**FUNDAÇÃO GETULIO VARGAS**
**ESCOLA de PÓS-GRADUAÇÃO em ECONOMIA**

**Alexis Maka**

# On Testing the Phillips Curves, the IS Curves, and the Interaction between Fiscal and Monetary Policies

**Rio de Janeiro**
**2013**

Alexis Maka

# On Testing the Phillips Curves, the IS Curves, and the Interaction between Fiscal and Monetary Policies

Tese submetida à Escola de Pós-Graduação em Economia da Fundação Getulio Vargas como requisito parcial para a obtenção do grau de Doutor em Economia

Orientador: Fernando de Holanda Barbosa

Rio de Janeiro
2013

Maka, Alexis
On testing the Phillips curves, the IS curves, and the interaction between fiscal and monetary policies / Alexis Maka. – 2013.
95 f.

Tese (doutorado) - Fundação Getulio Vargas, Escola de Pós-Graduação em Economia.
Orientador: Fernando de Holanda Barbosa.
Inclui bibliografia.

1. Phillips, Curva de. 2. Política tributária. 3. Polítca monetária. 4. Modelos econométricos. I. Barbosa, Fernando de Holanda. II. Fundação Getulio Vargas. Escola de Pós- Graduação em Economia. III. Título.

CDD – 332.46

FUNDAÇÃO
GETULIO VARGAS

**ALEXIS MAKA**

**"ON TESTING THE PHILLIPS CURVES, THE IS CURVES, AND THE INTERACTION BETWEEN FISCAL AND MONETARY POLICIES"**

Tese apresentada ao Curso de Doutorado em Economia da Escola de Pós-Graduação em Economia para obtenção do grau de Doutor em Economia.

Data da defesa: 27/11/2013

Aprovada em: 27/11/2013

**ASSINATURA DOS MEMBROS DA BANCA EXAMINADORA**

**Fernando de Holanda Barbosa**
Orientador (a)

**Renato Fragelli Cardoso**

**João Victor Issler**

**Elcyon Caiado Rocha Lima**

**André Minella**

# Acknowledgements

The process of writing and completing this dissertation has left me indebted to several people. Without their help and support this work would never have seen the light of day. I would like to take this opportunity to express my deepest gratitute towards them.

My main debt of gratitude goes to Fernando de Holanda Barbosa for his superb advice, constant dedication, and unfaltering encouragement.

I wish to thank Elcyon Caiado Rocha Lima for carefully reading my dissertation manuscript and providing me with many insightful and encouraging suggestions. I would also like to thank Renato Fragelli Cardoso, João Victor Issler, and André Minella for providing me with several detailed and valuable comments. Their help has greatly improved the quality of the dissertation.

I gratefully acknowledge the financial support from CAPES.

Finally, I would like to thank my beautiful wife and muse Katia for sharing life with me and making all this worthwhile.

# Abstract


This dissertation consists of three essays on empirical testing of Phillips curves, IS curves, and the interaction between fiscal and monetary policies.

The first essay ("Phillips Curves: An Encompassing Test") tests Phillips curves using an autoregressive distributed lag (ADL) specification that encompasses the accelerationist Phillips curve (APC), the New Keynesian Phillips curve (NKPC), the Hybrid Phillips curve (HPC), and the Sticky-Information Phillips curve (SIPC). We use data from the United States (1985Q1–2007Q4) and from Brazil (1996Q1–2012Q2), using the output gap and alternatively the real marginal cost as measure of inflationary pressure. The empirical evidence rejects the restrictions implied by the NKPC, the HPC, and SIPC, but does not reject those implied by the APC.

The second essay ("IS Curves: An Encompassing Test") tests IS curves using an ADL specification that encompasses the traditional Keynesian IS curve (KISC), the New Keynesian IS curve (NKISC), and the Hybrid IS curve (HISC). We use data from the United States (1985Q1–2007Q4) and from Brazil (1996Q1–2012Q2). The evidence rejects the restrictions implied by the NKISC and the HISC, but does not reject those of the KISC.

The third essay ("The Effects of Fiscal Policy and its Interactions with Monetary Policy in Brazil") analyzes the effects of fiscal policy shocks on the dynamics of the economy and the interaction between fiscal and monetary policy using structural vector autoregressions (SVARs). We test the Fiscal Theory of the Price Level for Brazil, analyzing the response of public sector liabilities to primary surplus shocks. For the hybrid identification we find that it is not possible to distinguish empirically between Ricardian (Monetary Dominance) and non-Ricardian (Fiscal Dominance) regimes. However, using sign restrictions there is some evidence that the government followed a Ricardian (Monetary Dominance) regime from January 2000 to June 2008.



**Keywords:** Phillips curve; IS curve; encompassing test; fiscal policy shocks.

# Resumo

Esta tese é composta por três ensaios sobre testes empíricos de curvas de Phillips, curvas IS e a interação entre as políticas fiscal e monetária.

O primeiro ensaio ("Curvas de Phillips: um Teste Abrangente") testa curvas de Phillips usando uma especificação autoregressiva de defasagem distribuída (ADL) que abrange a curva de Phillips Aceleracionista (APC), a curva de Phillips Novo Keynesiana (NKPC), a curva de Phillips Híbrida (HPC) e a curva de Phillips de Informação Rígida (SIPC). Utilizamos dados dos Estados Unidos (1985Q1–2007Q4) e do Brasil (1996Q1–2012Q2), usando o hiato do produto e alternativamente o custo marginal real como medida de pressão inflacionária. A evidência empírica rejeita as restrições decorrentes da NKPC, da HPC e da SIPC, mas não rejeita aquelas da APC.

O segundo ensaio ("Curvas IS: um Teste Abrangente") testa curvas IS usando uma especificação ADL que abrange a curva IS Keynesiana tradicional (KISC), a curva IS Novo Keynesiana (NKISC) e a curva IS Híbrida (HISC). Utilizamos dados dos Estados Unidos (1985Q1–2007Q4) e do Brasil (1996Q1–2012Q2). A evidência empírica rejeita as restrições decorrentes da NKISC e da HISC, mas não rejeita aquelas da KISC.

O terceiro ensaio ("Os Efeitos da Política Fiscal e suas Interações com a Política Monetária") analisa os efeitos de choques na política fiscal sobre a dinâmica da economia e a interação entre as políticas fiscal e monetária usando modelos SVARs. Testamos a Teoria Fiscal do Nível de Preços para o Brasil analisando a resposta do passivo do setor público a choques no superávit primário. Para a identificação híbrida, encontramos que não é possível distinguir empiricamente entre os regimes Ricardiano (Dominância Monetária) e não-Ricardiano (Dominância Fiscal). Entretanto, utilizando a identificação de restrições de sinais, existe evidência que o governo seguiu um regime Ricardiano (Dominância Monetária) de janeiro de 2000 a junho de 2008.

**Palavras-chave:** curva de Phillips; curva IS; teste abrangente; choques de política fiscal.

# Contents

# List of Figures

# List of Tables

# Chapter 1

# Introduction

In the 1970s, 1980s, and early 1990s, models used for monetary policy analysis combined the assumption of nominal rigidity with a simple structure that linked the quantity of money to aggregate spending. Although the theoretical foundations of these models were weak, the approach proved remarkably useful in addressing a wide range of monetary policy topics. Today, the standard approach in monetary economics and monetary policy analysis incorporates nominal wage or price rigidity into a dynamic stochastic general equilibrium (DSGE) framework that is based on optimizing behavior by the agents in the model.

These modern DSGE models with nominal frictions are commonly labeled New Keynesian models because, like older versions of models in the Keynesian tradition, aggregate demand plays a central role in determining output in the short run, and there is a presumption that some fluctuations both can be and should be dampened by countercylical monetary.

In its simplest form the canonical New Keynesian model is composed by: (1) an inflation equation, which resembles an expectations-augmented Phillips curve [the New Keynesian Phillips curve]; (2) Euler equation for output, which generalizes the consumption Euler equation to the whole economy [the New Keynesian IS curve] ; (3) an equation describing how interest rate policy is set, usually described as an explicit interest rate rule.

The New Keynesian model became the workhorse model for monetary analysis.[1] However, to use this model for policy analysis requires it to have a good fit to the data. The first two essays of this dissertation are concerned with the econometric analysis of the New Keynesian model.

The first essay ("Phillips Curves: An Encompassing Test") tests Phillips curves using an autoregressive distributed lag (ADL) specification that encompasses the Accelerationist Phillips curve (APC), the New Keynesian Phillips curve (NKPC), the Hybrid Phillips curve (HPC), and the Sticky-Information Phillips curve (SIPC). We use data from the United States (1985Q1–2007Q4) and from Brazil (1996Q1–2012Q2), using the output gap and alternatively the real marginal cost as measure of inflationary pressure.

The second essay ("IS Curves: An Encompassing Test") tests IS curves using an ADL specification that encompasses the traditional Keynesian IS curve (KISC), the New Keynesian IS curve (NKISC), and the Hybrid IS curve (HISC). We use data from the

[1] According to table C of Hammond (2010), 20 out of 27 inflation targeting central banks either use or are developing Dynamic Stochastic General Equilibrium (DSGE) models based on the New Keynesian framework for forecasting and policy analysis.

United States (1985Q1–2007Q4) and from Brazil (1996Q1–2012Q2).

The distinguishing feature of these essays is the empirical strategy followed that allows us to test different theoretical models within the same empirical framework. The empirical strategy begins with the formulation of a statistical model that encompasses alternative Phillips (IS) curve models as special cases–the unrestricted ADL Phillips (IS) curve model. The unrestricted ADL model specification takes the following general form in a closed economy:

$$\Delta y_t = \sum_{s=1}^{m} \alpha_s \Delta y_{t-s} + \sum_{s=0}^{n} \beta_s z_{t-s} + \gamma_1 y_{t-1} + \varepsilon_t$$

Each Phillips (IS) curve model implies restrictions on the set of parameters $(\alpha_1, ..., \alpha_m, \beta_0, ..., \beta_n, \gamma_1)$. The empirical strategy proceeds by estimating the unrestricted ADL model and reducing it based on statistical criteria. Finally, we test each Phillips (IS) curve model by comparing its implied restrictions with the values of the estimated parameters of the restricted ADL model.

Recently, there has been a worldwide movement toward the adoption of a policy regime in which the central bank is assigned the task of achieving an inflation target. At the same time, the independence of central banks to pursue this goal has also increased, suggesting that the choice of monetary policy to achieve the inflation target is a problem that can, and in fact ought to be, separated from the choice of fiscal policy or any other public policy. While there is an agreement between most economists regarding the effects of monetary policy shocks, the empirical literature has struggled so far to provide robust stylized facts on the effects of fiscal policy shocks. In particular, there is no agreement on even the qualitative effects of fiscal policy shocks on macroeconomic variables.

The third essay ("The Effects of Fiscal Policy and its Interactions with Monetary Policy in Brazil") analyzes the effects of fiscal policy shocks on the dynamics of the economy and the interaction between fiscal and monetary policy in Brazil.[2] It uses a structural vector autoregression (SVAR) model and the test proposed by Canzoneri, Cumby, and Diba (2001). The SVAR is identified by two alternative methodologies. The first methodology uses sign restrictions on impulse responses of the exogenous disturbances. The second methodology, developed by Lima, Maka, and Alves (2011) [LMA], combines sign restrictions with restrictions on the contemporaneous causal interrelationships among variables derived by Directed Acyclic Graphs (DAGs). LMA analysis is concerned mainly with the identification of the effects of monetary policy and exchange rate shocks, so no attention was given to fiscal policy. In this paper, we extend LMA analysis introducing a set of fiscal variables (budget surplus and public sector liabilities) in their VAR model. The hybrid identification strategy pursued consists of two steps. In the first step, we use DAGs to select over-identifying restrictions on the contemporaneous coefficients based on the conditional independence relations between the variables. These over-identifying restrictions allow us to identify monetary policy and demand shocks and to restrict the covariance matrix of the reduced-form residuals. In the second step, maintaining restricted the covariance matrix of reduced-form residuals, we keep the identified monetary policy and demand shocks, and impose sign restrictions on the impulse response functions on the other three shocks to identify the fiscal policy, supply, and exchange rate shocks.

[2]This paper is a joint work with Elcyon Caiado Rocha Lima and Amadeu Pumar. It was published in *Economia*, vol. 13, n.1, 149–160 (2012).

# Chapter 2

# Phillips Curves: An Encompassing Test

## 2.1 Introduction

The empirical evidence shows that inflation tends to be pro-cyclical: periods of above average inflation tend to be associated with above average economic activity. This statistical relationship is known as the Phillips curve. The Phillips curve was perceived in the 60's as a menu for monetary policymakers: they could choose between high inflation and low unemployment or low inflation and high unemployment. But this interpretation of the Phillips curve assumed that the relationship between unemployment and inflation was stable and would not break down once a policymaker attempts to exploit the tradeoff. After Friedman's (1968) paper and the high inflation episodes experienced by many economies in the 1970s, this interpretation of the Phillips curve was discredited. After a period of low inflation in the 1980s and early 1990s, economists have again worked on a theoretical framework for the Phillips curve. The New Keynesian Phillips curve (NKPC) provides an interpretation of the short-run inflation-unemployment trade-off by deriving it from an optimizing framework featuring rational expectations and nominal rigidities. This is a structural model, designed to be capable of explaining the behavior of inflation without being subject to the Lucas critique. The NKPC is part of the New Keynesian model which is the workhorse model for monetary analysis. However, to use the NKPC for policy analysis requires it to have a good econometric track record in describing inflation dynamics.

This paper tests Phillips curves using an autoregressive distributed lag (ADL) specification that encompasses the accelerationist Phillips curve (APC), the New Keynesian Phillips curve (NKPC), the Hybrid Phillips curve (HPC), the Sticky-Information Phillips curve (SIPC). We use data from the United States (1985Q1–2007Q4) and from Brazil (1996Q1–2012Q2), using the output gap and alternatively the real marginal cost as measure of inflationary pressure.[1] The empirical evidence rejects the restrictions implied by the NKPC, the HPC, and SIPC, but does not reject those implied by the APC.

A large body of research has used time series methods to estimate the NKPC. Initial attempts to estimate the NKPC using aggregate time series data for the U.S. were not very successful [Galí and Gertler (1999), GG henceforth]: the estimated coefficient on the output gap (proxied by detrended real GDP) was small and often negative in quarterly

[1] Given that we estimate reduced-form models, the choice of the sample in both countries is motivated by the attempt of avoiding the Lucas critique by selecting periods of economic regime stability.

data. One interpretation for the poor results using a standard output gap measure is that it is simply a poor proxy for real marginal cost, which according to the theory, would be the appropriate variable. GG report evidence in favor of the New Keynesian Phillips curve when labor's share of income, rather than a standard output gap variable, is used to proxy for real marginal cost. In order to capture the inflation persistence found in the data, GG modify the basic Calvo model of sticky prices to introduce lagged inflation into the Phillips curve, called hybrid Phillips curve. Based on U.S. data and using real marginal cost as the forcing variable, GG conclude that not only the forward looking behavior is predominant but, given the small estimate of the degree of backwardness, the pure forward looking model may do a reasonably good job of describing the data. Following the steps of GG, Galí, Gertler and López-Salido (2001) provided evidence on the fit of the NKPC for the Euro area.

