed in the frequency domain. One key advantage of Watson's (1993) procedure is that it naturally offers the possibility of directly focussing on the business cycle frequency ranges.

Our main results are as follow. First, we show that the combination of sticky prices and sticky wages is necessary in order to obtain a good empirical fit since the benchmark model is able to reproduce the whole characteristics of U.S. data and it is satisfactory according to Watson's (1993) criterion at business cycle frequencies. In addition, our analysis reveals is that a model with only sticky wages does not a good job of fitting the empirical spectra according to Watson's (1993) criterion while a model with only sticky prices, though less successful than the benchmark model, is better than the model with only sticky wages. This result contrasts with those previously obtained in the literature, which all suggest that wages stickiness implies more persistence effects than prices stickiness (Andersen, 1998; Huang and Liu, 2002; Christiano et al., 2005, for instance). However, contrary to these authors, we focus on the overall persistence rather than on the sole persistence conditional on monetary shocks.

1 Introduction

Models embedding nominal rigidities in prices and wages (New Keynesian models hereafter) have received considerable attention over the recent years and are being more and more used for policy analysis purposes. The theoretical developments associated with New Keynesian models have also sparked an important set of studies that have proposed to assess the empirical fit of these models, ranging from Rotemberg and Woodford (1997) to Ireland (2002, 2004) and many other.

The objective of the present paper is to contribute to this literature by providing a formal assessment of the goodness-of-fit of an augmented version of the basic New Keynesian model that is currently in use in the literature. To do that, we have to consider a model that can be reasonably taken to the data. Doing so requires that the model features two types of "frictions": *(i)* mechanisms designed to enhance the ability of small structural Dynamic Stochastic General Equilibrium (DSGE) models to reproduce the persistence properties of inflation and output, such as habit formations and price and wage indexation schemes; *(ii)* mechanisms designed to lower the degree of nominal price rigidity needed to match the data, such as the presence of material goods as a production input and a variable elasticity of demand for goods, as in Basu (1995), Bergin and Feenstra (2000), Kimball (1995), and Woodford (2003).[1] As these authors show, including these elements helps New Keynesian models to generate a large amount of inflation persistence without relying on an unrealistically high degree of nominal rigidities. The contribution of our paper is twofold.

First, although we are in line with a class of estimation methods named parametric estimation, we estimate the parameters of our model using an original method which allows us to address to some issues. Indeed, at least two broad alternative approaches have been adopted in the literature in order to estimate DSGE model. A first strand has opted for Full Information Maximum Likelihood (FIML). Using this estimation method, the results concerning the New Keynesian framework have typically been mixed[2]. However, FIML estimation of DSGE models has often been criticized in the

[1] See also Huang and Liu (2004).

[2] Authors such as Dennis (2004) argue that simple vector autoregressive (VAR) models outperform hybrid New Keynesian models in terms of relative information criteria. Authors such as Fuhrer and Rudebusch (2004), Fuhrer (1997) and Estrella and Furher (2002) argue that sheer versions of the IS and Phillips curves arising from microfounded DSGE models do not fit the data well or imply dynamics that are at odds with the data. Other, such as Roberts (2005) and Ireland (2002, 2004) argue that these models obtain a reasonable fit provided enough lags of inflation and output or enough rigidities are included in the theoretical framework.

literature. A well-known issue is that the method might not be robust to model mis-specification. Accordingly, FIML estimation might be too demanding for models which, by construction, are highly stylized representations of the real world. So as to avoid these problems, a second strand of the literature has increasingly relied on limited information approaches, which emphasize key moments from the data that the DSGE models under study ought to reproduce. In particular, a number of papers have adopted a common empirical strategy based on Minimum Distance Estimation (MDE), following the original work by Rotemberg and Woodford (1997, 1999), which consists in picking the DSGE parameters so as to mimic as closely as possible the responses of key variables to a monetary shock and/or a technology shock, as implied by a structural VAR model.[3] The central empirical message emerging from this literature is that tightly parameterized New Keynesian models can satisfactorily reproduce the economy's dynamic response to these structural VAR-based identified shocks. Yet, a potential issue with this popular approach is that it might not fully exploit the cross-equation restrictions implied by the DSGE models. In particular, there is no reason to expect that a set of parameters selected so as to reproduce the economy's response to a monetary shock will do a good job when confronted with other structural shocks.

Acknowledging these limitations, our contribution is to resort to an estimation method that might be viewed as a compromise between FIML estimation and standard MDE. The latter consists in picking the structural parameters of our DSGE model so as to replicate as closely as possible a truncated version of the autocovariances of output growth, inflation, wage inflation and the Federal Funds rate, as implied by a canonical VAR. More precisely, the parameters of the New Keynesian model are obtained by minimizing the distance between the theoretical autocovariances of these variables and their VAR-based counterparts. In doing so, we make sure that the investigated model is able to reproduce key persistence and co-movement characteristics of U.S. data. In addition, in order to assess the ability of our model to reproduce short- and long-run movements of these macroeconomic aggregates, we resort to Watson's (1993) procedure. The latter consists in decomposing in the frequency domain the error necessary to reconcile the model and the data. An advantage of combining our estimation method and Watson's (1993) procedure is that both are mutually consistent. Indeed, the model's parameters are estimated so as to replicate a truncated version of the VAR-based autocovariances, while Watson's (1993) procedure gives us a complete assessment

[3] A non exhaustive list of contributions includes Amato and Laubach (2003), Boivin and Giannoni (2003), Christiano et al. (2005), and Edge et al. (2003).

of the model, based on the full set of second order moments, further decomposed in the frequency domain. Moreover, Watson's (1993) procedure appears particularly suitable for our purpose. Indeed, as has been often acknowledged, DSGE models are not necessarily meant to account for all the dynamic movements in the data. One key advantage of Watson's (1993) procedure is that it naturally offers the possibility of directly focussing on the frequency ranges along which we seek to assess the performances of our model.

The second contribution of our paper is to pay particular attention to the relative merits of sticky prices versus sticky wages in obtaining a good fit. Indeed, since the inclusion of sticky wages in the baseline model has been promoted as an essential feature of the New Keynesian framework, we assess whether sticky prices and/or sticky wages are necessary to obtain a satisfactory model. Our approach is close in spirit to that retained by Rabanal and Rubio-Ramírez (2005). However, we complete their results since we resort to a different model-assessment tool, and in addition, we go further in our analysis since we assess the relative information of sticky prices and sticky wages.

The empirical results in this paper highlight the pretty good ability of the New Keynesian model to replicate the second order moments of postwar U.S data.[4] In particular, according to Watson's (1993) criterion, the model does a good job in reproducing the empirical spectra of output growth, inflation, the short-term interest rate, and wage inflation at business cycle frequencies. Our main result is that the combination of sticky prices and sticky wages is necessary in order to obtain a good empirical fit. This result has been previously emphasized in the literature. However, an original result which our analysis reveals is that a model with only sticky wages does not perform well according to Watson's criterion (1993) while a model with only sticky prices, though less successful than the benchmark model, is better than the model with only sticky wages. This result is interesting since it is opposed to the findings of Andersen (1998). However, contrary to him, our approach is based on the empirical assessment of a theoretical model.

The remainder is as follows. Section 2 expounds the model. Section 3 presents our estimation procedure and results. In addition, we detail the relative contribution of sticky prices and/or sticky wages. Section 4 assesses the goodness-of-fit of the models using Watson's (1993) procedure. The last section briefly concludes.

[4] Jung (2004) applies Watson's (1993) procedure to a sticky price model with external habit formation and reports very large relative mean square approximation errors (almost always close to 100%). See also Ellison and Scott (2000) for similarly disappointing results when applying Watson's procedure to a simple sticky price model.

2 The Model

2.1 Final and Material Goods

In this sector, perfectly competitive firms produce a homogeneous good that can either be consumed (y_t) or serve as an input in the production of material goods (q_t). Following Woodford (2003), Basu (1995), and Bergin and Feenstra (2000), this modelling device is included so as to increase the degree of strategic complementarity between price-setting firms. Let us define $d_t \equiv y_t + q_t$, the aggregate demand for final and material goods, and let P_t denote the nominal price of good d_t.[5]

So as to reinforce further the degree of strategic complementarity in price-setting decisions, it is also assumed that the aggregate final good is produced by combining intermediate goods through a production function characterized by a variable-elasticity. More precisely, following Kimball (1995), we assume that the final good production function is of the form

$$\int_0^1 G\left(\frac{d_t\left(\varsigma\right)}{d_t}\right)\mathrm{d}\varsigma = 1, \tag{1}$$

where $d_t\left(\varsigma\right)$ denotes the input of intermediate good $\varsigma \in [0,1]$, and the function $G\left(\cdot\right)$ is increasing, strictly concave, and satisfies the normalization $G\left(1\right) = 1$. If we let $P_t\left(\varsigma\right)$ denote the nominal price of intermediate good ς, the overall demand addressed to the producer of intermediate good $\breve{\varsigma}$, $d_t\left(\breve{\varsigma}\right)$, is implicitly defined by

$$G'\left(\frac{d_t(\breve{\varsigma})}{d_t}\right) = \frac{P_t(\breve{\varsigma})}{P_t}\int_0^1 \frac{d_t\left(\varsigma\right)}{d_t}G'\left(\frac{d_t\left(\varsigma\right)}{d_t}\right)\mathrm{d}\varsigma. \tag{2}$$

2.2 Aggregate Labor Index

Following Erceg et al. (2000), we assume for convenience that a set of differentiated labor inputs, indexed by $\upsilon \in [0,1]$, are aggregated into a single labor index ℓ_t by competitive firms, which will be referred to as labor intermediaries in the sequel. They produce the aggregate labor input according to the following CES technology

$$\ell_t = \left(\int_0^1 \ell_t\left(\upsilon\right)^{(\theta_w-1)/\theta_w}\mathrm{d}\upsilon\right)^{\theta_w/(\theta_w-1)},$$

where $\theta_w > 1$ is the elasticity of substitution between any two labor types, and $\ell_t\left(\upsilon\right)$ denotes the input of labor of type υ.

[5] A detailed technical appendix is available from the authors upon request.