Rudd and Whelan (2007) in a critical review of the New Keynesian model argue that the labor's share version of the model actually provides a very poor description of observed inflation behavior. This failure of the model extends along two dimensions: first, labor's share fails to provide a good measure of inflationary pressures; and second, this version of the model cannot account for the important role played by lagged inflation in empirical inflation regressions. They provide an alternative interpretation of the empirical estimates obtained from the hybrid Phillips curve, and argue that the data actually provide very little evidence of an important role for forward-looking behavior of the sort implied by these models.

As an alternative to the models of sticky prices, Mankiw and Reis (2002) [MR] argue that sticky information—the slow dispersion of information about macroeconomic conditions—can help account for the sluggish adjustment of prices and for the real effects that occur in response to monetary shocks. Kiley (2007) attempted to test the sticky information model of inflation against the sticky price for the United States using maximum-likelihood techniques. He finds that, once hybrid-behavior is allowed, hybrid sticky-price models provide a better description of inflation dynamics than a sticky-information model.

In a small open-economy exchange-rate movements play an important role in the transmission process that links monetary disturbances to output and inflation movements. Economic disturbances that originate in other countries have to be taken into account in monetary policy design that are absent in a closed-economy environment. In addition, part of the intermediate inputs are imported from abroad, which consist mainly of raw materials and energy. Usually the prices of imported inputs are more variable than those from domestic labor as well as domestically produced intermediate inputs. This should—other things equal—induce firms to change their prices more frequently and possibly also by a larger amount in response to more variable input costs. Price setting is also more complex as the choice of currency, competition from abroad and the pass-through of exchange rate changes into prices become an issue. It is therefore not surprising that model building becomes increasingly difficult when it comes to modelling inflation dynamics in an open economy where the relationship between inflation and exchange rate is a key concern.

Several authors have attempted to estimate the Phillips curve for Brazil and some of them were concerned specifically with the NKPC. For example, Areosa and Medeiros (2007) derive and estimate a structural model for inflation in a small open economy based on the models developed by Campos and Nakane (2003) and Galí and Monacelli (2005). Their model considers price rigidity according to Calvo (1983) and has inflationary inertia in a way similar to Woodford (2003) and GG. The estimated negative sign in the specification with the output gap contrasts with theory and although the impact of the estimate associated with the marginal cost is negligible, it is statistically significant. They find a small direct impact of the variables associated with economic openness, with the sum of their coefficients being close to zero. However, the indirect impact is significant, consistently changing the weights associated with lagged inflation and the expected future inflation. Mazali and Divino (2010) apply for Brazil the new Keynesian model of Blanchard and Galí (2007) with real wage rigidity and supply shocks. As the estimated coefficients satisfied a set of restrictions imposed by the theoretical model and over-identifying restrictions were not rejected, they concluded that the estimated New Keynesian Phillips curve adjusted very well to the Brazilian data.[2]

The paper is organized as follows. Section 2 reviews the literature on Phillips curves. Section 3 presents the ADL Phillips curve and shows how the different Phillips curve specifications considered in the literature can be seen as special cases of the ADL Phillips curve. Section 4 tests for the U.S. the restrictions implied by the different Phillips curve specifications. Section 5 estimates the ADL Phillips curve for Brazil and tests the fit of the alternative models. Finally, section 6 brings the concluding remarks.

## 2.2 Phillips Curves

In a seminal paper, Phillips (1958) showed that there was a negative and relatively stable relationship between nominal wage inflation and unemployment in the United Kingdom over the previous century. This relationship was found to work well for price inflation and for other economies, receiving the name of "Phillips curve". It became a key part of the standard Keynesian textbook model of the 1960s and as Keynesian economists saw it, the Phillips curve provided an exploitable tradeoff between inflation and unemployment: policy-makers could use demand management policies to increase output and decrease unemployment, but this could only be done at the expense of higher inflation. The Phillips curve relationship can be represented as

$$\pi_t = \alpha - \gamma u_t, \tag{2.1}$$

where $\pi_t$ is inflation, $u_t$ is the unemployment rate, and $\gamma > 0$.

The theoretical foundations of these early formulations were not completely sound, with a particular weak point being their treatment of how expectations would enter wage and price setting. This weakness was thoroughly criticized in the seminal contributions of Phelps (1967, 1968) and Friedman (1968).

### 2.2.1 Accelerationist Phillips Curve (APC)

Friedman argued that the correct formulation of the inflation-unemployment tradeoff was a Phillips curve of the form:

$$\pi_t = E_{t-1}\pi_t - \gamma(u_t - \overline{u}), \tag{2.2}$$

where inflation, $\pi_t$, is negatively correlated with deviations of the unemployment rate from its natural rate $\overline{u}$ ($\gamma > 0$) and where the entire curve is shifted up or down one-for-one

[2] Other articles that estimate the Phillips curve for Brazil include Cysne (1985), Lima (2003), Minella et al. (2003), Fasolo and Portugal (2004), Schwartzman (2006), Lima and Brito (2009), Holland and Mori (2010), Mendonça, Sachsida and Medrano (2012).

with changes in the rate of inflation that agents expected at time *t-1* to prevail at time $t$, $E_{t-1}\pi_t$. A common variant of this equation replaces $u_t - \overline{u}$ with the gap between actual and potential output, $y_t - \overline{y} \equiv x_t$.

Friedman predicted that attempts to keep unemployment low at the expense of higher inflation would just result in raised inflation expectations. Thus, the economy would not be able to sustain the low unemployment and would end up with higher inflation. In the Friedman-Phelps framework, then, there is no permanent tradeoff between the level of inflation and the unemployment rate—in the long run, $E_{t-1}\pi_t = \pi_t$, and the Phillips curve is vertical at $u_t = \overline{u}$ . However, to the extent that agents' expectations were slow to catch up with the reality, a policymaker could keep unemployment below the natural rate by constantly boosting the inflation rate. For this reason, the Friedman-Phelps characterization of the inflation process also came to be known as the "accelerationist hypothesis" since an acceleration in prices (an increase in the rate of inflation) would occur should policymakers attempt to permanently keep unemployment below its natural rate. Phelps assumed that inflation expectations evolved over time as a result of actual past experience—that is, that expectations were formed adaptively[3]. There is a long tradition in applied work that assumes backward-looking expectations: expected inflation is determined by past inflation. In the special case where $E_{t-1}\pi_t = \pi_{t-1}$, the Phillips curve becomes

$$\pi_t = \pi_{t-1} - \gamma(u_t - \overline{u}). \tag{2.3}$$

This so-called accelerationist Phillips curve—in which the acceleration of prices is related to unemployment—embodied two critical innovations in the literature. First, it eliminated the long-run trade-off between inflation and unemployment that was inherent in the original Phillips curve model. Second, it began to emphasize the importance of expectations in the price-setting process, a change that was to have dramatic implications in the evolution of inflation models for the next four decades.

In the decade following the publication of the Phelps and Friedman papers, the notion that the accelerationist view of the inflation process was correct gained wider acceptance. Several factors contributed to this attitude. The first, of course, was the strength of the theoretical arguments themselves. Second, it became apparent by the mid-1970s that the inflation-unemployment tradeoff implied by the short-run Phillips curve had shifted. Finally, it became easier to find that the lags of inflation in empirical Phillips curves summed to one. In addition, the important contribution of "supply shocks" to price acceleration in the early 1970s led to food, energy, and/or import prices receiving special treatment in empirical descriptions of inflation. What emerged in this period, therefore, was a benchmark econometric model of inflation of the form:

$$\pi_t = \alpha - \gamma u_t + B(L)\pi_{t-1} + \lambda z_t + \varepsilon_t, \tag{2.4}$$

where $B(L)$ is the distributed lag operator with $B(1) = 1$ , $z_t$ denotes a vector of supply shocks, and $\varepsilon_t$ is an error term. In this specification, then, inflation dynamics are determined by three sources: real activity (as summarized by the unemployment rate), supply shocks, and "inertia" (as captured by the lagged inflation terms); for this reason, it is sometimes called the "triangle model".

The basic framework of the triangle model has been widely adopted by applied macroeconomists, and has brought with it the important concept of the NAIRU, defined as

[3] In Phelps (1967), the appeal to adaptive expectations is explicit. The term is not used by Friedman (1968), who provides an informal discussion of a gradual adjustment process.

the unemployment rate consistent with unchanging inflation in the absence of supply shocks. In the context of the simple version of the triangle model presented above (equation 4), the NAIRU is calculated from the estimated coefficients as $u^* = \alpha/\gamma$ . More refined analyses of the NAIRU include Staiger, Stock, and Watson's (1997a) improved approach for quantifying the precision with which the NAIRU can be measured, attempts by Staiger, Stock, and Watson (1997b), Gordon (1998), and Ball and Mankiw (2002) to model variations over time in the NAIRU.

Taken literally, the characterization of inflation dynamics that the triangle model provides carries important implications for the conduct of macroeconomic policy. To the extent that lagged inflation captures true inertia in the price-setting process (resulting, for instance, from how expectations are formulated), the model implies that rapid reductions in inflation can only be produced at the cost of a substantial increase in unemployment. Hence, the model points to a gradualist approach as providing the best way to effect a large reduction in inflation. In addition, policymakers must be mindful of the presence of long time lags between macroeconomic shocks (including policy actions) and their full effects on inflation. Thus, this framework provides a strong argument in support of preemptive action to cut off the full effect of an inflationary shock.

The introduction of rational expectations into the modelling of economic dynamics had a significant influence on the development of macroeconomic theory from the mid-1970s onwards. The "demise" of the traditional Phillips curve, and the sense that it was due to inadequate modelling of expectations, was a major impetus for the rational expectations school led by Robert Lucas and Thomas Sargent. Lucas and Sargent also rejected the "accelerationist" reformulation of the Phillips curve because it relied on the assumption of adaptive expectations, which do not allow for the idea that agents process information in an optimal manner. In addition to being more precise about expectations formation, this school of economists relied more heavily on neoclassical "microfoundations" for macroeconomic models. Often, as well as rejecting the Phillips curve, these economists also questioned the whole basis for Keynesian economics, i.e. the assumption that monetary policy could systematically affect output even in the short-run. However, the Lucas supply curve advanced by Lucas (1972)—the so-called "islands" model in which a short-run inflation-output tradeoff resulted from imperfect information on the part of price-setters—generally failed to convince as a model of economic fluctuations. In part, this initial resistance to the rational expectations approach likely reflected the Lucas model's implication that only purely random and transitory policy shocks could affect output

The principal response of Keynesian economists to these theoretical critiques has been to attempt to build models that incorporate rational expectations and that provide a microeconomic justification for monetary policy having, at least, short-run effects. To explain why monetary policy might have effects on the economy, one needs a theory of why inflation is not just determined by some nominal anchor such as the money supply. The most common microeconomic rationale put forward has been sticky prices. With sticky prices, an increase in the money stock can produce a short-run increase in real spending power and thus can boost real output. This modern approach, featuring rational expectations and some form of microfoundations, focusing on the role played by various rigidities and market imperfections is known as New Keynesian macroeconomics. In the New Keynesian literature, models of sticky prices have been grouped into two general categories: "time-dependent"and "state-dependent"models. In state-dependent models, firms change prices when underlying determinants, such as demand and costs, reach certain bounds. In time dependent models, firms set their prices for fixed periods of time. Time dependent models can have explicit closed-form solutions relating current price changes to future price changes and the current state of demand. State-dependent models do not, in general, have simple closed-form solutions. Many academic economists have become convinced that certain theoretical new Keynesian models can provide a good description of the empirical inflation process. In part, this development stemmed from the realization that a number of popular new Keynesian models of price-setting each implied a sort of Phillips curve relationship. We will now describe one of the key New-Keynesian models and explore its implications for the behaviour of inflation and output.

### 2.2.2 New Keynesian Phillips Curve (NKPC)

Three papers that arrived on the scene around the same time provided partial answers to this question regarding the existence of rigid (or"sticky") wages and prices. Both authors, Calvo (1983) and Rotemberg (1982, 1983), offered different explanatory rationales, but when stripped to their respective cores, their models are nearly identical and bear essentially the same implications for macroeconomic dynamics.

Calvo's (1983) model of random price adjustment assumes that in each period a random fraction $(1-\omega)$ of firms can adjust their prices, while all the other firms keep their prices unchanged. The expected time between price adjustments is $1/(1-\omega)$ . The evolution of the (log) price level is given by

$$p_t = \omega p_{t-1} + (1-\omega) z_t, \tag{2.5}$$

where $z_t$ is the price chosen by those who can adjust their prices. Assuming a monopolistic competition structure such that, absent any frictions, firms would set their price as a fixed markup over marginal cost, a firm's optimal reset price is determined by

$$z_t = (1-\omega\beta) \sum_{j=o}^{\infty} (\omega\beta)^j \, E_t(MC_{t+j} + \mu), \tag{2.6}$$

where $\mu$ is the frictionless optimal markup, $\beta$ is the firm's discount factor, and $MC_t$ is nominal marginal cost. In other words, firms take into account that their prices will likely be fixed over some period by setting their price equal to a weighted average of expected future nominal marginal costs. These two equations can be combined to yield a NKPC of the form[4]:

$$\pi_t = \beta E_t \pi_{t+1} + \frac{(1-\omega)(1-\beta\omega)}{\omega}(mc_t + \mu), \tag{2.7}$$

which relates current inflation to next period's expected inflation rate and the deviation of current real marginal cost ($\mathrm{mc}_t$) from its optimal frictionless level.

Under relatively general conditions, aggregate real marginal cost is proportional to the gap between output and its potential level[5]. With this assumption, the NKPC becomes

$$\pi_t = \beta E_t \pi_{t+1} + \gamma x_t, \tag{2.8}$$

where $x_t$ is a measure of output gap.

[4] For a detailed derivation of the NKPC, see for example, Walsh (2010).

[5] In the standard sticky price framework without variable capital, there is an approximate proportionate relation between marginal cost and output. With variable capital the relation is no longer proportionate, but simulations suggest that the relation remain very close to proportionate.

Inflation is determined in a completely forward-looking manner, as can be seen by solving forward equation (2.8) to obtain the following closed-form solution for the "fundamental"inflation[6]:

$$\pi_t = \gamma \sum_{k=o}^{\infty} (\beta)^k E_t(x_{t+k}). \tag{2.9}$$

The idea that there is considerable inertia in inflation and hence that it is difficult to reduce inflation quickly, does not hold in this framework— indeed, according to the NKPC, there is no "intrinsic" inertia in inflation, in the sense that there is no structural dependence of inflation on its own lagged values. Thus, the NKPC has very different implications for monetary policy. This model implies that there is no need for gradualist policies to reduce inflation. According to the NKPC, low inflation can be achieved immediately by the central bank announcing (and the public believing) that it is committing itself to elminating positive output gaps in the future.

### 2.2.3 Hybrid Phillips Curve (HPC)

Many estimates of the NKPC find that lagged inflation helps to explain current inflation. GG consider augmenting the NKPC with a backward-looking element that is motivated by the presence of some firms that follow a simple rule of thumb in setting prices. Christiano, Eichenbaum, and Evans (2005) derive a similar specification under the assumption that price-setters who are unable to reset prices instead index their prices to the last period inflation rate. All of these variants imply a so-called HPC of the form

$$\pi_t = \gamma_b \pi_{t-1} + \gamma_f E_t \pi_{t+1} + \kappa x_t. \tag{2.10}$$

If $\gamma_b + \gamma_f = 1$, then the Phillips curve is vertical in the long-run, that is, there is no long-run tradeoff between inflation and real activity.

As shown by Barbosa (2013b), the estimates $\gamma_b$ and $\gamma_f$ can be misleading indicators of forward- and backward-looking behavior: it is possible to have $\gamma_f > \gamma_b$ but the model can in fact be totally backward-looking.

### 2.2.4 Sticky-Information Phillips Curve (SIPC)

The Mankiw and Reis (2002) [MR] model pioneers the literature on sticky information and argues that a Phillips curve with this rigidity is an adequate representation of the structural relationship between inflation and the real side of the economy. MR derives the SIPC based on the assumption that firms set price freely but the information for setting prices may or may not be updated. Because information constraints could apply to all economic agents, the sticky information model potentially provides a unifying framework for explaining inertial behavior in different macroeconomic variables.

The SIPC assumes that firms set their price every period, but firms gather information and re-compute optimal prices slowly over time. In each period, a fraction $\lambda$ of firms obtains new information about the state of the economy and computes a new path of optimal prices, and $(1 - \lambda)$ firms continue to set prices based on old plans and outdated information.

---

[6] In order to get the solution of inflation as the present discounted value of outputgaps we assume that $\lim_{T\to\infty} \gamma \sum_{i=0}^{T} \beta^i E_t x_{t+i}$ converges and that $\lim_{T\to\infty} \beta^T \pi_{t+T} = 0$.

Each firm has the same probability of being one of the firms updating their information sets, regardless of how long it has been since their last update. The expected time between price updates is therefore $1/\lambda$.

A firm's optimal price that maximizes expected profits at any given point in time is $p_t^* = p_t + \alpha x_t$ , where $p_t$ is the overall price level and $x_t$ is the output gap. A firm that lasts updated its plans j periods ago sets its price as the expected value of the optimal price j periods ago: $z_t^j = E_{t-j} p_t^*$.

The aggregate price level is the average of the prices of all firms in the economy, assuming that the arrival of decision dates is a Poisson process given by:

$$p_t = \lambda \sum_{j=0}^{\infty} (1-\lambda)^j z_t^j.$$

The price level is then defined by $p_t = \lambda \sum_{j=0}^{\infty} (1-\lambda)^j E_{t-j} \left(p_t + \alpha x_t\right)$, so that, the baseline SIPC is defined as:

$$\pi_t = \lambda \sum_{j=0}^{\infty} (1-\lambda)^j E_{t-1-j}(\pi_t + \alpha \Delta x_t) + \frac{\alpha\lambda}{1-\lambda} x_t + \varepsilon_t. \tag{2.11}$$

In this set-up, inflation depends on output, past expectations of current inflation and past expectations of changes in current output gap growth.

## 2.3 Autoregressive Distributed Lag (ADL) Phillips Curve

A model of inflation dynamics general enough to encompass the NKPC, the APC, the HPC, and the SIPC as special cases takes the form

$$\Delta\pi_t = \alpha_1 \Delta\pi_{t-1} + \beta_0 x_t + \beta_1 x_{t-1} + \gamma_1 \pi_{t-1} + \varepsilon_t. \tag{2.12}$$

Let us show how each model is embedded in equation (2.12). The simple APC (SAPC) is given by

$$\pi_t = \pi_{t-1} + \kappa x_t + \varepsilon_t,$$

which is equivalent to

$$\Delta\pi_t = \kappa x_t + \varepsilon_t.$$

This is a particular case of equation (2.12) when $\alpha_1 = 0$, $\beta_0 = \kappa > 0$, $\beta_1 = \gamma_1 = 0$. When expected inflation of the Phillips curve depends on additional inflation lags we have the APC. For the case in which expected inflation depends on the last two inflation lags the APC is given by

$$\pi_t = \omega_1 \pi_{t-1} + \omega_2 \pi_{t-2} + \kappa x_t + \varepsilon_t.$$

After some algebra, this can be rearranged as

$$\Delta\pi_t = -\omega_2 \Delta\pi_{t-1} - (1 - \omega_1 - \omega_2)\pi_{t-1} + \kappa x_t + \varepsilon_t.$$

Assuming that $\omega_1 + \omega_2 = 1$, that is, there is no long-run trade-off between inflation and real activity, then

$$\Delta\pi_t = -\omega_2 \Delta\pi_{t-1} + \kappa x_t + \varepsilon_t.$$

This specification is a particular case of equation (2.12) when $\alpha_1 = -\omega_2 < 0$, $\beta_0 = \kappa > 0$, $\beta_1 = \gamma_1 = 0$.

The NKPC is given by

$$\pi_t = \beta E_t \pi_{t+1} + \kappa x_t + \varepsilon_t.$$

Assuming rational expectations, $E_t \pi_{t+1} = \pi_{t+1} - \mu_{t+1}$ , the NKPC can be written as

$$\Delta\pi_t = \frac{1-\beta}{\beta}\pi_{t-1} - \frac{\kappa}{\beta} x_{t-1} + \eta_t, \;\; \beta \neq 0.$$

This follows from equation (2.12) when $\alpha_1 = \beta_0 = 0$, $\beta_1 = -\kappa/\beta < 0$, $\gamma_1 = (1-\beta)/\beta > 0$.

The HPC is given by[7]

$$\pi_t = \gamma_b \pi_{t-1} + \gamma_f E_t \pi_{t+1} + \kappa x_t + \varepsilon_t.$$

As with the NKPC, we assume rational expectations and after some algebra we obtain

$$\Delta\pi_t = \frac{\gamma_b}{\gamma_f}\Delta\pi_{t-1} - \frac{\kappa}{\gamma_f} x_{t-1} + \frac{(1-\gamma_b-\gamma_f)}{\gamma_f}\pi_{t-1} + \xi_t, \;\; \gamma_f \neq 0.$$

This is a particular case of equation (2.12) when $\alpha_1 = \gamma_b/\gamma_f > 0$, $\beta_0 = 0$, $\beta_1 = -k/\gamma_f < 0$, $\gamma_1 = (1-\gamma_b-\gamma_f)/\gamma_f > 0$.

The SIPC derived by MR is given by

$$\pi_t = \lambda \sum_{j=0}^{\infty} (1-\lambda)^j E_{t-1-j}(\pi_t + \alpha \Delta x_t) + \frac{\alpha\lambda}{1-\lambda} x_t + \varepsilon_t.$$

Assuming rational expectations and using the lag operator $L^j E_{t-1} = E_{t-1-j}$ , we obtain after some algebra the following expression for the acceleration of inflation

$$\Delta\pi_t = \frac{\alpha\lambda(2-\lambda)}{(1-\lambda)^2} x_t - \frac{2\alpha\lambda(1-\lambda)}{(1-\lambda)^2} x_{t-1} + \varsigma_t, \;\; \lambda \neq 0.$$

This expression is a particular case of equation (2.12) when $\alpha_1 = 0$, $\beta_0 = \alpha\lambda(2-\lambda)/(1-\lambda)^2 > 0$, $\beta_1 = -2\alpha\lambda(1-\lambda)/(1-\lambda)^2 < 0$, $\gamma_1 = 0$.

Table 2.1 shows the restrictions implied by each Phillips curve model subsumed by equation (2.12), which is an ADL(1,1). Once the parameters of the model have been estimated – $(\alpha_1, \beta_0, \beta_1, \gamma_1)$ – table 2.1 enables us to identify which Phillips curve model is consistent with the evidence. For example, if the estimated coefficient of $\Delta\pi_{t-1}(\alpha_1)$ is negative, then only the APC would be consistent with data.

| **Model** | **Parameters** | | | |
|---|---|---|---|---|
| | $\alpha_1$ | $\beta_0$ | $\beta_1$ | $\gamma_1$ |
| **APC** | - | + | 0 | 0 |
| **NKPC** | 0 | 0 | - | + |
| **HPC** | + | 0 | - | + |
| **SIPC** | 0 | + | - | 0 |

Table 2.1: Model Typology

[7] See, for example Galí and Gertler (1999).

The ADL Phillips curve [equation (2.12)] actually encompasses all Phillips curves as special cases. This provides a convenient framework for analysing the properties of each model used in empirical research, highlighting their respective strengths and weaknesses.

The unrestricted ADL Phillips curve includes additional lags of both output gap and inflation acceleration, according to:

$$\Delta\pi_t = \sum_{s=1}^{m} \alpha_s \Delta\pi_{t-s} + \sum_{s=0}^{n} \beta_s x_{t-s} + \gamma_1 \pi_{t-1} + \varepsilon_t. \tag{2.13}$$

In this paper we test the NKPC against the HPC, the SIPC, and the APC based on the encompassing ADL Phillips curve [equation (2.13)]. The empirical strategy, inspired in the general-to-specific approach to econometrics, is as follows: (i) formulate and estimate a Phillips curve model that encompasses the NKPC, the HPC, the SIPC and the APC as special cases—the unrestricted ADL Phillips curve; (ii) reduce the unrestricted ADL Phillips curve model based on statistical criteria—the restricted ADL Phillips curve; (iii) test each Phillips curve model by comparing its implied restrictions with the values of the estimated parameters of the restricted ADL model.[8]

## 2.4 Empirical Evidence: United States (1985Q1–2007Q4)

The U.S. sample goes from 1985Q1 to 2007Q4, the period known as "Great Moderation"due to the decline in the variability of both output and inflation.[9] Figure 2.1 plot the inflation rate and table 2.2 shows that it is stationary at the 5% level.[10] Figure 2.2 exhibits the output gap.[11]

| Null Hypothesis: Inflation has a unit root<br>Exogenous: Constant<br>Lag length: 2 (Spectral OLS AR based on SIC, maxlag=11)<br>Sample: 1985Q1 2007Q4 | | |
|---|---|---|
| | | P-Statistic |
| Elliott-Rothenberg-Stock Point Optimal test statistic | | 2.074 |
| Test critical values: | 1% level | 1.937 |
| | 5% level | 3.087 |
| | 10% level | 4.128 |

Table 2.2: Unit Root Test for Inflation (U.S.)

[8] For a survey on the general-to-specific modeling see Campos, Ericsson, and Hendry (2005).

[9] James Stock coined the phrase "the great moderation" while writing a research paper with Mark Watson in 2002 ("Has the Business Cycle Changed and Why?"). It was brought to the attention of the wider public by Ben Bernanke (then member and now chairman of the Board of Governors of the Federal Reserve) in a speech titled "The Great Moderation" in 2004. The results do not change significantly when we start or finish the sample one year earlier.

[10] The Data Appendix gives details on the definitions of the variables used in the baseline model and the robustness analysis.

[11] Proponents of the New Keynesian model criticize traditional measures of the output gap on the grounds that naive detrending procedures assume that potential GDP evolves smoothly over time. In theory, however, changes in potential output will be affected by any number of shocks, and so could fluctuate significantly (and stochastically) period to period. While the points regarding the difficulties surrounding the measurement of potential output are clearly valid and important, these points should not, on their own, be taken as reasons to jeopardize the usefulness of standard output gap measures.

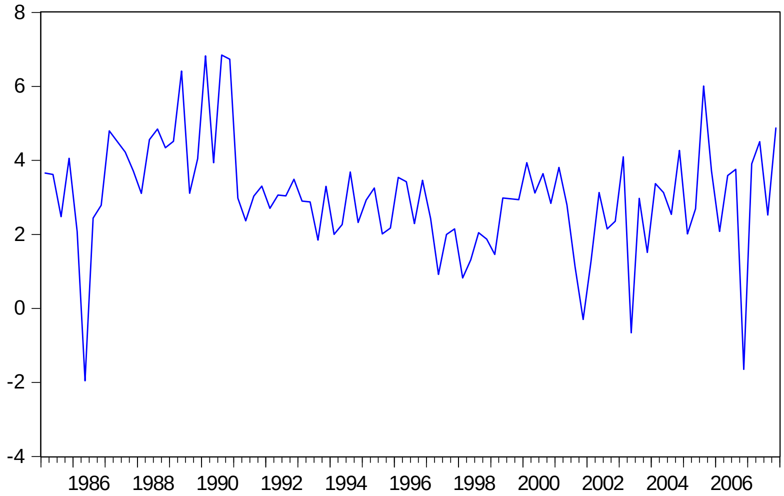



Figure 2.1: Inflation (U.S.)

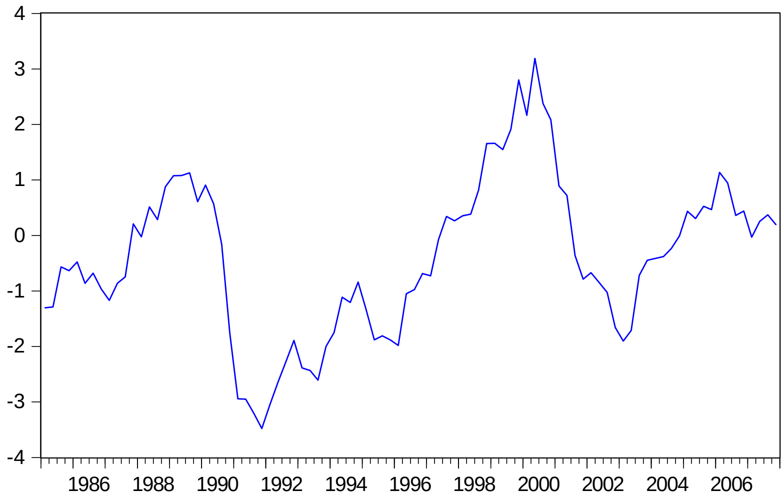



Figure 2.2: Output Gap (U.S.)