Let $W_t(\upsilon)$ denotes the nominal wage rate associated to type-υ labor, which labor intermediaries take as given. The first order conditions are

$$\ell_t(\upsilon) = \left(\frac{W_t(\upsilon)}{W_t}\right)^{-\theta_w} \ell_t, \tag{3}$$

where the aggregate nominal wage, W_t, is defined as

$$W_t = \left(\int_0^1 W_t(\upsilon)^{1-\theta_w}\, \mathrm{d}\upsilon\right)^{1/(1-\theta_w)}. \tag{4}$$

Notice that equation (4) is a direct consequence of the combination of equation (3) and the zero profits condition for labor intermediaries.

2.3 Intermediate Goods

In the third sector, monopolistic firms produce the intermediate goods $\varsigma \in [0, 1]$. Each firm ς is the sole producer of intermediate good ς. Let $\theta(\xi)$ denotes the elasticity of demand for a producer of intermediate good facing the relative demand $\xi = d_t(\varsigma)/d_t$. According to the implicit demand function (2), $\theta(\xi)$ obeys

$$\theta(\xi) = -\frac{G'(\xi)}{\xi G''(\xi)}.$$

This last equation illustrates that intermediate good firms face a varying elasticity of demand for their output, implying a varying markup, which is denoted by $\mu(\xi)$, and obeys

$$\mu(\xi) = \frac{\theta(\xi)}{\theta(\xi) - 1}.$$

Given a demand $d_t(\varsigma)$, we assume that monopolist ς faces the following production possibilities

$$\min \left\{\frac{\mathrm{e}^{z_t} F(n_t(\varsigma))}{1 - s_m}, \frac{m_t(\varsigma)}{s_m}\right\} \geq d_t(\varsigma), \tag{5}$$

where $F(\cdot)$ is an increasing and concave production function, $n_t(\varsigma)$ denotes the input of aggregate labor, $m_t(\varsigma)$ denotes the input of material goods, and s_m is the share of material goods in gross output. This specification, borrowed from Rotemberg and Woodford (1995), allows for an increased degree of strategic complementarity between price-setters, as shown in Woodford (2003). Finally, z_t is a productivity shock which evolves according to the following process

$$z_t = \log(a) + z_{t-1} + \zeta_t, \qquad \zeta_t = \rho_\zeta \zeta_{t-1} + \varepsilon_{\zeta,t},$$

where $\rho_\zeta \in (0,1)$, $\varepsilon_{\zeta,t} \sim \text{iid}(0,\sigma_\zeta)$, and $a > 1$ is the average, gross growth rate of technical progress. The autocorrelation of productivity shocks is meant to capture the effects of gradual technology diffusion, such as those rationalized by Rotemberg (2003).[6]

Cost minimization ensures that

$$m_t(\varsigma) = s_m d_t(\varsigma), \quad \text{and} \quad \mathrm{e}^{z_t} F(n_t(\varsigma)) = (1 - s_m) d_t(\varsigma),$$

so that the real cost $\mathbb{S}(d_t(\varsigma))$ of producing $d_t(\varsigma)$ units of goods ς is

$$\mathbb{S}(d_t(\varsigma)) = w_t F^{-1}\left((1 - s_m)\mathrm{e}^{-z_t} d_t(\varsigma)\right) + s_m d_t(\varsigma),$$

where $w_t \equiv W_t/P_t$.

Following Calvo (1983), we assume that in each period of time, a monopolistic firm can reoptimize its price with probability $1 - \alpha_p$, irrespective of the elapsed time since it last revised its price. If the firm cannot reoptimize its price, the latter is rescaled according to the simple revision rule

$$P_T(\varsigma) = (1 + \delta^p_{t,T}) P_t(\varsigma),$$

where

$$1 + \delta^p_{t,T} = \begin{cases} \prod_{j=t}^{T-1} (1+\pi)^{1-\gamma_p} (1 + \pi_j)^{\gamma_p} & \text{if } T > t \\ 1 & \text{otherwise} \end{cases}, \tag{6}$$

where $\pi_t = P_t/P_{t-1} - 1$ represents the inflation rate, $1 + \pi$ is the steady state inflation rate, and $\gamma_p \in (0,1)$ measures the degree of indexation to the most recently available inflation measure.

Let us consider the pricing behavior of monopolist $\breve{\varsigma}$. Firm $\breve{\varsigma}$ takes the demand function (2) into account when setting its price. Additionally, it takes into account the fact that this price rate will presumably hold for more than one period –except for the automatic revision. Let $P_t^\star(\breve{\varsigma})$ denote the price chosen in period t, and let $d^\star_{t,T}(\breve{\varsigma})$ denote the demand for good $\breve{\varsigma}$ in period T if firm $\breve{\varsigma}$ last reoptimized its price in period t. According to (2), $d^\star_{t,T}(\breve{\varsigma})$ obeys the relationship

$$G'\left(\frac{d^\star_{t,T}(\breve{\varsigma})}{d_T}\right) = \frac{(1 + \delta^p_{t,T}) P_t^\star(\breve{\varsigma})}{P_T} \int_0^1 \frac{d_t(\varsigma)}{d_t} G'\left(\frac{d_t(\varsigma)}{d_t}\right) \mathrm{d}\varsigma. \tag{7}$$

Then, $P_t^\star(\breve{\varsigma})$ is selected so as to maximize

$$\mathrm{E}_t \sum_{T=t}^{\infty} (\beta\alpha_p)^{T-t} \lambda_T \left\{ \frac{(1 + \delta^p_{t,T}) P_t^\star(\breve{\varsigma})}{P_T} d^\star_{t,T}(\breve{\varsigma}) - \mathbb{S}\left(d^\star_{t,T}(\breve{\varsigma})\right) \right\},$$

[6] Altig et al. (2004) and Galí et al. (2003) also consider autocorrelated growth rates of technical progress.

subject to the demand function (7), where λ_t is the representative household's marginal utility of wealth, and $\mathrm{E}_t\{\cdot\}$ is the expectation operator conditional on the information set available as of time t. That λ_t appears in the above maximization program reflects the fact that the representative household is the ultimate owner of firm $\breve{\varsigma}$.

Standard manipulations yield the approximate loglinearized first order condition

$$\hat{\pi}_t - \gamma_p \hat{\pi}_{t-1} = \varkappa \frac{(1-\alpha_p)(1-\beta\alpha_p)}{\alpha_p} (\hat{w}_t + \omega_p \hat{d}_t) + \beta \mathrm{E}_t \left\{ \hat{\pi}_{t+1} - \gamma_p \hat{\pi}_t \right\}, \tag{8}$$

with

$$\varkappa \equiv [(1-\mu s_m)^{-1}(1+\theta_p \epsilon_\mu) + \theta_p \omega_p]^{-1}.$$

In equation (8), $\hat{\pi}_t$ is the logdeviation of $1+\pi_t$, $\hat{d}_t$ and $\hat{w}_t$ are the logdeviations of $d_t \mathrm{e}^{-z_t}$ and $w_t \mathrm{e}^{-z_t}$, respectively[7], $\theta_p \equiv \theta(1)$ is the steady state elasticity of demand for a producer of intermediate good. Following Woodford (2003), we let ϵ_μ denote the elasticity of $\mu(\xi)$ in the neighborhood of $\xi = 1$, i.e. $\epsilon_\mu = \mu'(1)/\mu(1)$. Finally, the composite parameter ω_p obeys

$$\omega_p \equiv -\frac{F''(n)\,n}{F'(n)} \frac{F(n)}{F'(n)\,n}.$$

Here, $F(n)$, $F'(n)$, and $F''(n)$ denote the value of F and its first and second derivatives, evaluated at the steady state value of n.

The composite parameter $\varkappa$ is linked to the degree of strategic complementarity between the price-setting decisions of the intermediate goods producers. More precisely, the smaller $\varkappa$, the higher the degree of strategic complementarity. When the latter is high, as explained by Woodford (2003), inflation and output turn out to be persistent, i.e. adjust gradually to shocks. As is clear from the above definition of $\varkappa$, we can observe that the share of material goods s_m reduces the responsiveness of inflation. Furthermore, ϵ_μ and θ_p play a similar role to that of s_m.

In empirical estimations of the New Keynesian Phillips Curve, it is not uncommon to obtain a very low partial elasticity of inflation with respect to the real marginal cost.[8] Similarly, in estimated DSGE models, this partial elasticity is also found to be very small. Given a certain calibrated value for $\varkappa$, such a small elasticity translates into a high probability of not reoptimizing prices (high α_p),

[7] Given the presence of a stochastic trend in technical progress, consumption and real wages grow at the same rate while labor is constant through time. To obtain a bounded steady state, trending variables dated t are divided through by e^{z_t}.

[8] See Galí and Gertler (1999) and Eichenbaum and Fisher (2004).

e.g. Rabanal and Rubio-Ramírez (2005). It turns out, however, that this probability is thought to be small in practice, as suggested by the results reported by Bils and Klenow (2004). Thus, if one is mainly interested in obtaining a realistic estimate of α_p, it is important to calibrate $\varkappa$ to a small value.[9]

2.4 Households

The economy is inhabited by a continuum of differentiated households, indexed by $\upsilon \in [0, 1]$. A typical household υ acts as a monopoly supplier of type-υ labor. It is assumed that at each point in time only a fraction $1 - \alpha_w$ of the households can set a new wage, which will remain fixed until the next time period the household is drawn to reset its wage. The remaining households simply revise their wages according to the simple rule

$$W_T(\upsilon) = a^{T-t}(1 + \delta^w_{t,T})W_t(\upsilon),$$

where

$$1 + \delta^w_{t,T} = \begin{cases} \prod_{j=t}^{T-1}(1+\pi)^{1-\gamma_w}(1+\pi_j)^{\gamma_w} & \text{if } T > t \\ 1 & \text{otherwise} \end{cases},$$

where $\gamma_w \in (0, 1)$ measures the degree of indexation to the most recently available inflation measure. Notice that we let the households index their nominal wage to past inflation as well as to the average growth rate of technical progress. In addition to being economically realistic, this assumption contributes to ensuring the existence of a well-behaved deterministic steady state.

A typical household must select a sequence of consumption plans and nominal bonds holdings. As such, the above described problem makes the choices of wealth accumulation contingent upon a particular history of wage decisions, thus leading to households heterogeneity. For the sake of tractability, we assume that the momentary utility function is separable across consumption and leisure. Combining this with the assumption of a complete set of contingent claims market, all the households will make the same choices regarding consumption and will only differ by their wage rate and labor supply. This is directly reflected in our notations.

[9] See Eichenbaum and Fisher (2004) for a related discussion in the context of an estimated New Keynesian Phillips Curve.