Our empirical analysis starts with the estimation of the unrestricted ADL Phillips curve [equation (2.13)]. Given the presence of the current output gap ($x_t$) as a regressor, one may suspect that the Ordinary Least Square (OLS) assumption of orthogonality between regressors and the error term ($\varepsilon_t$) may not hold. If $x_t$ is correlated with $\varepsilon_t$, then the OLS estimates will be inconsistent. A solution to this endogeneity problem lies in the use of Instrumental Variable (IV)/Generalized Method of Moments (GMM) estimators. However, if $x_t$ is not correlated with $\varepsilon_t$ and we use IV/GMM estimators, there will be a loss of efficiency, as the asymptotic variance of the IV/GMM is always larger (sometimes much larger) than the asymptotic variance of the OLS estimator. Therefore, before proceeding with the estimation of equation (2.13) we test the endogeneity of $x_t$ in order to check the appropriateness of OLS or the necessity to resort to IV/GMM methods.[12] We apply the GMM distance test (also known as C test) of a subset of orthogonality conditions to test the endogeneity of $x_t$.[13,14] The endogeneity test displayed on table 2.3, indicates that we can treat $x_t$ as an exogenous variable.[15] Therefore, we proceed with our empirical analysis by estimating the unrestricted ADL Phillips curve [equation (2.13)] using OLS, setting $m = n = 4$ [see table 2.4].

| Null Hypothesis: Output Gap is exogenous | | | |
|---|---|---|---|
| | value | d.f. | Probability |
| Difference in J-statistics | 1.085 | 1 | 0.297 |

Table 2.3: Endogeneity Test for Output Gap (U.S.)

[12] If orthogonality between regressors and the error term cannot be rejected, we should use OLS rather than IV or GMM, especially in small samples.

[13] In our context, the GMM distance test of a subset of orthogonality conditions is equivalent to the Durbin-Wu-Hausman endogeneity test. See Baum, Schaffer, and Stillman (2003, 2007).

[14] The GMM distance test is computed as the difference between two J-statistics: that for the (restricted, fully efficient) regression using the entire set of overidentifying restrictions, versus that for the (unrestricted, inefficient but consistent) regression using a smaller set of restrictions, in which a specified set of instruments is removed from the set. For instruments included in the equation to be estimated, the GMM distance test places them in the list of included endogenous variables; in essence, treating them as endogenous regressors. The GMM distance test, distributed $\chi^2$ with degrees of freedom equal to the loss of overidentifying restrictions (i.e., the number of suspect instruments being tested), has the null hypothesis that the specified variables are proper instruments, that is, they are exogenous.

[15] Given the structure of the error term when the ADL Phillips curve is restricted to be either the NKPC or the HPC, another souce of inconsistency of the OLS estimator would be the one-period lagged regressors. We performed instrument orthoganality tests (C tests) for all one-period lagged regressors and for all of them it was not possible to reject the hypothesis that they are exogenous.

| Dependent Variable: $\Delta\pi_t$ | | |
|---|---|---|
| Estimation Method: OLS | | |
| sample: 1985Q1–2007Q4 | | |
| Variable | Coefficient | Standard Error |
| $\Delta\pi_{t-1}$ | -0.589 | 0.113 |
| $\Delta\pi_{t-2}$ | -0.422 | 0.118 |
| $\Delta\pi_{t-3}$ | 0.023 | 0.105 |
| $x_t$ | -0.266 | 0.340 |
| $x_{t-1}$ | 0.841 | 0.497 |
| $x_{t-2}$ | 0.284 | 0.485 |
| $x_{t-3}$ | -1.294 | 0.497 |
| $x_{t-4}$ | 0.471 | 0.338 |
| $\pi_{t-1}$ | -0.036 | 0.049 |
| Diagnostic Tests | | |
| AR 1-5 test: F(5,75) | 2.862[0.020]* | |
| ARCH 1-4 test: F(4,80) | 3.087 [0.020]* | |
| Normality test: Chi^2(2) | 5.052 [0.0800] | |
| Hetero test: F(16,71) | 1.064 [0.404] | |
| Hetero-X test: F(44,43) | 0.691 [0.886] | |
| RESET23 test: F(2,78) | 1.262[0.288] | |

The values in brackets denote probability values (p-values).
Significant diagnostic tests at 5% level are shown by one star.

Table 2.4: Unrestricted ADL Phillips Curve (U.S.)

Next, we reduce the unrestricted ADL Phillips curve by sequential elimination of regressors based on t-ratios of the parameters estimators. We sequentially delete those regressors with the smallest t-ratios and re-estimate the model until the t-ratios are all greater than the threshold $\gamma = 2$.[16] The resulting restricted ADL model is displayed on table 2.5.

[16] In a context of VAR models, Brüggemann and Lütkepohl (2001) showed that this strategy is equivalent to the sequential elimination based on model selection criteria if the threshold value $\gamma$ is chosen accordingly.

| Dependent Variable: $\Delta\pi_t$ | | |
|---|---|---|
| Estimation Method: OLS | | |
| sample: 1985Q1–2007Q4 | | |
| Variable | Coefficient | Standard Error |
| $\Delta\pi_{t-1}$ | -0.641 | 0.092 |
| $\Delta\pi_{t-2}$ | -0.474 | 0.092 |
| $x_{t-1}$ | 0.585 | 0.207 |
| $x_{t-3}$ | -0.532 | 0.204 |
| Diagnostic Tests | | |
| AR 1-5 test: F(5,80) | 3.174[0.011]* | |
| ARCH 1-4 test: F(4,81) | 3.889 [0.006]** | |
| Normality test: Chi^2(2) | 10.728 [0.003]** | |
| Hetero test: F(8,80) | 1.701 [0.110] | |
| Hetero-X test: F(14,74) | 1.264 [0.250] | |
| RESET23 test: F(2,83) | 1.142 [0.324] | |

The values in brackets denote probability values (p-values).
Significant diagnostic tests at 5% level are shown by one star.
Significant diagnostic tests at 1% level are shown by two stars.

Table 2.5: Restricted ADL Phillips Curve (U.S.)

According to Table 2.5, the restricted model is composed by the first two lags of the change in inflation, and the lags of order one and three of the output gap. Using equation (2.12) and comparing the values of Tables 2.1 and 2.5, we observe that only one of the restrictions implied by the SIPC is verified ($\gamma_1 = 0$). Only one of the restrictions implied by the NKPC and the HPC is accepted ($\beta_0 = 0$), while all their other restrictions are rejected by the U.S. data. The APC has two restrictions correct ($\alpha_1 < 0$ and $\gamma_1 = 0$) and restrictions on the output gap ($\beta_0 > 0$ and $\beta_1 = 0$) that are consistent with the fact that their sum is positive when inflation depends on additional lags of output gap.

We conclude that only the APC model is consistent with inflation dynamics in the U.S. from 1985Q1–2007Q4. It should be noticed, however, that the diagnostic tests of Table 2.5 indicate some misspecification problems with the estimated model, as the hypothesis of normality of the distribution of the residuals is rejected, and there is evidence of autocorrelation and conditional heteroscedasticity of the residuals.

The restricted ADL Phillips curve of table 2.5 can be written in the Error-Correction form as[17]

$$\Delta\left(\Delta\pi_t\right) = -\alpha_2\Delta\left(\Delta\pi_{t-1}\right) - \beta_3(\Delta x_{t-1} + \Delta x_{t-3}) - (1-\alpha_1-\alpha_2)\left[\Delta\pi_{t-1} - \frac{\beta_1+\beta_3}{1-\alpha_1-\alpha_2}x_{t-1}\right] + \varepsilon_t$$

where $\left[\Delta\pi_{t-1} - \frac{\beta_1+\beta_3}{1-\alpha_1-\alpha_2}x_{t-1}\right]$ measures the deviation of $\Delta\pi_{t-1}$ from its the long-run solution. Table 2.6 shows that the coefficient of $x_{t-1}$ is not significantly different from zero, implying that inflation depend on output gap *changes*, not on its level.

[17] The Error Correction Model (ECM) represents the dynamics in a different way from the ADL model and captures short-run adjustments to past disequilibria and contemporaneous changes in the explanatory variables. All the information in an ADL model is contained in the ECM version and vice versa. It is just a matter of how the information is parameterised and interpreted.

| Null Hypothesis: $(\beta_1 + \beta_3) / (1 - \alpha_1 - \alpha_2) = 0$ | | | |
|---|---|---|---|
| | value | d.f. | Probability |
| Chi-square test statistic | 0.262 | 1 | 0.608 |

Table 2.6: Wald Test of Coefficient Restrictions (U.S.)

*Robustness*

Since Gordon (1982), many empirical researchers add supply shocks to the Phillips curve. Excluded supply shock variables are positively correlated with inflation and positively correlated with the output gap, so the omission of these supply shock variables causes the coefficient on the output gap to be biased towards zero. Following Ball and Mazumber (2011), we define core inflation as the part of inflation explained not by supply shocks. With this definition one can measure core inflation by removing the effects of supply shocks from total inflation. If supply shocks are asymmetries in the distribution of price changes, then a measure of core inflation should eliminate the effects of these asymmetries. A simple measure, proposed by Bryan and Cecchetti (1994), is the weighted median of price changes across industries (median inflation) [see Figure 2.3]. As a robustness check, we estimate the unrestricted ADL Phillips curve using median inflation, instead of the headline inflation. Tables 2.7 and 2.8 show that median inflation is difference stationary, which implies that the coefficient of lagged inflation in equation (2.13) is equal to zero ($\gamma_1 = 0$), while $t$-tests are still appropriate to evaluate the other (stationary) variables.[18] Table 2.10 displays the results of the estimated restricted ADL Phillips curve. With the median inflation the lag of order four of the output gap is now relevant, instead of order three, and the values of the coefficients are different. However, we reach the same conclusion as before: the APC is still consistent with inflation dynamics in the U.S.. Furthermore, the diagnostic tests of Table 2.10 indicate that the estimated model performs better when a measure of core inflation is used—median inflation—suggesting that the omission of supply shocks may be jeopardizing the estimation with headline inflation.

Regarding the Error-Correction representation, table 2.11 shows that the coefficient measures the deviation of $\Delta\pi_{t-1}$ from its long-run solution is significantly different from zero, implying that inflation depends on the level of the output gap and its changes.

| Null Hypothesis: Median Inflation has a unit root<br>Exogenous: Constant<br>Lag length: 1 (Spectral OLS AR based on SIC, maxlag=11)<br>Sample: 1985Q1 2007Q4 | | |
|---|---|---|
| | | P-Statistic |
| Elliott-Rothenberg-Stock Point Optimal test statistic | | 7.790 |
| Test critical values: | 1% level | 1.937 |
| | 5% level | 3.087 |
| | 10% level | 4.128 |

Table 2.7: Unit Root Test for Median Inflation (U.S.)

[18] See Hamilton (1995).

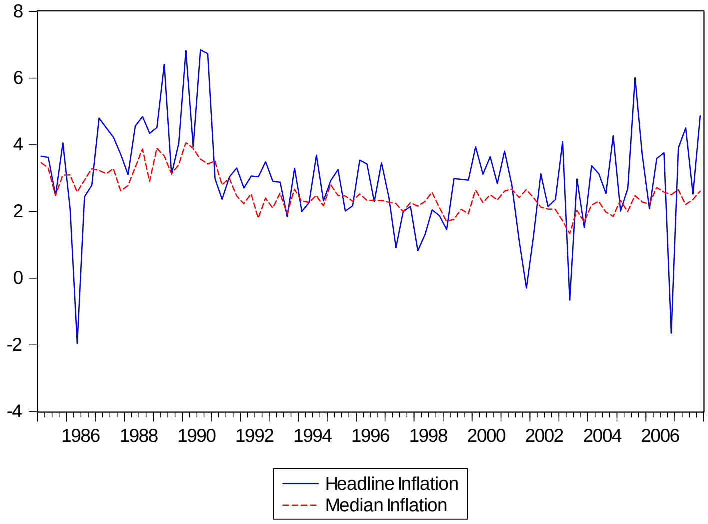



Figure 2.3: Median Inflation x Headline Inflation (U.S.)

| Null Hypothesis: $\Delta$ (median inflation)$_t$ has a unit root<br>Exogenous: Constant<br>Lag length: 1 (Spectral OLS AR based on SIC, maxlag=11)<br>Sample: 1985Q1 2007Q4 | | |
|---|---|---|
| | | P-Statistic |
| Elliott-Rothenberg-Stock Point Optimal test statistic | | 0.827 |
| Test critical values: | 1% level | 1.937 |
| | 5% level | 3.087 |
| | 10% level | 4.128 |

Table 2.8: Unit Root Test for the first difference of Median Inflation (U.S.)

| Null Hypothesis: Output Gap is exogenous | | | |
|---|---|---|---|
| | value | d.f. | Probability |
| Difference in J-statistics | 0.894 | 1 | 0.344 |

Table 2.9: Endogeneity Test for Output Gap when median inflation is used (U.S.)

| Dependent Variable: $\Delta\,(\text{median inflation})_t$<br>Estimation Method: OLS<br>sample: 1985Q1–2007Q4 | | |
|---|---|---|
| Variable | Coefficient | Standard Error |
| $\Delta\,(\text{median inflation})_{t-1}$ | -0.648 | 0.097 |
| $\Delta\,(\text{median inflation})_{t-2}$ | -0.308 | 0.097 |
| $x_{t-1}$ | 0.122 | 0.035 |
| $x_{t-4}$ | -0.071 | 0.034 |
| Diagnostic Tests | | |
| AR 1-5 test: F(5,79) | 1.210 [0.312] | |
| ARCH 1-4 test: F(4,80) | 0.549 [0.700] | |
| Normality test: Chi^2(2) | 0.630 [0.729] | |
| Hetero test: F(8,79) | 0.701 [0.689] | |
| Hetero-X test: F(14,73) | 0.946 [0.514] | |
| RESET23 test: F(2,82) | 0.976 [0.380] | |

The values in brackets denote probability values (p-values).

Table 2.10: Restricted ADL Phillips Curve, using median inflation (U.S.)

| Null Hypothesis: $(\beta_1 + \beta_4)\,/\,(1 - \alpha_1 - \alpha_2) = 0$ | | | |
|---|---|---|---|
| | value | d.f. | Probability |
| Chi-square test statistic | 4.563 | 1 | 0.032 |

Table 2.11: Wald Test of Coefficient Restrictions when median inflation is used (U.S.)

Motivated by the NKPC literature, we also modified the ADL Phillips curve by replacing the output gap with the real marginal cost, proxied by the labor share of income (see Figure 2.4).[19] The results of the estimated restricted ADL Phillips curve are displayed on Table 2.13 (for headline inflation) and Table 2.15 (for median inflation). In none of these cases the labor share of income appears to be a (statistically) relevant variable for explaining inflation dynamics. Furthermore, the signs of inflation acceleration are consistent only with the APC. Once again, the diagnostic tests indicate that the estimated model performs better when median inflation is used.

| Null Hypothesis: Labor Share is exogenous | | | |
|---|---|---|---|
| | value | d.f. | Probability |
| Difference in J-statistics | 1.397 | 1 | 0.237 |

Table 2.12: Endogeneity Test for Labor Share (U.S.)

[19] Most papers in the recent literature on the NKPC have referred to labor's share of income as real unit labor costs.

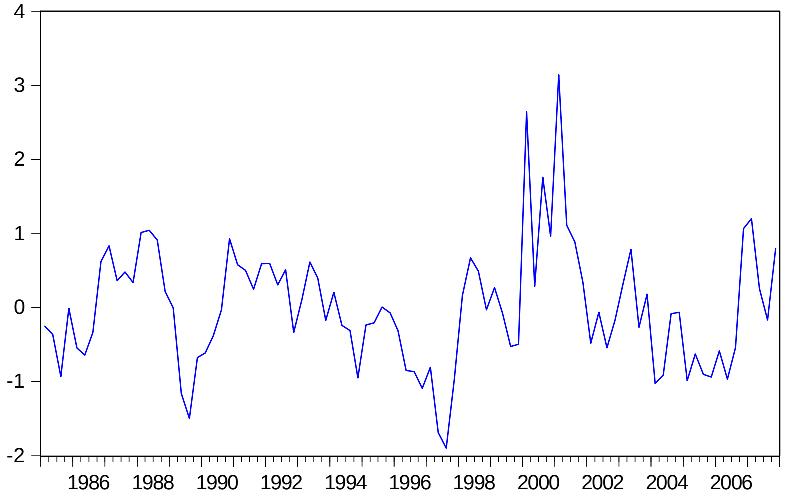



Figure 2.4: Labor Share of Income (U.S.)

| Dependent Variable: $\Delta\pi_t$ | | |
|---|---|---|
| Estimation Method: OLS | | |
| sample: 1985Q1–2007Q4 | | |
| Variable | Coefficient | Standard Error |
| $\Delta\pi_{t-1}$ * | -0.582 | 0.101 |
| $\Delta\pi_{t-2}$ * | -0.437 | 0.098 |
| $(labor\ share)_{t-2}$ | -0.099 | 0.192 |
| $\pi_{t-1}$ | -0.041 | 0.189 |
| Diagnostic Tests | | |
| AR 1-5 test: F(5,80) | 2.429 [0.042]* | |
| ARCH 1-4 test: F(4,81) | 1.953 [0.109] | |
| Normality test: Chi^2(2) | 9.421 [0.009]** | |
| Hetero test: F(8,80) | 1.272 [0.269] | |
| Hetero-X test: F(14,74) | 1.207 [0.288] | |
| RESET23 test: F(2,83) | 0.417 [0.248] | |

The values in brackets denote probability values (p-values).
Significant variables and diagnostic tests at 5% level are shown by one star.

Table 2.13: Restricted ADL Phillips Curve using labor income share (U.S.)

| Null Hypothesis: Labor Share is exogenous | | | |
|---|---|---|---|
| | value | d.f. | Probability |
| Difference in J-statistics | 0.002 | 1 | 0.957 |

Table 2.14: Endogeneity Test for Labor Share when median inflation is used

| Dependent Variable: $\Delta\,(\text{median inflation})_t$ | | |
|---|---|---|
| Estimation Method: OLS | | |
| sample: 1985Q1–2007Q4 | | |
| Variable | Coefficient | Standard Error |
| $\Delta\,(\text{median inflation})_{t-1}\,*$ | -0.582 | 0.100 |
| $\Delta\,(\text{median inflation})_{t-2}\,*$ | -0.229 | 0.100 |
| $(labor\ share)_{t-2}$ | -0.042 | 0.042 |
| Diagnostic Tests | | |
| AR 1-5 test: F(5,81) | 0.451 [0.810] | |
| ARCH 1-4 test: F(4,81) | 1.115 [0.355] | |
| Normality test: Chi^2(2) | 1.111 [0.573] | |
| Hetero test: F(6,82) | 0.793 [0.577] | |
| Hetero-X test: F(9,79) | 1.114 [0.362] | |
| RESET23 test: F(2,84) | 0.030 [0.970] | |

The values in brackets denote probability values (p-values).
Significant variables at 5% level are shown by one star.

Table 2.15: Restricted ADL Phillips Curve using median inflation and labor income share (U.S.)

## 2.5 Empirical Evidence: Brazil (1996Q1–2012Q2)

We extend the unrestricted ADL Phillips curve [equation (2.13)] to introduce the exchange rate, an important open-economy modelling feature.

Including the exchange rate in the study of inflation dynamics is important because the exchange rate allows additional channels for the transmission of monetary policy. In an open economy, the real exchange rate will affect the relative price between domestic and foreign goods, which, in turn, will affect both domestic and foreign demand for domestic goods, and hence contribute to the aggregate-demand channel for the transmission of monetary policy. There is also a direct exchange rate channel for the transmission of monetary policy to inflation, in that the exchange rate affects domestic currency prices of imported final goods, which enter the consumer price index (CPI) and hence CPI inflation. Finally, there is an additional exchange rate channel to inflation: the exchange rate will affect the domestic currency prices of imported intermediate inputs, affecting the cost of domestically produced goods and hence domestic inflation (inflation in the prices of domestically produced goods).

In order to capture the importance of the exchange rate for inflation dynamics we supplement the unrestricted ADL Phillips curve [equation (2.13)] with a 'level' term in the real exchange rate gap $(q_t - \overline{q}_t)$:

$$\Delta\pi_t = \sum_{s=1}^{m}\alpha_s\Delta\pi_{t-s} + \sum_{s=0}^{n}\beta_s x_{t-s} + \gamma_1\pi_{t-1} + \sum_{s=0}^{p}\delta_s\left(q_{t-s} - \overline{q}_{t-s}\right) + \zeta D_t + \varepsilon_t, \tag{2.14}$$

where $D_t$ is a vector of intervention dummies.

For Brazil, the sample goes from 1996Q1 to 2012Q2, the period following the Real Plan. Figures 2.5–2.7 plot the Brazilian inflation rate, the output gap, and the real exchange rate gap.[20] Table 2.16 shows that inflation is stationary at the 1% level.

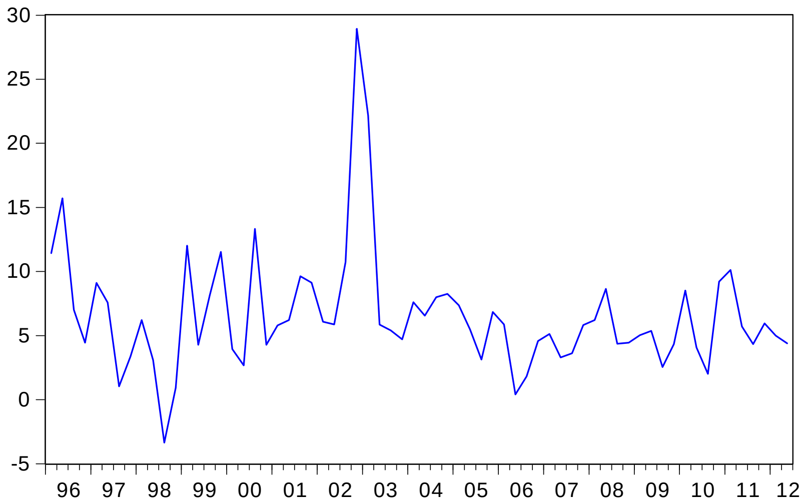



Figure 2.5: Inflation (Brazil)

[20] The Data Appendix gives details on the definitions of the variables used in the baseline model and the robustness analysis.

| Null Hypothesis: Inflation has a unit root<br>Exogenous: Constant<br>Lag length: 0 (Spectral OLS AR based on SIC, maxlag=10)<br>Sample: 1996Q1 2012Q2 | | |
|---|---|---|
| | | P-Statistic |
| Elliott-Rothenberg-Stock Point Optimal test statistic | | 1.563 |
| Test critical values: | 1% level | 1.895 |
| | 5% level | 3.014 |
| | 10% level | 3.993 |

Table 2.16: Unit Root Test for Inflation (Brazil)

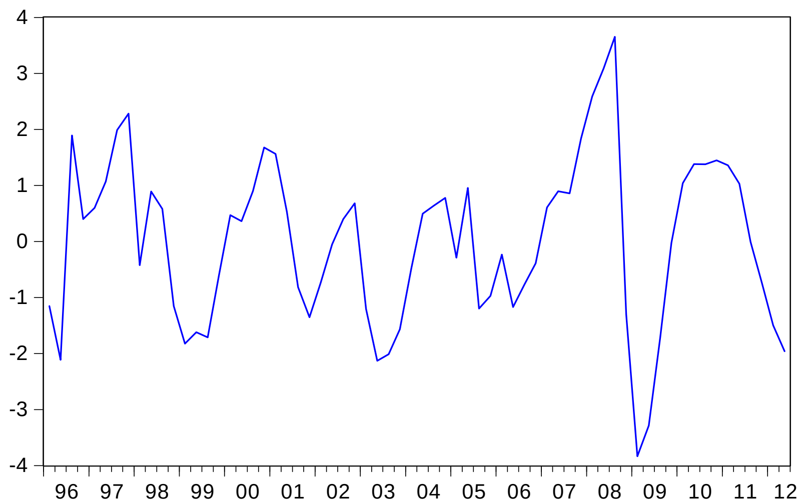



Figure 2.6: Output Gap (Brazil)

Our empirical analysis starts with the estimation of the unrestricted open-economy ADL Phillips curve [equation (2.14)]. The endogeneity test displayed on table 2.17, indicates that both $x_t$ and $(q_t - \overline{q}_t)$ can be treated as exogenous variables.[21] Then, we estimate the unrestricted ADL Phillips curve using OLS, setting $m = n = p = 4$ [see table 2.18].

| Null Hypothesis: Output Gap and Real Exchange Rate Gap are exogenous | | | |
|---|---|---|---|
| | value | d.f. | Probability |
| Difference in J-statistics | 1.721 | 2 | 0.422 |

Table 2.17: Endogeneity Test for Output Gap and Real Exchange Rate Gap (Brazil)

[21] Additional exogeneity tests (not displayed) suggest also that one-period lagged regressors are exogenous.

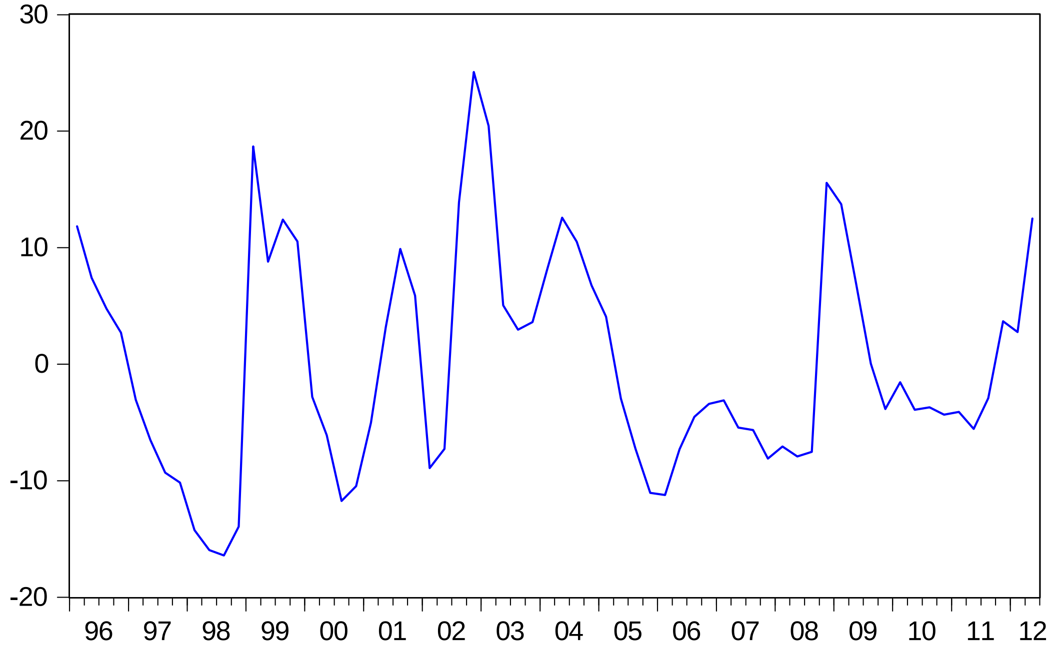



Figure 2.7: Real Exchange Rate Gap (Brazil)

| Dependent Variable: $\Delta\pi_t$ | | |
|---|---|---|
| Estimation Method: OLS | | |
| sample: 1996Q1–2012Q2 | | |
| Variable | Coefficient | Standard Error |
| $\Delta\pi_{t-1}$ | -0.532 | 0.151 |
| $\Delta\pi_{t-2}$ | -0.662 | 0.128 |
| $\Delta\pi_{t-3}$ | -0.263 | 0.141 |
| $x_t$ | 1.148 | 0.527 |
| $x_{t-1}$ | -0.678 | 0.637 |
| $x_{t-2}$ | 0.698 | 0.579 |
| $x_{t-3}$ | -0.591 | 0.520 |
| $x_{t-4}$ | -0.487 | 0.455 |
| $(q_t - \overline{q}_t)$ | 0.143 | 0.080 |
| $(q_{t-1} - \overline{q}_{t-1})$ | -0.034 | 0.102 |
| $(q_{t-2} - \overline{q}_{t-2})$ | 0.080 | 0.101 |
| $(q_{t-3} - \overline{q}_{t-3})$ | -0.121 | 0.099 |
| $(q_{t-4} - \overline{q}_{t-4})$ | -0.151 | 0.095 |
| $\pi_{t-1}$ | -0.166 | 0.068 |
| lula | 9.040 | 2.736 |
| apagão | 2.151 | 2.098 |
| Diagnostic Tests | | |
| AR 1-5 test: F(4,42) | 1.216 [0.318] | |
| ARCH 1-4 test: F(4,54) | 0.314 [0.867] | |
| Normality test: Chi^2(2) | 4.294 [0.116] | |
| Hetero test: F(30,31) | 2.329 [0.011]* | |
| RESET23 test: F(2,44) | 1.371 [0.264] | |

The values in brackets denote probability values (p-values).
Significant diagnostic tests at 5% level are shown by one star.