Household υ's goal in life is to maximize

$$\mathbb{U}_t = \mathrm{E}_t \sum_{T=t}^{\infty} \beta^{T-t} \{ \mathrm{e}^{g_T} \log(y_T - b y_{T-1}) - \mathrm{e}^{\chi_T} V(\ell_T(\upsilon)) \}, \tag{9}$$

where $\beta \in (0,1)$ is the subjective discount factor, $V(\cdot)$ is a convex function measuring labor disutility, and $b \in (0,1)$ is the parameter of habit in consumption. The variable $\ell_t(\upsilon)$ is household υ's labor supply at period t. Finally, g_t and χ_t are preference shocks, to be specified below. Household υ maximizes (9) subject to the sequence of constraints

$$y_t + \frac{b_t}{1+i_t} \leq w_t(\upsilon) \ell_t(\upsilon) + \frac{b_{t-1}}{1+\pi_t} + \mathrm{div}_t, \tag{10}$$

where div_t denotes profits redistributed by monopolistic firms; $w_t(\upsilon) \equiv W_t(\upsilon)/P_t$ is the real wage rate earned by type-υ labor; $b_t \equiv B_t/P_t$, where B_t denotes the nominal bonds acquired in period t and maturing in period $t+1$; $1+i_t$ denotes the nominal interest rate. The first order conditions with respect to y_t and b_t are

$$\lambda_t = \frac{\mathrm{e}^{g_t}}{(y_t - b y_{t-1})} - \beta b \mathrm{E}_t \left\{ \frac{\mathrm{e}^{g_{t+1}}}{y_{t+1} - b y_t} \right\}, \tag{11}$$

$$\lambda_t = (1+i_t) \beta \mathrm{E}_t \left\{ \frac{\lambda_{t+1}}{1+\pi_{t+1}} \right\}, \tag{12}$$

where, as explained before, λ_t is the multiplier on constraint (10). Let us define $\hat{\imath}_t$ and $\hat{y}_t$ as the logdeviations of $1+i_t$ and $y_t \mathrm{e}^{-z_t}$, respectively, and $\hat{\lambda}_t$ as that of $\lambda_t \mathrm{e}^{z_t}$. Additionally, let us define $\bar{b} = b/a$. We thus obtain the approximate loglinear first order conditions

$$(1+\beta\bar{b}^2)\hat{y}_t = \bar{b}\hat{y}_{t-1} + \beta\bar{b}\mathrm{E}_t\{\hat{y}_{t+1}\} - (1-\beta\bar{b})(1-\bar{b})\hat{\lambda}_t + \breve{g}_t + \beta\bar{b}\mathrm{E}_t\{\zeta_{t+1}\} - \bar{b}\zeta_t, \tag{13}$$

$$\hat{\lambda}_t = \hat{\imath}_t + \mathrm{E}_t\{\hat{\lambda}_{t+1} - \hat{\pi}_{t+1} - \zeta_{t+1}\}, \tag{14}$$

where the shock $\breve{g}_t$ is defined by

$$\breve{g}_t = (1-\bar{b})g_t - \beta\bar{b}(1-\bar{b})\mathrm{E}_t\{g_{t+1}\}.$$

Let us now consider the wage setting decision confronting a household drawn to reoptimize its nominal wage rate in period t, say household υ. In the sequel, it will be convenient to define wage inflation $\pi_t^w \equiv W_t/W_{t-1} - 1$. Since the household is a monopoly supplier, it will take the demand function (3) into account when setting its wage. Additionally, it takes into account the fact that this wage rate will presumably hold for more than on period -except for the automatic revision. Now, let $W_t^{\star}(\upsilon)$ denote the nominal wage rate chosen in date t, and $\ell_{t,T}^{\star}(\upsilon)$ denote hours worked in period T if household υ last reoptimized its wage in period t. According to equation (3), $\ell_{t,T}^{\star}(\upsilon)$ obeys the relationship

$$\ell_{t,T}^{\star}(\upsilon) = \left(\frac{a^{T-t}(1+\delta_{t,T}^{w})W_t^{\star}(\upsilon)}{W_T} \right)^{-\theta_w} \ell_T. \tag{15}$$

Then, $W_t^{\star}(\upsilon)$ is selected to maximize

$$\mathrm{E}_t \sum_{T=t}^{\infty} (\beta\alpha_w)^{T-t} \left\{ \lambda_T \frac{a^{T-t}(1+\delta_{t,T}^{w})W_t^{\star}(\upsilon)}{P_T} \ell_{t,T}^{\star}(\upsilon) - \mathrm{e}^{\chi_t} V\left(\ell_{t,T}^{\star}(\upsilon)\right) \right\}.$$

Standard manipulations yield the approximate loglinear relation

$$\hat{\pi}_t^w - \gamma_w \hat{\pi}_{t-1} = \frac{(1-\alpha_w)(1-\beta\alpha_w)}{\alpha_w(1+\omega_w\theta_w)}(\phi\omega_w\hat{y}_t - \hat{\lambda}_t - \hat{w}_t) + \beta\mathrm{E}_t\{\hat{\pi}_{t+1}^w - \gamma_w\hat{\pi}_t\} + \breve{\chi}_t, \tag{16}$$

with

$$\breve{\chi}_t = \frac{(1-\alpha_w)(1-\beta\alpha_w)}{\alpha_w(1+\omega_w\theta_w)}\chi_t.$$

In the above relationship, $\hat{\pi}_t^w$ and $\hat{w}_t$ are the logdeviations of $1+\pi_t^w$ and $w_t\mathrm{e}^{-z_t}$, respectively, and we defined the parameters

$$\omega_w \equiv \frac{V_{\ell\ell}\ell}{V_\ell}, \quad \phi = \frac{F(n)}{F'(n)\,n}.$$

Finally, we assume that

$$\breve{g}_t = \rho_g \breve{g}_{t-1} + \varepsilon_{g,t}, \quad \rho_g \in (0,1), \quad \varepsilon_{g,t} \sim \mathrm{iid}(0,\sigma_g),$$

$$\breve{\chi}_t = \rho_\chi \breve{\chi}_{t-1} + \varepsilon_{\chi,t}, \quad \rho_\chi \in (0,1), \quad \varepsilon_{\chi,t} \sim \mathrm{iid}(0,\sigma_\chi).$$

2.5 Monetary Policy

Monetary authorities are assumed to obey the following interest rate rule

$$\hat{\imath}_t = a_p \mathrm{E}_t\{\hat{\pi}_{t+1}\} + a_y \hat{y}_t + \epsilon_t,$$

where ϵ_t is an monetary shock which evolves according to

$$\epsilon_t = \rho_\epsilon \epsilon_{t-1} + \varepsilon_{\epsilon,t}, \quad \rho_\epsilon \in (0,1), \quad \varepsilon_{\epsilon,t} \sim \mathrm{iid}(0,\sigma_\epsilon).$$

This rule incorporates the usual feedback terms according to which monetary authorities react to the expected logdeviation of inflation as well as the logdeviation of output from its stochastic trend.

A large set of the literature assumes that monetary authorities take their decisions according to an inertial policy rule (e.g. Clarida et al., 2001). In this context, inertia is modelled by assuming that the current nominal interest rate also reacts to its past level. Based on evidence from the term structure of interest rates, Rudebusch (2002) argues convincingly against this specification. The latter would imply forecastability of the nominal interest rate at horizons of more than a quarter, which is not apparent in the data. Instead, we follow Rudebusch (2002) and assume that monetary authorities face persistent shocks in the conduct of their policy[10].

2.6 Equilibrium

In equilibrium, it must be the case that $\hat{d}_t = \hat{y}_t$. The final linear system can then be summarized as follows

$$(1+\beta\bar{b}^2)\hat{y}_t = \bar{b}\hat{y}_{t-1} + \beta\bar{b}\mathrm{E}_t\{\hat{y}_{t+1}\} - (1-\beta\bar{b})(1-\bar{b})\hat{\lambda}_t + \breve{g}_t + \beta\bar{b}\mathrm{E}_t\{\zeta_{t+1}\} - \bar{b}\zeta_t, \tag{17}$$

$$\hat{\lambda}_t = \hat{\imath}_t + \mathrm{E}_t\{\hat{\lambda}_{t+1} - \hat{\pi}_{t+1} - \zeta_{t+1}\}, \tag{18}$$

$$\hat{\pi}_t^w - \gamma_w\hat{\pi}_{t-1} = \frac{(1-\alpha_w)(1-\beta\alpha_w)}{\alpha_w(1+\omega_w\theta_w)}(\phi\omega_w\hat{y}_t - \hat{\lambda}_t - \hat{w}_t) + \beta\mathrm{E}_t\{\hat{\pi}_{t+1}^w - \gamma_w\hat{\pi}_t\} + \breve{\chi}_t, \tag{19}$$

$$\hat{\pi}_t - \gamma_p\hat{\pi}_{t-1} = \frac{(1-\alpha_p)(1-\beta\alpha_p)}{[(1-\mu s_m)^{-1}(1+\theta_p\epsilon_\mu) + \theta_p\omega_p]\alpha_p}(\hat{w}_t + \omega_p\hat{y}_t) + \beta\mathrm{E}_t\{\hat{\pi}_{t+1} - \gamma_p\hat{\pi}_t\}, \tag{20}$$

$$\hat{\pi}_t^w = \hat{\pi}_t + \hat{w}_t - \hat{w}_{t-1} + \zeta_t, \tag{21}$$

$$\hat{\imath}_t = a_p\mathrm{E}_t\{\hat{\pi}_{t+1}\} + a_y\hat{y}_t + \epsilon_t. \tag{22}$$

This system is solved with the AIM package proposed by Anderson and Moore (1985).

3 Model Estimation

In this section, we estimate the above model using the Asymptotic Least Square (ALS) method applied on the autocovariances implied by a canonical VAR. First, we describe the model calibration and second, we introduce the estimation method. Finally, our estimation results are discussed.

[10] In an experiment not reported here, we make sure that our model generates little predictability of the short-term interest rate, as in U.S. data.