Table 2.18: Unrestricted ADL Phillips Curve (Brazil)

Next, we reduce the unrestricted ADL Phillips curve by sequential elimination of regressors based on t-ratios of the parameters estimators. We sequentially delete those regressors with the smallest t-ratios and re-estimate the model, until the t-ratios are all greater than the threshold $\gamma = 2$. The estimated restricted ADL model is displayed on table 2.19.

Using equation (2.14) and comparing the values of tables 2.1 and 2.19, we observe that all restrictions implied by the NKPC ($\alpha_1 = 0$, $\beta_0 = 0$, $\beta_1 < 0$, $\gamma_1 > 0$) and the HPC ($\alpha_1 > 0$, $\beta_0 = 0$, $\beta_1 < 0$, $\gamma_1 > 0$) are rejected by the Brazilian data. The SIPC gets one restriction correct ( $\beta_0 > 0$) and three wrong ($\alpha_1 = 0$, $\beta_1 < 0$, $\gamma_1 = 0$). The APC gets three restrictions correct ($\alpha_1 < 0$, $\beta_0 > 0$, $\beta_1 = 0$) and one incorrect ($\gamma_1 < 0$).[22] Therefore, in a small open economy like Brazil only the APC model is consistent with inflation dynamics from 1996Q1–2012Q2. However, the diagnostic tests of Table 2.19 indicate some misspecification problems with the estimated model, as there is evidence of heteroscedasticity in the residuals.

| Dependent Variable: $\Delta\pi_t$<br>Estimation Method: OLS<br>sample: 1996Q1–2012Q2 | | |
|---|---|---|
| Variable | Coefficient | Standard Error |
| $\Delta\pi_{t-1}$ | -0.582 | 0.123 |
| $\Delta\pi_{t-2}$ | -0.666 | 0.096 |
| $\Delta\pi_{t-3}$ | -0.328 | 0.106 |
| $x_t$ | 0.843 | 0.359 |
| $x_{t-4}$ | -0.741 | 0.357 |
| $(q_t - \overline{q}_t)$ | 0.134 | 0.060 |
| $\left(q_{t-4} - \overline{q}_{t-4}\right)$ | -0.224 | 0.059 |
| $\pi_{t-1}$ | -0.155 | 0.062 |
| $lula$ | 9.334 | 2.419 |
| Diagnostic Tests | | |
| AR 1-5 test: F(4,49) | 0.083 [0.987] | |
| ARCH 1-4 test: F(4,54) | 0.826 [0.514] | |
| Normality test: Chi^2(2) | 4.070 [0.130] | |
| Hetero test: F(12,49) | 4.237 [0.000]** | |
| Hetero-X test: F(27,34) | 2.532 [0.023]* | |
| RESET23 test: F(2,54) | 1.073 [0.349] | |

The values in brackets denote probability values (p-values).
Significant diagnostic tests at 5% level are shown by one star.
Significant diagnostic tests at 1% level are shown by two stars.

Table 2.19: Restricted ADL Phillips Curve (Brazil)

*Robustness*

In order to test the robustness of the results, we estimate the ADL Phillips curve using core inflation instead of headline inflation (see Figure 2.8). Tables 2.20 and 2.21 show that

[22] However, remember from section 3 that if we do not impose the assumption of vertical long-run Phillips curve, the APC implies that $\gamma_1 < 0$.

core inflation is difference stationary, which implies that the coefficient of lagged inflation in equation (2.14) is equal to zero ($\gamma_1 = 0$), while $t$-tests are still appropriate to evaluate the other (stationary) variables. Table 2.23 displays the results of the estimated restricted ADL Phillips curve. With the exception of lagged inflation, the selected variables are the same that those when headline inflation is used. We reach the same conclusion as before: only the APC is consistent with inflation dynamics in Brazil. However, the diagnostic tests still indicate some heteroscedasticity problems.

Given that the Brazilian economic policy changed in January 1999, when the exchange rate was allowed to float and in June 1999, when inflation targeting was implemented, we tested the stability of our results by re-estimating the ADL model for two different subsamples: 1998Q2–2012Q2 and 1999Q3–2012Q2. Not only the restricted models are equal to the 1996Q1–2012Q2 sample, but the parameters estimates are quite stable. None of the conclusions changes.

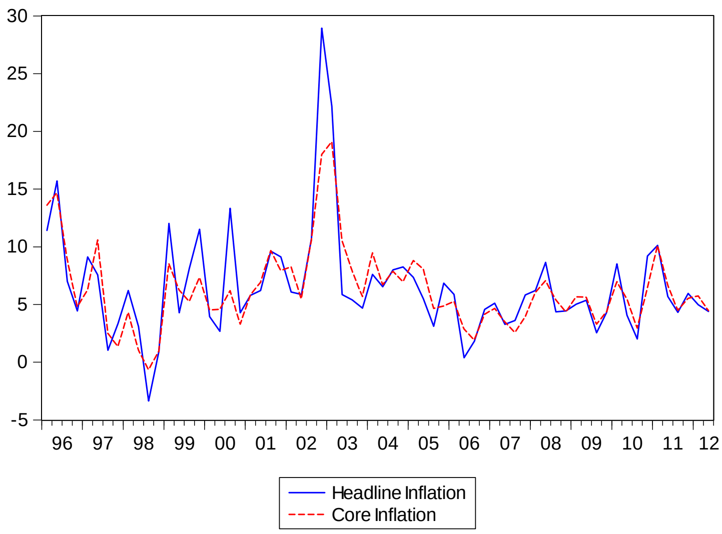



Figure 2.8: Core Inflation x Headline Inflation (Brazil)

| | | |
|---|---|---|
| Null Hypothesis: Core Inflation has a unit root | | |
| Exogenous: Constant | | |
| Lag length: 0 (Spectral OLS AR based on SIC, maxlag=10) | | |
| Sample: 1996Q1 2012Q2 | | |
| | | P-Statistic |
| Elliott-Rothenberg-Stock Point Optimal test statistic | | 4.820 |
| Test critical values: | 1% level | 1.895 |
| | 5% level | 3.014 |
| | 10% level | 3.993 |

Table 2.20: Unit Root Test for Core Inflation (Brazil)

| Null Hypothesis: $\Delta$ (core inflation)$_t$ has a unit root<br>Exogenous: Constant<br>Lag length: 1 (Spectral OLS AR based on SIC, maxlag=10)<br>Sample: 1996Q1–2012Q2 | | |
|---|---|---|
| | | P-Statistic |
| Elliott-Rothenberg-Stock Point Optimal test statistic | | 0.567 |
| Test critical values: | 1% level | 1.895 |
| | 5% level | 3.014 |
| | 10% level | 3.993 |

Table 2.21: Unit Root Test for the first difference of Core Inflation (Brazil)

| Null Hypothesis: Output Gap and Real Exchange Rate Gap are exogenous | | | |
|---|---|---|---|
| | value | d.f. | Probability |
| Difference in J-statistics | 2.263 | 2 | 0.322 |

Table 2.22: Endogeneity Test for Output Gap and Real Exchange Rate Gap (Brazil)

| Dependent Variable: $\Delta(\textit{core inflation})_t$<br>Estimation Method: OLS<br>sample: 1996Q1–2012Q2 | | |
|---|---|---|
| Variable | Coefficient | Standard Error |
| $\Delta(\textit{core inflation})_{t-1}$ | -0.473 | 0.107 |
| $\Delta(\textit{core inflation})_{t-2}$ | -0.605 | 0.091 |
| $\Delta(\textit{core inflation})_{t-3}$ | -0.258 | 0.095 |
| $x_t$ | 0.486 | 0.234 |
| $x_{t-4}$ | -0.437 | 0.216 |
| $(q_t - \overline{q}_t)$ | 0.105 | 0.038 |
| $(q_{t-4} - \overline{q}_{t-4})$ | -0.150 | 0.034 |
| *lula* | 4.089 | 1.361 |
| Diagnostic Tests | | |
| AR 1-5 test: F(4,50) | 0.781 [0.542] | |
| ARCH 1-4 test: F(4,54) | 2.909 [0.029]* | |
| Normality test: Chi^2(2) | 0.545 [0.761] | |
| Hetero test: F(15,46) | 1.041 [0.433] | |
| Hetero-X test: F(36,25) | 2.297 [0.016]* | |
| RESET23 test: F(2,52) | 2.583 [0.085] | |

The values in brackets denote probability values (p-values).
Significant diagnostic tests at 5% level are shown by one star.

Table 2.23: Restricted ADL Phillips Curve using core inflation (Brazil)

We also modified the ADL Phillips curve by replacing the output gap with the real marginal cost, proxied by the labor share of income (see Figure 2.9). The results of the estimated restricted ADL Phillips curve are displayed on Table 2.25 (for headline inflation) and Table 2.27 (for core inflation). In both cases, labor share of income is not a (statistically) relevant variable and the APC is still the only model consistent with inflation dynamics.

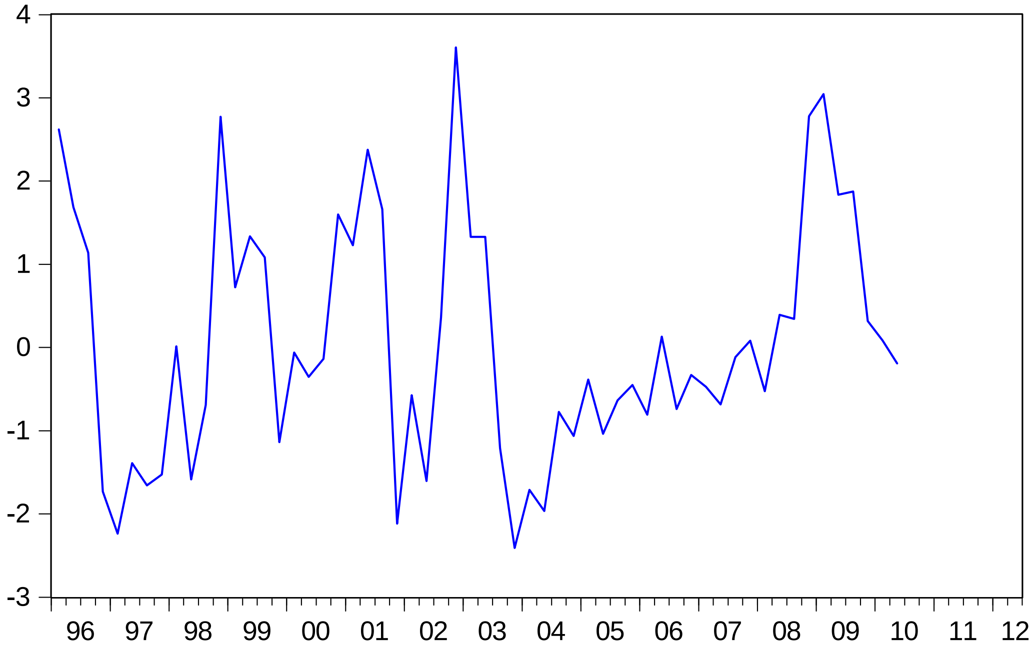



Figure 2.9: Labor Share of Income (Brazil)

| Null Hypothesis: Labor Share and Real Exchange Rate Gap are exogenous | | | |
|---|---|---|---|
| | value | d.f. | Probability |
| Difference in J-statistics | 1.392 | 2 | 0.498 |

Table 2.24: Endogeneity Test for Labor Share and Real Exchange Rate Gap when headline inflation is used

| Dependent Variable: $\Delta\pi_t$<br>Estimation Method: OLS<br>sample: 1996Q1–2010Q2 | | |
|---|---|---|
| Variable | Coefficient | Standard Error |
| $\Delta\pi_{t-1}*$ | -0.340 | 0.107 |
| $\Delta\pi_{t-2}*$ | -0.512 | 0.098 |
| $(labor\ share)_{t-2}$ | -0.585 | 0.364 |
| $(q_t - \overline{q}_t)*$ | 0.136 | 0.058 |
| $\left(q_{t-3} - \overline{q}_{t-3}\right)*$ | -0.129 | 0.053 |
| $\pi_{t-1}*$ | -0.157 | 0.074 |
| $lula*$ | 7.206 | 2.734 |
| Diagnostic Tests | | |
| AR 1-5 test: F(4,46) | 0.822 [0.517] | |
| ARCH 1-4 test: F(4,49) | 0.464 [0.761] | |
| Normality test: Chi^2(2) | 1.663 [0.435] | |
| Hetero test: F(13,43) | 2.981 [0.003]** | |
| Hetero-X test: F(28,28) | 4.620 [0.000]** | |
| RESET23 test: F(2,48) | 1.732 [0.187] | |

The values in brackets denote probability values (p-values).
Significant variables at 5% level are shown by one star.
Significant diagnostic tests at 1% level are shown by two stars.

Table 2.25: Restricted ADL Phillips Curve using labor share (Brazil)

| Null Hypothesis: Labor Share and Real Exchange Rate Gap are exogenous | | | |
|---|---|---|---|
| | value | d.f. | Probability |
| Difference in J-statistics | 4.701 | 2 | 0.095 |

Table 2.26: Endogeneity Test for Labor Share and Real Exchange Rate Gap when core inflation is used

| Dependent Variable: $\Delta(core\ inflation)_t$ | | |
|---|---|---|
| Estimation Method: OLS | | |
| sample: 1996Q1–2010Q2 | | |
| Variable | Coefficient | Standard Error |
| $\Delta(core\ inflation)_{t-1}*$ | -0.445 | 0.120 |
| $\Delta(core\ inflation)_{t-2}*$ | -0.558 | 0.106 |
| $\Delta(core\ inflation)_{t-3}*$ | -0.299 | 0.109 |
| $(labor\ share)_{t-2}$ | 0.399 | 0.085 |
| $(q_t - \overline{q}_t)*$ | 0.073 | 0.035 |
| $(q_{t-4} - \overline{q}_{t-4})*$ | -0.083 | 0.036 |
| lula* | 4.075 | 1.516 |
| Diagnostic Tests | | |
| AR 1-5 test: F(4,43) | 1.756 [0.155] | |
| ARCH 1-4 test: F(4,46) | 0.921 [0.459] | |
| Normality test: Chi^2(2) | 1.733 [0.420] | |
| Hetero test: F(13,40) | 1.188 [0.322] | |
| Hetero-X test: F(28,25) | 2.350 [0.017]* | |
| RESET23 test: F(2,45) | 3.703 [0.032]* | |

The values in brackets denote probability values (p-values).
Significant variables at 5% level are shown by one star.
Significant diagnostic tests at 1% level are shown by two stars.