3.1 Structural Parameters Calibration

We partition the model parameters into two groups. The first one collects the parameters which we calibrate prior to estimation. These include parameters that can be given a value based on first order moments, as well as parameters that cannot be separately identified. Let $\psi^{\mathrm{c}} = (\beta, \phi, \omega_p, \theta_w, s_m, \theta_p, \epsilon_\mu)'$ denote the vector of calibrated parameters, whose values are reported in table 1. We choose $\beta = 0.99$ as is conventional in the literature for models confronted with quarterly data. Assuming that F is Cobb-Douglas, i.e. $y = n^{1/\phi}$, we set $\phi = 3/2$, implying a labor share close to 66%, as in the data. Notice that we implicitly assume that profits are redistributed proportionately to factors income, so that $1/\phi$ is indeed the steady state labor share, as in Chari et al. (2000). Given that F is Cobb-Douglas, the definition of ω_p implies $\omega_p = \phi - 1$.

Finally, we chose to calibrate s_m, ϵ_μ, θ_p and θ_w because these parameters cannot be separately identified as long as we want to estimate the probabilities of price and wage fixity, namely α_p and α_w. The reason why is simple. Notice that α_p and θ_p (resp. α_w and θ_w) appear only in equation (20) (resp. equation (19)). Fundamentally, the data allow us only to estimate the partial elasticity of inflation (resp. wage inflation) with respect to the real marginal cost (resp. labor disutility wedge), and many combinations of α_p and θ_p (resp. α_w and θ_w) are compatible with a given estimate of this partial elasticity, as explained by Rotemberg and Woodford (1997) and Amato and Laubach (2003). Thus, α_p and θ_p (resp. α_w and θ_w) are not separately identified. Here, we chose to estimate α_p and α_w, which requires that θ_p and θ_w be calibrated prior to estimation. Similar arguments hold for ϵ_μ and s_m, which cannot be separately identified when one wants to estimate α_p.

We set $\theta_w = 21$, as in Christiano et al. (2005) and Schmitt-Grohé and Uribe (2005). We assume that in the deterministic steady state, the markup charged by intermediate goods producers amounts to 20%, implying $\theta_p = 6$, as is conventional in the literature[11]. Following Rotemberg and Woodford (1995), we set $s_m = 0.50$, implying that the share of material goods in gross output is 50%. Under this calibration, the share of materials goods in production costs is $\mu s_m = 0.60$, as proposed by Woodford (2003). Finally, we assume that $\epsilon_\mu = 1$. The rationale for this choice is as follows.

According to Kimball (1995), ϵ_μ could be set at an arbitrarily large value. However, as pointed out by Chari et al. (2000), one is not completely free to choose an arbitrary value for ϵ_μ. Doing so

[11] e.g. Rabanal and Rubio Ramirez (2005), Rotemberg and Woodford (1995).

might result in an implausibly convex demand function. Indeed, a first order Taylor expansion on $\theta_p(\xi)$ near the steady state yields

$$\theta_p(\xi) \approx \theta_p - (1 + \theta_p - \tau)(\xi - 1),$$

where τ governs the curvature of the demand function.[12] Following Kimball (1995), Chari et al. (2000) choose a parameterization in which a 1 percent increase in market share ξ leads to a decline in the elasticity of demand from 10 to 7. They find that the value of τ consistent with this assumption is $\tau = -289$. If we let $\mathbb{D}(P(\varsigma)/P)$ denote the demand function, taking a second order Taylor expansion on $\mathbb{D}(\cdot)$ near the steady state yields

$$\mathbb{D}\left(\frac{P(\varsigma)}{P}\right) \approx 1 - \theta_p\left(\frac{P(\varsigma)}{P} - 1\right) + \frac{\theta_p \tau}{2}\left(\frac{P(\varsigma)}{P} - 1\right)^2,$$

so that, under Chari et al.'s calibration, a 2% increase in relative prices results in a 78% decline in demand, and a 2.3% increase in relative prices results in almost zero demand. They conclude that a demand function with such an extreme level of convexity is implausible.

In our own calibration, assigning values to θ_p and ϵ_μ amounts to assigning a value to τ.[13] Setting $\theta_p = 6$ and $\epsilon_\mu = 1$, implies that $\tau = -23$. Thus, using the above formula, we obtain that a 2% increase in relative prices results in a 14.8% decline in demand. This value is close to what obtains in the case of constant elasticity of demand ($\epsilon_\mu = 0$), in which case a 2% increase in relative prices results in a 11.2% decline in demand.

As discussed above, the degree of strategic complementarity in price setting decisions results from the calibration of the parameters s_m, θ_p and ϵ_μ. It turns out that this modest degree of curvature of the demand function, combined with the assumptions that material goods are used in production, permits us to obtain a small value for $\varkappa$ ($\varkappa = 0.049$). As explained above, setting $\varkappa$ to a small value is important if one is willing to obtain a good empirical fit without relying on an extremely high degree of nominal rigidities.

[12] More precisely, τ is defined as

$$\tau \equiv \frac{G'(1)}{G''(1)} \frac{G'''(1)}{G''(1)}.$$

[13] The exact link between ϵ_μ and τ is

$$\epsilon_\mu = \frac{1 + \theta_p - \tau}{\theta_p(\theta_p - 1)}.$$

3.2 Structural Parameters Estimation

In a first step, we estimate an empirical VAR model. We use U.S. data from the Non Farm Business (NFB) sector over the sample 1965(1)-2002(4). We estimate the VAR model on the following variables

$$\mathbf{y}_t = (\Delta\hat{y}_t \quad \hat{\pi}_t \quad \hat{\imath}_t \quad \hat{\pi}_t^w)',$$

where $\Delta\hat{y}_t$ is output growth, $\hat{\pi}_t$ is the inflation rate, $\hat{\imath}_t$ is the short-term nominal rate, and $\hat{\pi}_t^w$ is the wage inflation rate. We construct $\Delta\hat{y}_t$ as the first difference of the log of real GDP, $\hat{\pi}_t$ as the growth rate of GDP's implicit deflator, $\hat{\pi}_t^w$ as the growth rate of nominal hourly compensation, and $\hat{\imath}_t$ is simply the quarterly Fed Funds rate. The estimated, canonical VAR is of the form

$$\mathbf{y}_t = \Phi_0 + \Phi_1\mathbf{y}_{t-1} + ... + \Phi_p\mathbf{y}_{t-p} + \mathbf{e}_t, \quad \mathbf{e}_t \sim \text{iid}\left(0, \Sigma_e\right), \tag{23}$$

where p is the maximal lag. In our case, we set $p = 4$, as is conventional in the literature. Using the estimated VAR, we can easily compute the empirical autocovariances of $\mathbf{y}_t$. Let Γ_j denote the jth autocovariance of $\mathbf{y}_t$, i.e. $\Gamma_j = \text{E}\{(\mathbf{y}_t - \bar{\mathbf{y}})(\mathbf{y}_{t-j} - \bar{\mathbf{y}})'\}$, where $\bar{\mathbf{y}} \equiv \text{E}\{\mathbf{y}_t\}$. In the sequel, we define

$$\theta = \left(\text{vech}\left(\Gamma_0\right)', \text{vec}\left(\Gamma_1\right)', ..., \text{vec}\left(\Gamma_k\right)'\right)', \qquad k > 0,$$

where the $\text{vec}\left(\cdot\right)$ operator transforms an $(n \times m)$ matrix into an $(nm \times 1)$ vector by stacking the columns of the original matrix, and the $\text{vech}\left(\cdot\right)$ operators transforms an $(n \times n)$ matrix into an $(n(n+1)/2 \times 1)$ vector by vertically stacking those elements on or below the principal diagonal. Let $\hat{\theta}_T$ denote the empirical estimate of θ resulting from the estimated VAR, where T is the sample size. As shown in Lütkepohl (1993)

$$\sqrt{T}(\hat{\theta}_T - \theta) \xrightarrow{d} \text{N}\left(0, \Sigma_\theta\right),$$

where Σ_θ depends on the VAR parameters.

We then seek to estimate the structural parameters defined by

$$\psi = \left(\bar{b}, \gamma_w, \gamma_p, \alpha_p, \alpha_w, a_p, a_y, \rho_g, \sigma_g, \rho_\chi, \sigma_\chi, \rho_\zeta, \sigma_\zeta, \rho_\epsilon, \sigma_\epsilon\right)'.$$

The vector ψ is estimated via the ALS method. Formally, if we let $h\left(\cdot\right)$ denote the mapping from ψ to the DSGE counterpart of θ, the ALS estimate of ψ is then

$$\hat{\psi}_T = \arg\min_{\psi\in\Psi} \ (h(\psi) - \hat{\theta}_T)'W_T(h(\psi) - \hat{\theta}_T),$$

where Ψ is the set of admissible values for ψ and W_T is a positive semi-definite weighting matrix. In practice, W_T is a diagonal matrix with the inverse of the asymptotic variances of each element of $\hat{\theta}_T$ along the diagonal. With this choice for W_T, the parameters are selected so that the model-based autocovariances lie as much as possible inside the confidence interval of their VAR-based counterpart. In addition, we select arbitrarily $k = 10$ in our empirical implementation of the above procedure[14].

Under standard regularity conditions, we have

$$\sqrt{T}(\hat{\psi}_T - \psi) \xrightarrow{d} \mathrm{N}\left(0, D\Sigma_\theta D'\right),$$

where

$$D = \left(\frac{\partial h'}{\partial \psi} W_T \frac{\partial h}{\partial \psi'}\right)^{-1} \frac{\partial h'}{\partial \psi} W_T.$$

In practice, all the partial derivatives are evaluated at the point estimate.

3.3 Estimation Results

First, we estimate our benchmark model with prices and wages both sticky. The estimated parameters as well as the $\mathcal{J}$ statistic are reported in the column entitled M1 of table 2. Second, we decompose our model into two alternatives. We consider a model with only sticky wages and a model with only sticky prices (columns M2 and M3 of table 2, respectively). This latter step will help us to assess the relative merits of sticky wages and sticky prices.

3.3.1 The Model with Price and Wage Both Sticky

We start the discussion of our results with model M1. In a first step, we tried to estimate all the parameters in ψ. Three parameters were characterized by binding constraints, namely $\gamma_w = 1$, $\rho_\zeta = 0$ and $\rho_\chi = 0$. In a second step, we enforced these equalities and estimated the remaining parameters. The first step suggests that the degree of wage indexation to past inflation is very high. This result is consistent with the findings of Giannoni and Woodford (2004) and the assumption of Christiano et al. (2005). Moreover, the zero autocorrelations in the innovation to productivity growth and labor supply shock seem to suggest that the model contains enough internal propagation mechanisms that it does not need to rely on an exogenous channel of persistence.

[14] We also estimated the model with $k = 5$ and $k = 15$. We obtained results which are close to our case $k = 10$.

We then proceed to estimate the remaining parameters. When it comes to the price setting side of the model, we obtain the following results. The probability of no price adjustment is $\alpha_p = 0.704$ and the probability of no wage reoptimization is $\alpha_w = 0.775$. Thus, on average, an intermediate good producer does not reoptimize its price for more than three quarters and a half, and the representative household does not reoptimize its wage for roughly one year. It thus appears that, under reasonable calibrations, our model is able to provide pretty small degrees of nominal price and wage rigidities. Indeed, taking sampling uncertainty into account, our estimate of α_p is consistent with the results reported by Bils and Klenow (2004). Thus, thanks to the strategic complementarity in price setting behaviour, our model is able to respond to the challenge of having a degree of price rigidity compatible with microeconometrics studies.

Thus, our estimated model is characterized by a moderate amount of nominal price rigidities and a higher degree of nominal wage rigidities. In addition, the degree of price indexation to past inflation is $\gamma_p = 0.736$, which implies that during each quarters, fixed prices incorporate roughly 74% of past inflation.

When it comes to the monetary policy rule parameters, we obtain $a_p = 1.451$, suggesting that over the period 1965-2002, the response of monetary authorities to the expected deviations of inflation was rather strong. In addition, the response of monetary authorities to the deviation of output from its stochastic trend is very small and not significant ($a_y = 0.007$). This suggests that in our sample, monetary authorities did not particularly grant attention to this variable. We obtain a pretty high degree of habits in consumption, with $\bar{b}$ almost equal to 0.90. However, this result is not completely surprising given previous estimates in the literature (Boivin and Giannoni, 2003). Finally, ω_w is estimated to a value of 0.606, with little precision.

The standard error of productivity shocks is similar to previous estimates obtained in the literature ($\sigma_\varsigma = 0.943\%$)[15]. The standard error of the preference shock is only 0.274%. The standard errors of the monetary shock is close to 0.40%. The large and significant estimate $\rho_\epsilon = 0.881$ implies that the monetary shock is highly persistent. To a lesser extent, the preference shock is also significantly persistent ($\rho_g = 0.506$).

Figure 1 plots the theoretical and empirical autocovariances (both from model M1 and from the VAR), together with the VAR-based 95% asymptotic confidence interval. This figure shows that the model does a pretty good job of reproducing the cross covariances between the four variables.

[15] Burnside and Eichenbaum (1996); Ireland (2002) for instance.

In particular, the model is able to reproduce the high autocovariance of output growth at short horizons. This has traditionally represented a challenge for studies of business cycle, following the seminal paper by Cogley and Nason (1995). This relative success simply reflects the fact that our model incorporates a sufficient number of internal sources of propagation. The autocovariances of inflation, the Fed Funds rate, and wage inflation are also fairly well replicated. The model also performs reasonably well when it comes to the cross-covariances between lagged output growth and inflation, wage inflation and the interest rate. In addition, notice that despite the fact that the estimation algorithm drove a_y to zero, the model is also able to reproduce the correlation pattern between the nominal interest rate and output growth. In particular, the model is able to replicate the inverted-leading indicator property of the nominal interest rate. Overall, the theoretical autocovariances are well within the confidence intervals of their VAR-based counterparts.

3.3.2 The Sticky Wages Model

In this second step, we propose to deconstruct the relative importance of sticky wages in explaining the results obtained in the previous section. To do so, we re-estimate the benchmark model assuming that prices are fully flexible ($\alpha_p = \gamma_p = 0$). The column entitled M2 in table 2 reports the values of the estimated parameters when we impose these constraints. As previously, we were confronted with binding constraints. In particular the estimation algorithm drives ω_w to an implausibly large value. We thus decide to calibrate $\omega_w = 3$, which seems to correspond to a upper bound for this elasticity. Under this specification, the estimation algorithm drives the probability of not reoptimizing wages to a lower value than in model M1 ($\alpha_w = 0.575$), implying that, on average the representative household does not reoptimize its wage for slightly more than two quarters. Thus, at face value, model M2 seems to be characterized by a smaller overall degree of nominal rigidities than model M1. Moreover, wages are no longer indexed to past inflation since γ_w is driven to zero in the estimation stage. This implies that, in each quarter, fixed wages increase at a constant rate. However, closer scrutiny seems to suggest that the persistence channels have simply changed. In particular, the habit parameter is higher than in model M1 ($\bar{b} = 0.93$). Thus, the fully flexible prices assumption seems to need much more consumption smoothing in order to replicate the VAR-based autocovariances. Similarly, the innovation of the productivity shock becomes persistent ($\rho_\zeta = 0.807$) with a standard error of 0.178%, which suggests that the model suffers from a lack of internal propagation mechanisms alleviated by exogenous sources of persistence. A possible interpretation is that this exogenous persistence mechanism has substituted for nominal price rigidities. Notice also that under this assumption, the labor supply shock is no longer necessary to replicate the autocovariances of U.S. data, since the estimation algorithm drives σ_χ to zero. Lastly, the preference shock is not significantly persistent and its standard error is higher than in the benchmark model ($\sigma_g = 0.37\%$).

In addition, the monetary policy rule is also affected by this specification since $a_p = 1.369$, which implies that monetary authorities are mildly less reactive to variations in expected inflation. However, the monetary shock remains quite persistent ($\rho_\epsilon = 0.831$) and the standard error is close to 0.19%.

Figure 1 highlights the performances of model M2. It clearly appears that the model with only sticky wages is less successful in replicating the autocovariances of the four variables than model M1. In particular, the model is not able to mimic the autocovariance of output growth. In the same fashion, the VAR-based cross covariances between output growth and inflation are overestimated by model M2. Although the model-based autocovariances of inflation, the interest rate and wage inflation are within the VAR-based confidence interval, the fit is not very satisfying. This is evidence that wage stickiness is not sufficient in itself to fit the autocovariances between wage inflation and the other variables.

3.3.3 The Sticky Prices Model

We now alter our model by assuming that wages are fully flexible while prices are sticky ($\alpha_w = \gamma_w = 0$). The column entitled M3 in table 2 reports the estimated parameters in this case. First, the degree of price rigidities remains almost constant, compared to the benchmark model M1. Indeed, the probability of no price reoptimization (α_p) varies from 0.704 to 0.682. In addition, the degree of price indexation to past inflation increases compared to model M1 ($\gamma_p = 0.797$). The habit parameter is also close to the estimate obtained in the benchmark model since $\bar{b}$ now equals 0.886. Finally, in terms of the interest rate rule, the response of the monetary authorities to a variation in expected inflation is close to the value obtained by Clarida et al. (2001) ($a_p = 1.26$), whereas a_y is driven to zero, thus slightly smaller than in the benchmark model.

When it comes to the stochastic shocks parameters, the model's dynamics are different from what obtains in model M1. The monetary shock does not seem to be necessary in order to replicate the autocovariances of the four variables, since the estimation algorithm drives σ_ϵ to zero. On the contrary, the model requires a large amount of serial correlation in the labor supply shock. Indeed, during the estimation stage, ρ_χ was driven toward its upper bound. Accordingly, we reestimated the model with the imposed constraint $\rho_\chi = 0.99$, so as to avoid non-stationarity issues. Finally, the productivity shock and the preference shock are not much altered compared to the benchmark model.

Figure 1 shows that model M3 replicates approximately the autocovariances of the variables. More precisely, the VAR-based autocovariances of output growth and the interest rate are close to the one provided by the benchmark model. The variance of wage inflation is overestimated. A possible reason is that the real wage is too volatile compared to price inflation, thus translating into a volatile wage inflation. Assuming that wages are fully flexible seems to also affect the covariance between lagged output growth and wage inflation which is overestimated. In addition, in a model with only sticky prices, the cross covariances between inflation and output growth is rather well replicated.

4 Assessing the Model's Fit

To assess formally the models' fit, we now resort to Watson's (1993) procedure. The latter is not a test based on a null hypothesis but rather offers a simple measure of fit that allows us to provide a quantitative assessment of the model's ability to replicate the dynamics of the data.

4.1 Watson's Procedure

Watson's (1993) procedure consists in decomposing the performances of a model into the frequency domain. The procedure amounts to augmenting the data generated by the model, $\mathbf{x}_t$, with an approximation error $\mathbf{u}_t$ designed to reconcile the second order moments of the model with those from the data, $\mathbf{y}_t$. If the added error is small, then the model is judged to do a good job of accounting for these moments. Here, we will study the spectral properties of the process $\mathbf{y}_t$, either taken in level or Hodrick-Prescott (HP) filtered.[16]

Formally, the procedure underlying Watson's (1993) procedure might be described as follows. First, the error induced by the model is defined as the difference between the two data sets, i.e. $\mathbf{u}_t = \mathbf{y}_t - \mathbf{x}_t$. Assuming that $\mathbf{y}_t$ and $\mathbf{x}_t$ are jointly stationary, we can define the spectral density matrix

[16] As is customary with quarterly data, we set the smoothing parameter to 1600.

of $\mathbf{u}_t$ at frequency ω by the formula

$$A_{\mathbf{u}}(\mathrm{e}^{-\mathrm{i}\omega}) = A_{\mathbf{y}}(\mathrm{e}^{-\mathrm{i}\omega}) + A_{\mathbf{x}}(\mathrm{e}^{-\mathrm{i}\omega}) - A_{\mathbf{xy}}(\mathrm{e}^{-\mathrm{i}\omega}) - A_{\mathbf{xy}}(\mathrm{e}^{-\mathrm{i}\omega})',$$

where a prime denotes the transpose-conjugate operation and $A_{\mathbf{xy}}(\mathrm{e}^{-\mathrm{i}\omega})$ is the cross-spectrum of the model and data.