Table 2.27: Restricted ADL Phillips Curve using core inflation and labor share (Brazil)

## 2.6 Concluding Remarks

There is no consensus on two key Phillips curves issues. First, are inflation expectations forward-looking or backward-looking? If expectations are forward-looking, future events (including changes in monetary policy) can influence the current inflation rate. If, instead, expectations are backward-looking, inflation has inertia. Such inertia affects the design of monetary policy. Second, what is a proper measure of inflationary pressures—the output gap or the real marginal cost?

In this paper we test Phillips curves using an ecompassing framework. We conclude that the NKPC does not provide a useful description of the inflation process in the cases of the U.S. and Brazil. The evidence presented here rejects the restrictions implied by the NKPC, the HPC, and SIPC, but does not reject those of the APC.

Regarding the two fundamental Phillips curves issues, our results suggest that inflation expectations are backward-looking and that the output gap is a better measure of inflationary pressures than the real marginal cost.

## 2.A Data Appendix

U.S. Data

The inflation rate ($\pi_t$) is measured as the quarter-to-quarter, annualized growth rate of the Consumer Price Index for All Urban Consumers (All Items).

The output gap ($x_t$) is given by 100 times the difference between the log of the quarterly real GDP and the log of the Congressional Budget Office's estimate of the potential real GDP.

Median inflation is the annualized percentage change of the quarterly median inflation reported by the Federal Reserve Bank of Cleveland.

Our measure of labor share of income (equivalently, real unit labor cost) corresponds to the GG measure of log-real unit cost as described by King and Watson (2012), with the difference that we applied the Hodrick-Prescott (HP) filter to get the deviation from steady-state.

The capacity utilization gap is total industry capacity utilization, detrended by the HP filter.

Brazilian Data

The inflation rate ($\pi_t$) is measured as the quarter-to-quarter, annualized growth rate of the Consumer Price Index (IPCA).

The output gap ($x_t$) is given by 100 times the log of the quarterly real GDP seasonally adjusted, detrended by the HP filter.

The real exchange rate gap ($q_t - \overline{q}_t$) is calculed as 100 times the log of the quarterly average of monthly indexes of real effective exchange rates, detrended by the HP filter.

We employed two intervention dummies ($D_t$). One is "Lula", which tries to capture the effect of the election of President Lula on inflation (2002Q3– 2003Q1). The other "Apagão", which tries to capture the negative effect of the energy crisis on the output gap (2001Q3–2002Q1).

Core inflation is the annualized percentage change of the quarterly exclusion core inflation reported by the Central Bank of Brazil. (In the exclusion core inflation, food, energy and regulated & administered prices are excluded from the measurement of inflation).

Our measure of labor share of income is based on Bastos (2012), which ends in 2010Q2. To his series we applied logs, seasonally adjusted it, and applied the HP filter.

The capacity utilization gap is general installed capacity utilization, detrended by the HP filter.

# Chapter 3

# IS Curves: An Encompassing Test

## 3.1 Introduction

The New Keynesian model has become the workhorse framework for monetary policy analysis. In its simplest form, it is a three-equation model composed by the New Keynesian Phillips curve (NKPC); the New Keynesian IS curve (NKISC); and an interest rate rule. A key difference between the traditional Keynesian IS curve (KISC) and the NKISC is that the KISC determines the *level* of output gap, while the NKISC determines its *rate of change*.[1] The NKPC has been subject of extensive empirical investigation.[2] In contrast, few studies have analysed the empirical performance of the NKISC.

This paper tests IS curves using an autoregressive distributed lag (ADL) specification that encompasses the traditional Keynesian IS curve (KISC), the New Keynesian IS curve (NKISC), and the Hybrid IS curve (HISC). We use data from the United States (1985Q1–2007Q4) and from Brazil (1996Q1–2012Q2).[3] The evidence rejects the restrictions implied by the NKISC and the HISC, but does not reject those of the KISC.

Rudebusch and Svensson (1999) and Peersman and Smets (1999) estimate Keynesian (backward-looking) IS curves for the US and the Euro Area respectively and both find a significantly negative real interest rate semi-elasticity of around 0.1. Rotemberg and Woodford (1999) are among the first to estimate an output Euler equation in the context of an optimization based macroeconomic model in which there is no weight on past inflation. However, Estrella and Fuhrer (2002) show that the NKISC provides a remarkably poor fit to the time series data on aggregate output. Furthermore, Fuhrer and Rudebusch (2004) in an empirical analysis of the NKISC find that the estimated real interest rate semi-elasticity is not significantly different from zero. This result poses a serious challenge for standard theory, which predicts that this real interest rate elasticity is unambiguously negative. There is growing evidence that purely forward-looking specifications generate counterfactual dynamics for output. Some adjustment process must be added to the structural equations in order to match the inertial responses of output that are apparent in the data. Hybrid IS curve (HISC) specifications, with both forward- and backward-looking components seem better suited to characterize actual dynamics. Fuhrer (2000)

[1] For a critical review of the New Keynesian model and a discussion about the distinction between the KISC and NKISC see Barbosa (2013b).

[2] See, for example, Galí and Gertler (1999), Galí, Gertler and López-Salido (2001), Rudd and Whelan (2007).

[3] Given that we estimate reduced-form models, the choice of the sample in both countries is motivated by the attempt of avoiding the Lucas critique by selecting periods of economic regime stability.

obtains better empirical results by enhancing the model of consumer behavior with a habit formation process that adds significant inertial output dynamics.

Carneiro and Wu (2003) estimate a KISC for Brazil and when they include only data after the beginning of the Real Plan, they fail to capture the influence of some important variables. Santos and Holland (2008) estimate different specifications of a NKISC curve for the Brazilian economy in which they include external variables like the global aggregate demand, real exchange rate and the terms of trade. They find that the real exchange rate is not significant for the aggregate demand behavior. Carvalho (2008) estimates an HISC for Brazil and finds that the real interest rate semi-elasticity is either positive or not significant, depending on the estimator used.

The paper is organized as follows. Section 2 reviews the literature on IS curves. Section 3 presents the ADL IS curve and shows how the different IS curve specifications considered in the literature can be seen as special cases of the ADL IS curve. Section 4 tests for the U.S. the restrictions implied by the different IS curve specifications. Section 5 estimates the ADL IS curve for Brazil and tests the fit of the alternative models. Finally, section 6 brings the concluding remarks.

## 3.2 IS Curves

This section reviews three approaches followed in the literature to model the relationship between real interest rate and output in the short-run, the IS curve. In the Keynesian approach, the IS equation is motivated by behavior relationships that are not based on optimization models. In the New Keynesian approach, the IS equation is derived from an optimizing framework. The Hybrid approach adds habit formation to the New Keynesian model, a feature that improves the short-run dynamics of the model. Each approach provides different IS curve specifications, which can be tested empirically.

### 3.2.1 Keynesian IS Curve (KISC)

The output market is in equilibrium when real domestic output, $Y$, equals the aggregate demand for domestic output

$$Y = C(Y - \tau) + I(r) + G + NX(Q) \tag{3.1}$$

where $C$ is consumption, $I$ is investment, $G$ are government purchases, $NX$ are net exports of goods and services, $\tau$ denotes net taxes collected by the government, $r$ is the real interest rate, and $Q$ is the real exchange rate.

Linearization of equation (3.1) around the steady-state $(\overline{Y}, \overline{I}, \overline{G}, \overline{NX}, \overline{r}, \overline{\tau}, \overline{Q})$ gives us[4]

$$x = -\alpha(r - \overline{r}) + \beta(q - \overline{q}) + \gamma(f - \overline{f}) + (g - \overline{g}), \tag{3.2}$$

where $x \equiv \log Y - \log \overline{Y} \cong \left(Y - \overline{Y}\right)/\overline{Y}$ is the output gap, $r - \overline{r}$ is the real interest rate gap (the deviation of the real interest rate from its steady-state value), $q - \overline{q} \equiv \log Q - \log \overline{Q} \cong \left(Q - \overline{Q}\right)/\overline{Q}$ is the real exchange rate gap (the deviation of the real exchange rate from its steady-state value), $f = (g - \tau)/\overline{Y}$ is the public deficit over potential output, $f - \overline{f}$ is the public deficit gap (the deviation of public deficit-to-potential output from its steady-state

[4] For the step-by-step derivation see Barbosa (2013a).

value), $g = G/\overline{Y}$, $g - \overline{g}$ is the government expenditure gap (the deviation of government expenditure-to-potential output from its steady-state value).

According to equation (3.2), the output gap depends on the gaps of the real interest rate, the real exchange rate, the public deficit, and the government expenditures.

### 3.2.2 New Keynesian IS Curve (NKISC)

The New Keynesian model incorporates nominal wage or price rigidity into a dynamic stochastic general equilibrium (DSGE) framework that is based on optimizing behavior by the agents in the model. These DSGE models with nominal frictions are labeled New Keynesian models because, like older versions of models in the Keynesian tradition, prices are sticky and aggregate demand plays a central role in determining output in the short run, and there is a presumption that some fluctuations both can be and should be dampened by countercylical monetary or fiscal policy. Within the model, monetary policy affects the real economy in the short run, much as in the traditional Keynesian IS/LM framework. A key difference, however, is that the aggregate behavioral equations evolve explicitly from optimization by households and firms. One important implication is that current economic behavior depends critically on expectations of the future course of monetary policy as well as on current policy.

It is possible to represent the canonical New Keynesian model in terms of two equations: an "IS" curve that relates the output gap negatively to the real interest rate; and a Phillips curve that relates inflation positively to the output gap. In what follows, we show how the NKISC can be derived from the households' optimization problem.[5]

In the canonical New Keynesian model the preferences of the representative household are defined over a composite consumption good $C_t$,[6] real money balances $M_t/P_t$, and the time devoted to market employment $N_t$. Households maximize the expected present discounted value of utility:

$$E_0 \sum_{t=0}^{\infty} \beta^t \left[ \frac{C_t^{1-1/\sigma}}{1-1/\sigma} + \frac{\gamma}{1-b} \left( \frac{M_t}{P_t} \right)^{1-b} - \kappa \frac{N_t^{1+\chi}}{1+\chi} \right], \tag{3.3}$$

subject to the budget constraint

$$C_t + \frac{M_t}{P_t} + \frac{B_t}{P_t} = \frac{W_t}{P_t} N_t + \frac{M_{t-1}}{P_t} + \frac{(1+i_{t-1})\, B_{t-1}}{P_t} + \frac{D_t}{P_t}$$

where $B_t$ and $D_t$ denote household bond holdings and dividends revenues, respectively, $W_t$ denotes the nominal wage, and $i_t$ denotes the nominal return on bonds.

The first-order conditions with respect to $C_t$ and $B_t$ are:

$$C_t^{-1/\sigma} - \lambda_t = 0 \tag{3.4}$$

$$-\lambda_t + \beta E_t \lambda_{t+1} (1+i_t) = 0 \tag{3.5}$$

where $\lambda_t$ denotes the Lagrange multiplier.

[5] For an exposition of the New Keynesian model see Clarida, Galí, and Gertler (1999). For a detailed textbook presentation see Galí (2008), Walsh (2010), or Woodford (2003).

[6] The composite consumption good consists of differentiated products produced by monopolistically competitive final goods producers.

We can use equations (3.4) and (3.5) to derive the Euler equation for the optimal intertemporal allocation of consumption which must hold in equilibrium:

$$C_t^{-1/\sigma} = \beta E_t \left[ C_{t+1}^{-1/\sigma} \frac{(1+i_t)}{(1+\pi_{t+1})} \right] \tag{3.6}$$

where $\pi_{t+1} = (P_{t+1}/P_t) - 1$

Because consumption is equal to output in this model (there is no government or investment because capital has been ignored), (3.6) can be log-linearized around the steady-state to give us an expression for the NKISC:

$$x_t = E_t x_{t+1} - \sigma(r_t - \overline{r}_t) \tag{3.7}$$

where $x_t \equiv y_t - \overline{y}_t$ is the output gap, $y_t = \log Y_t$, $\overline{y}_t = \log \overline{Y}_t$, $r_t = i_t - E_t\pi_{t+1}$, $\overline{r}_t = \rho + \frac{1}{\sigma}\Delta\overline{y}_{t+1}$, and $\rho = (1-\beta)/\beta$.

### 3.2.3 Hybrid IS Curve (HISC)

A stylized fact about the monetary policy transmission mechanism is the delayed response of spending to policy actions, with the peak response at about one year and the entire effect lasting two years or more. The New Keynesian model implies instead that spending act like a jump variable, pulling forward in time in response to shocks. The challenge is to build models that imply a gradual and hump-shaped response of spending to monetary policy shocks. Fuhrer (2000) shows that including habit formation in the consumer's utility function allows the model to replicate key dynamic correlations among consumption, output, interest rates, and inflation to a degree the standard New Keynesain model cannot. In particular, the augmented model can match the hump-shaped response of consumption to income, interest rate, and inflation shocks. The habit formation specification improves upon the standard specification because it imparts a motive for consumers to smooth the change, as well as the level of consumption. We show below how the introduction of habit formation allows us to get an expression for the HISC.[7]

Assuming that the utility function is of the isoelastic form and that the habit stock enters the utility function additively, that is, utility depends on the difference between consumption and the habit stock, the period utility function is

$$u(C_t, H_t) = \frac{(C_t - H_t)^{1-1/\sigma}}{1-1/\sigma}, \tag{3.8}$$

where $H_t$ evolves according to

$$H_t = \eta \mathbb{C}_{t-1}, \tag{3.9}$$

where $\mathbb{C}_{t-1}$ is aggregate consumption at *t-1* and $\eta \in [0, 1)$.

The first-order conditions of the households' maximization problem with respect to $C_t$ and $B_t$ are now:

$$(C_t - \eta C_{t-1})^{-1/\sigma} - \lambda_t = 0 \tag{3.10}$$
$$-\lambda_t + \beta E_t \lambda_{t+1}(1+r_t) = 0 \tag{3.11}$$

[7]For a detailed analysis of habit formation and its consequences for the business cycles properties of the standard New Keynesian model, see Dennis (2009).

We can use equations ((3.10) and (3.11) to derive the Euler equation

$$(C_t - \eta C_{t-1})^{-1/\sigma} = \beta E_t(1 + r_t)(C_{t+1} - \eta C_t)^{-1/\sigma}. \tag{3.12}$$

Combining equation (3.12) with the market equilibrium condition gives us an expression for the HISC:

$$x_t = \frac{\eta}{1+\eta} x_{t-1} + \frac{1}{1+\eta} E_t x_{t+1} - \frac{\sigma(1-\eta)}{(1+\eta)}(r_t - \overline{r}_t).$$

Thus, the introduction of habit formation in the consumer's utility function makes the output gap in period $t$ dependent on output gap of the previous period yielding inertia in the model.