To compute $A_{\mathbf{u}}(\mathrm{e}^{-\mathrm{i}\omega})$ three terms are needed. First, $A_{\mathbf{y}}(\mathrm{e}^{-\mathrm{i}\omega})$, the spectral density of $\mathbf{y}_t$, is built from the VAR previously estimated. Second, $A_{\mathbf{x}}(\mathrm{e}^{-\mathrm{i}\omega})$, the spectral density of $\mathbf{x}_t$, is computed from the approximate solution to the DSGE model. Third, we need the cross spectral density $A_{\mathbf{xy}}(\mathrm{e}^{-\mathrm{i}\omega})$. However, this quantity cannot be determined by the model, nor can it be estimated from the data. Watson (1993) points out that any restriction used to identify $A_{\mathbf{xy}}(\mathrm{e}^{-\mathrm{i}\omega})$, and hence $A_{\mathbf{u}}(\mathrm{e}^{-\mathrm{i}\omega})$, is arbitrary. However, he proposes a way of computing a lower bound for the variance of $\mathbf{u}_t$ by selecting $A_{\mathbf{xy}}(\mathrm{e}^{-\mathrm{i}\omega})$ so as to minimize $\mathrm{tr}(\Omega A_{\mathbf{u}}(\mathrm{e}^{-\mathrm{i}\omega}))$, subject to the requirement that the spectral density matrix of $(\mathbf{x}_t', \mathbf{y}_t')'$ be positive semidefinite at all frequencies. Here, $\mathrm{tr}(\cdot)$ is the trace operator and Ω is a pre-specified weighting matrix.

For each frequency, we can determine a lower bound of the variance of the approximation error divided by the variance of the data. Let $r(\omega)$ denote this bound and $r_j(\omega)$ denote the jth component of $r(\omega)$. In the same fashion, let $[A_{\mathbf{u}}(\mathrm{e}^{-\mathrm{i}\omega})]_{jj}$ and $[A_{\mathbf{y}}(\mathrm{e}^{-\mathrm{i}\omega})]_{jj}$ denote the (j,j) elements of matrices $A_{\mathbf{u}}\left(\mathrm{e}^{-\mathrm{i}\omega}\right)$ and $A_{\mathbf{y}}\left(\mathrm{e}^{-\mathrm{i}\omega}\right)$, respectively. We can then define

$$r_j(\omega) = \frac{[A_{\mathbf{u}}(\mathrm{e}^{-\mathrm{i}\omega})]_{jj}}{[A_{\mathbf{y}}(\mathrm{e}^{-\mathrm{i}\omega})]_{jj}}.$$

Watson (1993) proposes to integrate separately both the numerator and denominator of the above expression, defining so the Relative Mean Square Approximation Error (RMSAE) induced by the model. The smaller it is, the better the model reproduces the spectral behavior of the data. This statistic is nothing more than the variance of the error relative to the variance of the data, decomposed in the frequency domain.

4.2 Results of Watson's (1993) procedure

Table 3 reports the estimated RMSAEs with their standard errors. The standard errors of the RMSAEs are based on the sampling error in the estimated VAR coefficients used to estimate the data spectrum. The latter are based on the sampling uncertainty in the estimated VAR coefficients used to estimate the data spectrum. As before, the columns labelled M1, M2, and M3 correspond to the sticky prices – sticky wages, sticky wages, and sticky prices models, respectively. Watson's procedure (1993) is not a test based on a null hypothesis but rather offers a simple measure of fit of the three models. Thus, we decide to base our discussion about the goodness-of-fit of the models both on the value and the significance of the RMSAEs. More particularly, we will argue that a model does a good job of fitting the data when the RMSAEs are small and when all (or almost all) RMSAEs are not significant.

Panel A reports the RMSAEs over the frequency range $[0, \pi]$ so as to get a feel for the overall models' behavior. However, as has been often acknowledged in the literature, DSGE models are not necessarily meant to account for all the dynamic movements in the data. Naturally, in this paper, our primary focus is on the business cycle. Thus, in a second step, we restrict our attention to the frequency band $[\pi/16, \pi/3]$. This interval insulates the frequencies typically attached to business cycle, i.e. cyclical movements the reproduction period of which runs from 6 to 32 quarters. The results of this second step are included under the heading B. Additionally, we complete this study by considering HP-filtered variables, which are reported under the heading C. Given that this filter is widely used as an operational definition of the business cycle, it seems natural to include it in our empirical analysis.

Moreover, in the top panel of table 3, results are obtained under the assumption that all variables are given the same weight in Watson's procedure, i.e. Ω is the identity matrix. The bottom panel of the table resorts to a diagonal matrix Ω containing the VAR-based variances of $\mathbf{y}_t$ along its diagonal. This alternative weighting scheme permits us to downsize the importance of those elements of $\mathbf{y}_t$ that have a high variance relative to the others.

4.2.1 The Model with Price and Wage Both Sticky

The results of Watson's procedure applied to our benchmark model M1 confirm our previous conclusion, i.e. the model does a very good job of fitting the main characteristics of post-war U.S. data. With the first weighting matrix, when evaluated over the whole frequency range, all RMSAEs are small and only the RMSAE associated with output growth is significant, which reinforces our conclusion about the goodness-of-fit of our benchmark model. Since we assess a business cycle model, we focus our analysis on business cycle frequencies. In this case, the RMSAEs for inflation, wage inflation, and the Fed Funds rate increase but it remains very small when compared, for example, with what Jung (2004) obtains. In addition, the errors are not significant, except for inflation. These results are very encouraging about the quality of fit of our benchmark model. When we focus on HP filtered variables, the errors for output growth and inflation are significant, but the RMSAEs are close to the case over the business cycle frequencies. Still, the model performs well when the sampling error is taken into account.

Figure 2 provides a graphical interpretation of these results, under the minimum error representation with equal weight. As expected, the model is able to reproduce the peak in the spectrum of output growth at business cycle frequencies. Notice that for output growth, the error is mainly concentrated at low or high frequencies. The spectrum of $\mathbf{u}_t$ slightly peaks at business cycle frequencies for inflation and the Fed Funds rate. Finally, the variance of the error is mainly concentrated at low frequencies for wage inflation.

4.2.2 The Sticky Wages Model

Figure 3 gives us the results of Watson's procedure for model M2. In this case, the error needed to match the VAR-based and model-based spectra for all variables is much higher than in model M1. In particular, as was to be expected from figure 1, model M2 fails to replicate the VAR-based spectrum of output growth.

Table 3 confirms this result since the errors appear to be significant in almost all the cases considered. In particular the RMSAEs are strongly affected by the assumption that prices are fully flexible. Over the whole frequency range, the error is significant for output growth (18%), wage inflation (34%) and inflation (49%). Model M2 does also a poor job of fitting the data when the focus is on business cycle frequencies. More precisely, the variance of the error needed to match the VAR-based and model-based spectra is greater than 45% for all variables, except for output growth (26%), and it even reaches 108% for inflation. When it comes to HP-filtered variables, the results are even worse, since the RMSAEs for HP-filtered variables are all large and significant. This is especially true for inflation which requires an error whose variance is at least 160% of its empirical counterpart. To sum up, since most of the RMSAEs are large and significant, we can conclude that Watson's procedure is not overwhelmingly supportive of model M2.

4.2.3 The Sticky Prices Model

Figure 4 gives us the results of Watson's procedure for model M3. It appears that the match between the VAR-based and the model-based spectra is not as good as in model M1. In particular, model M3 seems to suffer from a lack of amplification properties, since the spectrum of output growth is much lower than its empirical counterpart at all frequencies. The variance of the error (relative to that of the data) for the other variables is large, particularly at low frequencies.

Columns M3 of table 3 report the estimated RMSAEs when wages are perfectly flexible. Over the whole frequency range, the RMSAEs are larger than for model M1. However, notice that none is significant. This result contrasts with the findings of Ellisson and Scott (2000), who showed that their sticky price model obtains a very poor fit (especially at high frequencies). Compared to the benchmark model, the RMSAE rises from 13% to 39% for inflation, from 10% to 38% for the interest rate, and from 4% to 40% for wage inflation. Although the RMSAEs are all not significant over the whole frequency range, the results are less successful at business cycle frequencies. Indeed, the RMSAEs of inflation, wage inflation, and the Fed Funds rate are significant. Finally, when it comes to the HP-filtered variables, the RMSAEs are overall significant.

Thus, the sticky price model seems to do a relatively poor job compared to the benchmark model. Although it is not completely wrong according to Watson's (1993) criterion over the whole frequency range – contrary to the sticky wage model, the model does not perform well on business cycle frequencies. This comparison confirms that prices and wages both sticky are a necessary ingredient in order to reproduce the empirical autocovariances of our four variables.

5 Conclusion

The present paper proposed to provide a systematic assessment of the goodness-of-fit of a small-scale, typical sticky prices – sticky wages New Keynesian model, similar to those currently in use in the literature. The model incorporates a large number of modelling ingredients that have been shown to improve the fit of New Keynesian models: habit formations, material goods, a variable elasticity of demand for goods, as well as prices and wages indexation schemes. The model was confronted to the data with a two-step procedure. First, the structural parameters were estimated so as to reproduce a truncated version of the autocovariances of key macroeconomic variables, as implied by a simple canonical VAR. Second, the overall goodness-of-fit of the model was assessed by resorting to Watson's (1993) procedure.

The empirical results in this paper highlight the ability of the sticky prices – sticky wages model to replicate the second order moments of postwar U.S data. It appears that the New Keynesian model is able to reproduce satisfactorily the empirical spectra of output growth, inflation, the short-term nominal interest rate, and wage inflation, according to Watson's (1993) criterion. Our main result is that the combination of sticky prices and sticky wages is necessary in order to obtain a good empirical fit. An original result which our analysis reveals is that a model with only sticky wages does a very poor job of fitting the data, according to Watson's (1993) criterion. This result contrasts with those previously obtained in the literature, which all suggest that wages stickiness implies more persistence effects than prices stickiness (Andersen, 1998; Huang and Liu, 2002; Christiano et al., 2005, for instance). However, contrary to these authors, we focus on the overall persistence rather than on the sole persistence conditional on monetary shocks.

However, our results suggest that prices and wages stickiness is central for New Keynesian models to provide a reasonable explanation of U.S. postwar business. This result reinforces the importance of the conclusion reached by Erceg et al. (2000) who showed that in such an environment, it is no longer feasible for monetary authorities to completely stabilize price inflation.

To conclude, one must notice that in spite of its good performances, the model still needs improvements especially when it comes to reproducing the dynamics of inflation and the nominal interest rate at business cycle frequencies. Concerning the Federal Funds rate, these shortcomings might be alleviated by modifying the monetary policy rule. As to inflation a possible solution could consist in including further lags of inflation in the so-called New Keynesian Phillips curve, as suggested by Kiley (2005). We leave these for future research.

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Table 1. Calibrated Parameters

Parameters	Value
β	0.99
ϕ	1.50
ω_p	0.50
s_m	0.50
θ_p	6.00
ϵ_μ	1.00
θ_w	21.00

Table 2.Results of ALS estimation

	Models		
Parameters	M1	M2	M3
α_p	0.704 (0.045)	—	0.682 (0.045)
α_w	0.775 (0.087)	0.575 (0.092)	—
γ_p	0.736 (0.091)	—	0.797 (0.147)
γ_w	1.000 (*)	0.000 (*)	—
$\bar{b}$	0.891 (0.023)	0.930 (0.016)	0.886 (0.034)
ω_w	0.606 (0.703)	3.000 (*)	0.260 (0.212)
a_y	0.007 (0.079)	0.000 (*)	0.000 (*)
a_p	1.451 (0.750)	1.369 (0.520)	1.260 (0.157)
ρ_ζ	0.000 (*)	0.807 (0.074)	0.000 (*)
σ_ζ	0.943 (0.291)	0.178 (0.041)	0.624 (0.118)
ρ_g	0.506 (0.052)	0.090 (0.097)	0.442 (0.076)
σ_g	0.274 (0.037)	0.374 (0.031)	0.323 (0.045)
ρ_χ	0.000 (*)	0.000 (*)	0.990 (*)
σ_χ	0.531 (0.056)	0.000 (*)	1.811 (0.685)
ρ_ϵ	0.881 (0.052)	0.831 (0.061)	0.000 (*)
σ_ϵ	0.402 (0.108)	0.495 (0.141)	0.000 (*)

Notes: Models codes as follows. M1: benchmark model; M2: sticky wages model; M3: sticky prices model. The number in parentheses are the standard errors of the parameters. A star refers to a constraint imposed during the estimation stage to avoid convergence issues.

Table 3. Results of Watson procedure (RMSAE)

Variables	Definition of Data Spectrum and Range of Evaluation								
	A. All Frequencies			B. 6–32 Quarters			C . HP, All Frequencies		
	Equal Weights, $\Omega = \Omega_1$								
	M1	M2	M3	M1	M2	M3	M1	M2	M3
Output Growth	0.053† (0.024)	0.178† (0.043)	0.048 (0.025)	0.013 (0.019)	0.257† (0.064)	0.014 (0.020)	0.044† (0.019)	0.158† (0.037)	0.032† (0.016)
Inflation	0.134 (0.085)	0.490† (0.231)	0.390 (0.207)	0.223† (0.100)	1.084† (0.441)	0.407† (0.106)	0.411† (0.073)	1.608† (0.355)	0.582† (0.082)
Interest Rate	0.094 (0.086)	0.168 (0.105)	0.384 (0.296)	0.309 (0.242)	0.474 (0.304)	0.214† (0.099)	0.292 (0.160)	0.551† (0.224)	0.312† (0.073)
Wage Inflation	0.037 (0.057)	0.339† (0.108)	0.405 (0.214)	0.030 (0.066)	0.490† (0.150)	0.256† (0.130)	0.044 (0.034)	0.672† (0.077)	0.593† (0.072)
	Unequal Weights, $\Omega = \Omega_2$								
	M1	M2	M3	M1	M2	M3	M1	M2	M3
Output Growth	0.069† (0.027)	0.276† (0.048)	0.066† (0.026)	0.030 (0.028)	0.370† (0.078)	0.024 (0.024)	0.056† (0.021)	0.257† (0.045)	0.047† (0.019)
Inflation	0.128 (0.084)	0.384† (0.190)	0.364 (0.192)	0.208† (0.099)	0.833† (0.414)	0.374† (0.101)	0.401† (0.075)	1.260† (0.305)	0.557† (0.079)
Interest Rate	0.088 (0.083)	0.171 (0.106)	0.412 (0.316)	0.278 (0.235)	0.455 (0.301)	0.226† (0.094)	0.263 (0.155)	0.552† (0.226)	0.324† (0.071)
Wage Inflation	0.028 (0.056)	0.309† (0.099)	0.374 (0.199)	0.026 (0.066)	0.417† (0.131)	0.247 (0.127)	0.033 (0.031)	0.624† (0.068)	0.559† (0.065)

Notes: Models codes as in table 2. The particular weighting matrices Ω used to minimize $\mathrm{tr}(\Omega A_{\mathbf{u}})$, are Ω_1: identity matrix; Ω_2: matrix containing the inverse of the VAR-based variances of $\mathbf{y}_t$ along its diagonal. The number in parentheses are standard errors based on the sampling error in the estimated VAR coefficients used to estimate the data spectrum. The † indicates that the RMSAE is significant at the 5% level.

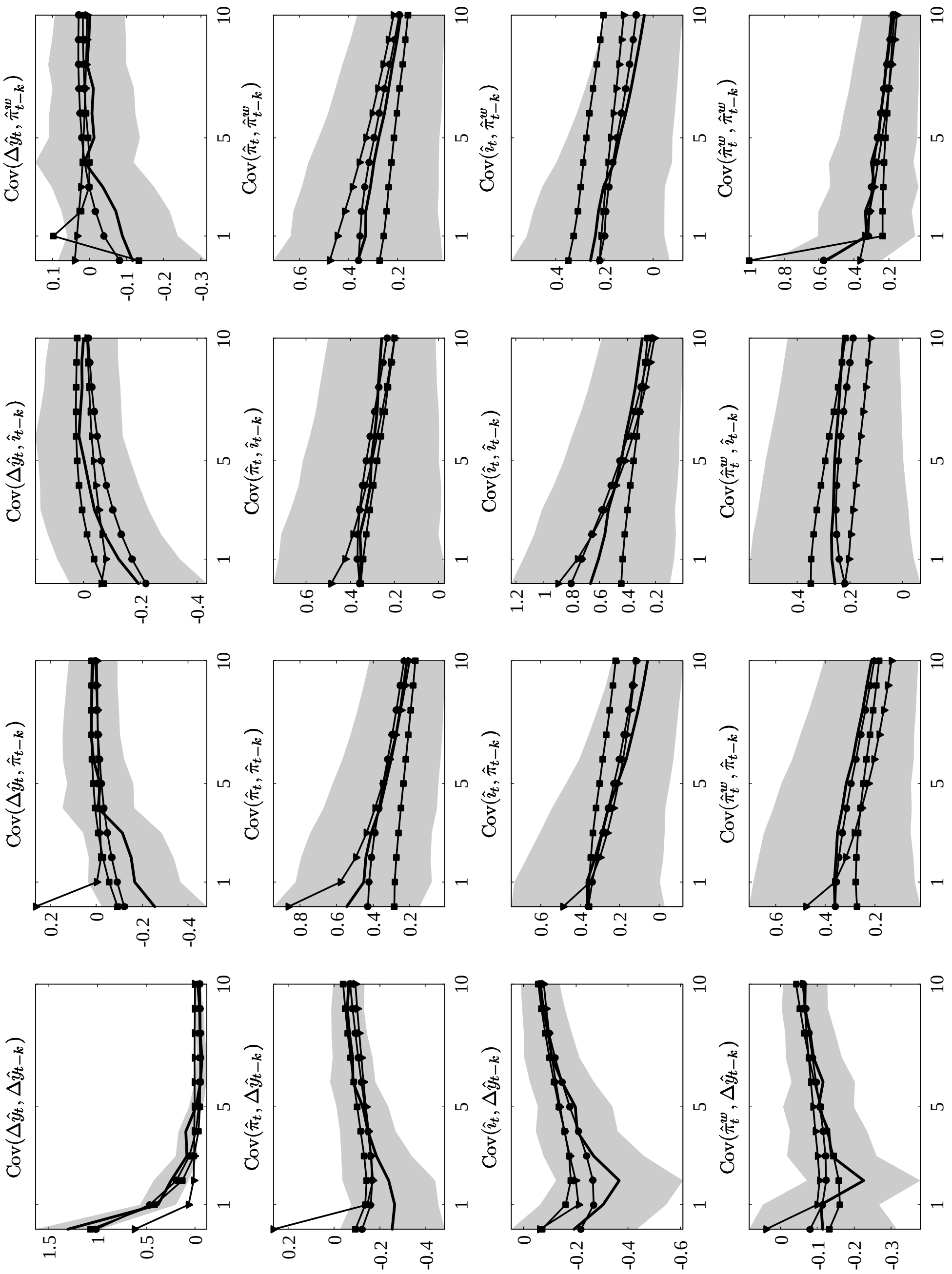

Figure 1: SVAR-Based and Model-Based Autocovariances. Plain line: VAR model; Line with circles: model M1; Line with triangles: model M2; Line with squares: model M3.

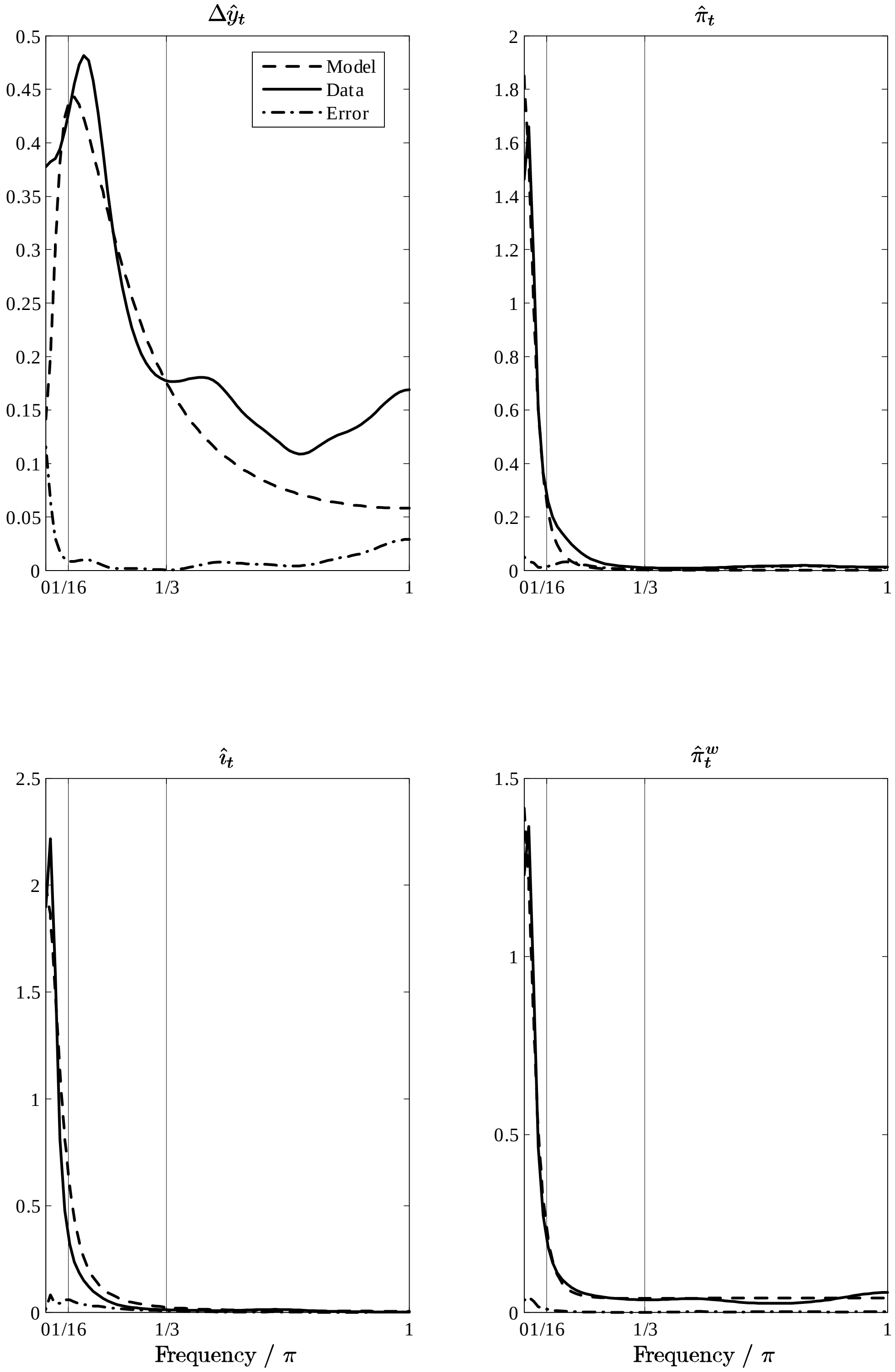

Figure 2: Model M1 spectral densities (dashed line), VAR-based spectral densities (solid line), and error spectral densities (dash-dot line) for output growth, inflation, short term nominal interest rate, and wage inflation. The data have been multiplied by 100. The error is computed under the minimum error representation with equal weight.

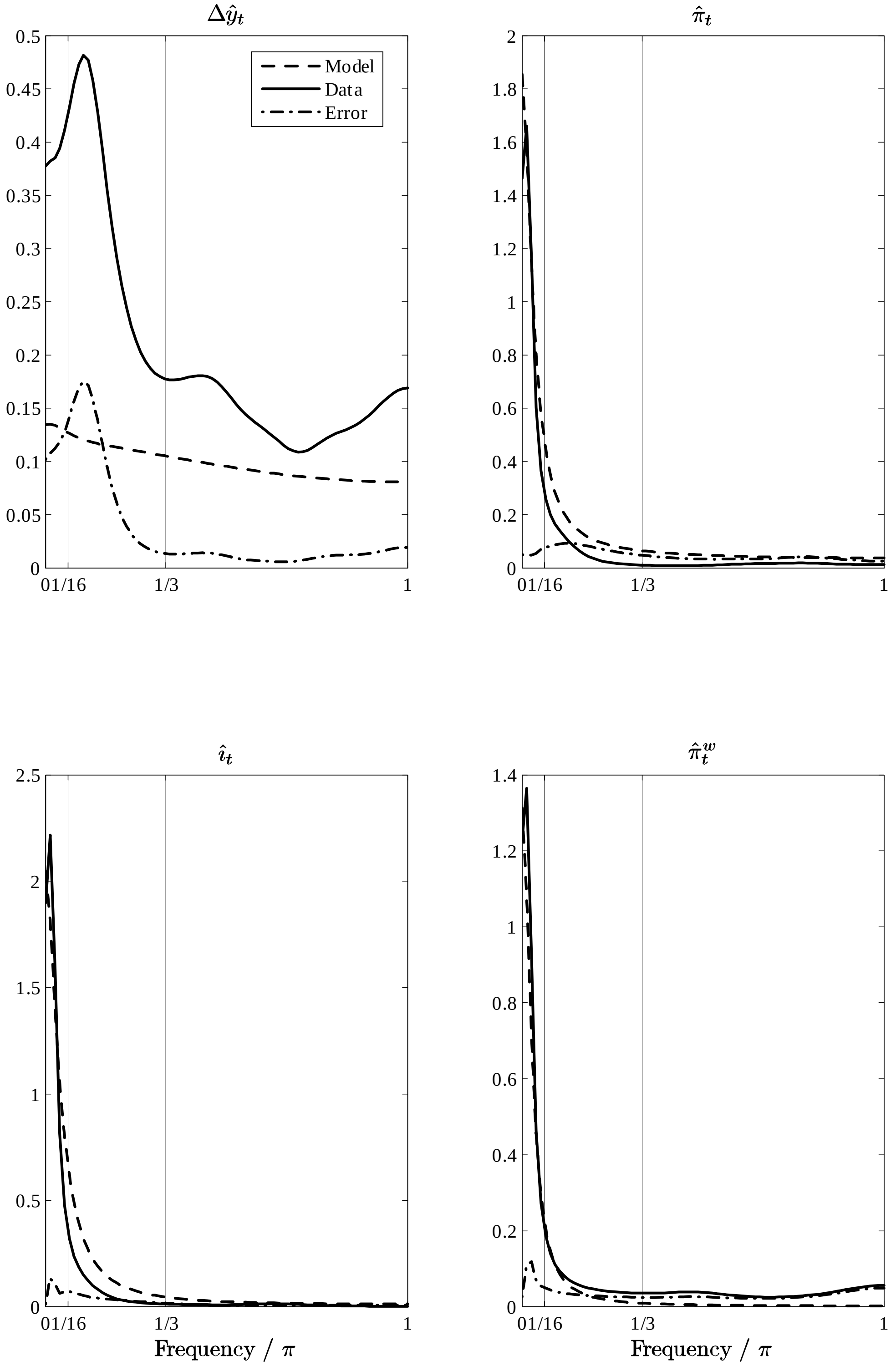

Figure 3: Model M2 spectral densities (dashed line), VAR-based spectral densities (solid line), and error spectral densities (dash-dot line) for output growth, inflation, short term nominal interest rate, and wage inflation. The data have been multiplied by 100. The error is computed under the minimum error representation with equal weight.

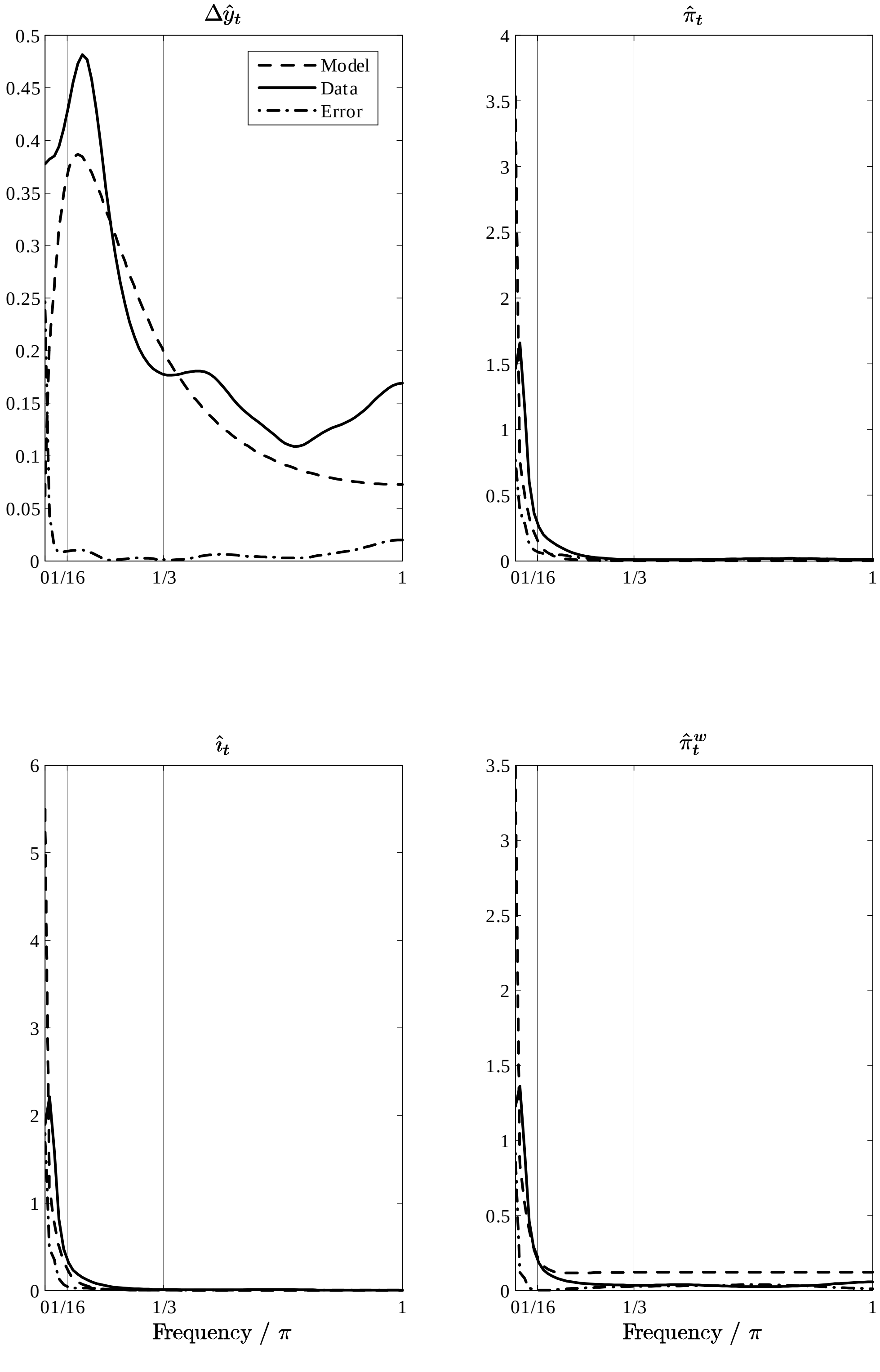

Figure 4: Model M3 spectral densities (dashed line), VAR-based spectral densities (solid line), and error spectral densities (dash-dot line) for output growth, inflation, short term nominal interest rate, and wage inflation. The data have been multiplied by 100. The error is computed under the minimum error representation with equal weight.

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