## 3.3 IS Curve: An Encompassing Specification

A model of aggregate demand general enough that encompasses the KISC, the NKISC, and the HISC as special cases takes the form

$$\Delta x_t = \alpha_1 \Delta x_{t-1} + \beta_0 (r_t - \overline{r}_t) + \beta_1 (r_{t-1} - \overline{r}_{t-1}) + \gamma_1 x_{t-1} + \varepsilon_t, \tag{3.13}$$

where $\varepsilon_t$ is an error term. Let us show how each model is embedded in this equation. The simple KISC can be written as

$$x_t = \lambda x_{t-1} - \alpha(r_t - \overline{r}_t) + \varepsilon_t, \quad 0 < \lambda < 1,$$

which is equivalent to

$$\Delta x_t = -\left(1 - \lambda\right) x_{t-1} - \alpha(r_t - \overline{r}_t) + \varepsilon_t.$$

This is a particular case of equation (3.13) when $\alpha_1 = 0$, $\beta_0 = -\alpha < 0$, $\beta_1 = 0$, $\gamma_1 = -\left(1 - \lambda\right) < 0$.
When the output gap depends on additional output gap lags we have the KISC. For the case in which output gap depends on the last two output gap lags the KISC is given by:

$$x_t = \lambda_1 x_{t-1} + \lambda_2 x_{t-2} - \alpha(r_t - \overline{r}_t) + \varepsilon_t.$$

After some algebra, this expression can be rearranged as

$$\Delta x_t = -\lambda_2 \Delta x_{t-1} - \left(1 - \lambda_1 - \lambda_2\right) x_{t-1} - \alpha(r_t - \overline{r}_t) + \varepsilon_t.$$

This specification is a particular case of equation (3.13) when $\alpha_1 = -\lambda_2 \lesseqgtr 0$, $\beta_0 = -\alpha < 0$, $\beta_1 = 0$, $\gamma_1 = -\left(1 - \lambda_1 - \lambda_2\right) < 0$.[8]

The NKISC is given by

$$x_t = E_t x_{t+1} - \alpha(r_t - \overline{r}_t) + \varepsilon_t.$$

Assuming rational expectations, $E_t x_{t+1} = x_{t+1} - \mu_{t+1}$ , the NKISC can be written as

$$\Delta x_t = \alpha(r_{t-1} - \overline{r}_{t-1}) + \eta_t.$$

[8]The sign of $\alpha_1 = -\lambda_2$ will be determined by the stability condition that the roots of $\left(\phi^2 - \lambda_1 \phi - \lambda_2 = 0\right)$ lie inside the unit circle.

This follows from equation (3.13) when $\alpha_1 = \beta_0 = \gamma_1 = 0$, $\beta_1 = \alpha > 0$.

The HISC is given by

$$x_t = \omega x_{t-1} + (1-\omega)E_t x_{t+1} - \alpha(r_t - \overline{r}_t) + \varepsilon_t.$$

As with the NKISC, we assume rational expectations and after some algebra we obtain

$$\Delta x_t = \frac{\omega}{(1-\omega)}\Delta x_{t-1} + \frac{\alpha}{(1-\omega)}(r_{t-1} - \overline{r}_{t-1}) + \xi_t.$$

This is a particular case of equation (3.13) when $\alpha_1 = \omega/\,(1-\omega) > 0$, $\beta_0 = \gamma_1 = 0$, $\beta_1 = \alpha/\,(1-\omega) > 0$.

Table 3.1 shows the restrictions implied by each IS curve model subsumed by equation (3.13) which is an ADL(1,1). Once the parameters of the model have been estimated–$(\alpha_1, \beta_0, \beta_1, \gamma_1)$–table 3.1 enables us to identify which IS curve model is consistent with the evidence. For example, if the estimated coefficient of $x_{t-1}(\gamma_1)$ is negative, then only the KISC would be consistent with data.

| **Model** | **Parameters** | | | |
|---|---|---|---|---|
| | $\boldsymbol{\alpha_1}$ | $\boldsymbol{\beta_0}$ | $\boldsymbol{\beta_1}$ | $\boldsymbol{\gamma_1}$ |
| **KISC** | **-/0/+** | **-** | **0** | **-** |
| **NKISC** | **0** | **0** | **+** | **0** |
| **HISC** | **+** | **0** | **+** | **0** |

Table 3.1: Model Typology

The ADL IS curve [equation (3.13)] encompasses all IS curves as special cases. The unrestricted ADL IS curve includes additional lags of output gap changes and real interest rate gap, according to:

$$\Delta x_t = \sum_{s=1}^{m} \alpha_s \Delta x_{t-s} + \sum_{s=0}^{n} \beta_s \left(r_{t-s} - \overline{r}_{t-s}\right) + \gamma_1 x_{t-1} + \varepsilon_t. \tag{3.14}$$

In this paper we test the NKISC against the KISC and the HISC based on the encompassing ADL IS curve [equation (3.14)]. The empirical strategy, inspired in the general-to-specific approach to econometrics, is as follows: (i) formulate an IS curve model that encompasses the NKISC, the KISC, and the HISC as special cases—the unrestricted ADL IS curve; (ii) reduce the unrestricted ADL IS curve model based on statistical criteria—the restricted ADL IS curve; (iii) test each IS curve model by comparing its implied restrictions with the values of the estimated parameters of the restricted ADL model.[9]

## 3.4 Empirical Evidence: United States (1985Q1 - 2007Q4)

The U.S. sample goes from 1985Q1 to 2007Q4, the period known as "Great Moderation"due to the decline in the variability of both output and inflation.[10] Figure 3.1

[9] For a survey on the general-to-specific modeling see Campos, Ericsson, and Hendry (2005).

[10] James Stock coined the phrase "the great moderation" while writing a research paper with Mark Watson in 2002 ("Has the Business Cycle Changed and Why?"). It was brought to the attention of the wider public by Ben Bernanke (then member and now chairman of the Board of Governors of the Federal Reserve) in a speech titled "The Great Moderation" in 2004. The results do not change significantly when we start or finish the sample one year earlier.

plots the output gap and real interest rate gap for the U.S.[11,12]

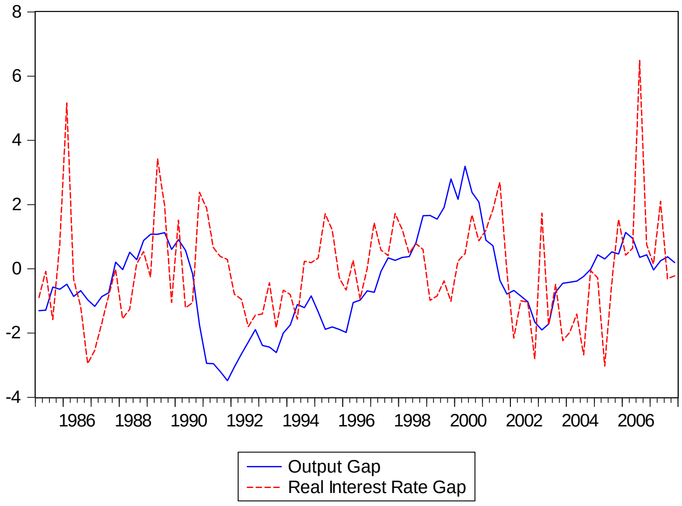



Figure 3.1: Output Gap and Real Interest Rate Gap (U.S.)

Our empirical analysis starts with the estimation of the unrestricted ADL IS curve [equation (3.14)]. Given the presence of the current real interest rate gap $(r_t - \overline{r}_t)$ as a regressor, one may suspect that the Ordinary Least Square (OLS) assumption of orthogonality between regressors and the error term $(\varepsilon_t)$ may not hold. If $(r_t - \overline{r}_t)$ is correlated with $\varepsilon_t$, then the OLS estimates will be biased and inconsistent. A solution to this endogeneity problem lies in the use of Instrumental Variable (IV)/Generalized Method of Moments (GMM) estimators. However, if $(r_t - \overline{r}_t)$ is not correlated with $\varepsilon_t$ and we use IV/GMM estimators, there will be a loss of efficiency, as the asymptotic variance of the IV/GMM is always larger (sometimes much larger) than the asymptotic variance of the OLS estimator. Therefore, before proceeding with the estimation of equation (3.14) we test the endogeneity of $(r_t - \overline{r}_t)$ in order to check the appropriateness of OLS or the necessity to resort to IV/GMM methods.[13] We apply the GMM distance test (also known as C test) of a subset of orthogonality conditions to test the endogeneity of $(r_t - \overline{r}_t)$.[14,15]

[11] The Data Appendix gives details on the definitions of the variables used in the baseline model and the robustness analysis.

[12] Proponents of the New Keynesian model criticize traditional measures of the output gap on the grounds that naive detrending procedures assume that potential GDP evolves smoothly over time. In theory, however, changes in potential output will be affected by any number of shocks and so could fluctuate significantly (and stochastically) period to period. While the points regarding the difficulties surrounding the measurement of potential output are clearly valid and important, these points should not, on their own, be taken as reasons to jeopardize the usefulness of standard output gap measures.

[13] If orthogonality between regressors and the error term cannot be rejected, we should use OLS rather than IV or GMM, especially in small samples.

[14] In our context, the GMM distance test of a subset of orthogonality conditions is equivalent to the Durbin-Wu-Hausman endogeneity test. See Baum, Schaffer, and Stillman (2003, 2007).

[15] The GMM distance test is computed as the difference between two J-statistics: that for the (restricted, fully efficient) regression using the entire set of overidentifying restrictions, versus that for the (unrestricted, inefficient but consistent) regression using a smaller set of restrictions, in which a specified set of instruments is removed from the set. For instruments included in the equation to be estimated, the GMM distance test places them in the list of included endogenous variables; in essence, treating them as endogenous regressors. The GMM distance test, distributed $\chi^2$ with degrees of freedom equal to the

The endogeneity test displayed on table 3.2, indicates that we can treat $(r_t - \overline{r}_t)$ as an exogenous variable.[16] Therefore, we proceed with our empirical analysis by estimating the unrestricted ADL IS curve [equation (3.14)] using OLS, setting $m = n = 4$ [see table 3.3].

| Null Hypothesis: Real Interest Rate Gap is exogenous | | | |
|---|---|---|---|
| | value | d.f. | Probability |
| Difference in J-statistics | 1.370 | 1 | 0.241 |

Table 3.2: Endogeneity Test for Real Interest Rate Gap (U.S.)

| Dependent Variable: $\Delta x_t$<br>Estimation Method: OLS<br>sample: 1985Q1–2007Q4 | | |
|---|---|---|
| Variable | Coefficient | Standard Error |
| $\Delta x_{t-1}$ | 0.127 | 0.110 |
| $\Delta x_{t-2}$ | 0.356 | 0.107 |
| $\Delta x_{t-3}$ | -0.100 | 0.109 |
| $(r_t - \overline{r}_t)$ | -0.054 | 0.035 |
| $(r_{t-1} - \overline{r}_{t-1})$ | 0.029 | 0.036 |
| $(r_{t-2} - \overline{r}_{t-2})$ | -0.019 | 0.034 |
| $(r_{t-3} - \overline{r}_{t-3})$ | -0.002 | 0.034 |
| $(r_{t-4} - \overline{r}_{t-4})$ | -0.016 | 0.033 |
| $x_{t-1}$ | -0.068 | 0.037 |
| Diagnostic Tests | | |
| AR 1-5 test: F(5,75) | 0.358 [0.875] | |
| ARCH 1-4 test: F(4,80) | 1.374 [0.250] | |
| Normality test: Chi^2(2) | 0.364 [0.833] | |
| Hetero test: F(16,71) | 1.244 [0.258] | |
| Hetero-X test: F(44,43) | 0.929 [0.595] | |
| RESET23 test: F(2,78) | 1.950 [0.149] | |

The values in brackets denote probability values (p-values).

Table 3.3: Unrestricted ADL IS Curve (U.S.)

Next, we reduce the unrestricted ADL IS curve by sequential elimination of regressors based on t-ratios of the parameters estimators. We sequentially delete those regressors with the smallest t-ratios and re-estimate the model until the remaining variables are significant at least at the 10% level.[17] The resulting restricted ADL model is displayed on table 3.4.

loss of overidentifying restrictions (i.e., the number of suspect instruments being tested), has the null hypothesis that the specified variables are proper instruments, that is, they are exogenous.

[16] Given the structure of the error term when the ADL IS curve is restricted to be either the NKISC or the HISC, another souce of inconsistency of the OLS estimator would be the one-period lagged regressors. We performed instrument orthoganality tests (C tests) for all one-period lagged regressors and for all of them it was not possible to reject the hypothesis that they are exogenous.

[17] In a context of VAR models, Brüggemann and Lütkepohl (2001) showed that this strategy is equivalent to the sequential elimination based on model selection criteria if the threshold value $\gamma$ is chosen accordingly.

| Dependent Variable: $\Delta x_t$ | | |
|---|---|---|
| Estimation Method: OLS | | |
| sample: 1985Q1–2007Q4 | | |
| Variable | Coefficient | Standard Error |
| $\Delta x_{t-2}$ | 0.359 | 0.097 |
| $x_{t-1}$ | -0.068 | 0.033 |
| $(r_t - \overline{r}_t)$ | -0.055 | 0.030 |
| Diagnostic Tests | | |
| AR 1-5 test: F(5,81) | 0.730 [0.602] | |
| ARCH 1-4 test: F(4,81) | 1.348 [0.259] | |
| Normality test: Chi^2(2) | 1.615 [0.445] | |
| Hetero test: F(6,82) | 0.405 [0.873] | |
| Hetero-X test: F(9,79) | 0.858 [0.565] | |
| RESET23 test: F(2,84) | 1.889 [0.157] | |

The values in brackets denote probability values (p-values).

Table 3.4: Restricted ADL IS Curve (U.S.)

Table 3.4 shows that the selected model is composed by the second lag of the change in output gap ($\Delta x_{t-2}$), the first lag of the output gap ($x_{t-1}$), and the current value of the real interest rate gap ($r_t - \overline{r}_t$). The diagnostic tests do not indicate any specification problems. Using equation (3.13) and comparing the values of Tables 3.1 and 3.4, we observe that none of the restrictions implied by the HISC is verified. Only one of the restrictions implied by the NKISC is accepted ($\alpha_1 = 0$), while all its other restrictions are rejected by the U.S. data. The KISC has three restrictions correct ($\beta_0 < 0$, $\beta_1 = 0$, $\gamma_1 < 0$), with the estimated sign of $\Delta x_{t-2}$ ($> 0$) being consistent with the KISC when output gap depends on additional output gap lags.

We conclude that only the KISC is consistent with the evidence for the U.S. from 1985Q1–2007Q4.

*Robustness*

In order to test the robustness of the results we use Autometrics, an automated model-selection algorithm, to reduce the ADL IS curve model, instead of the sequential elimination of the regressors based on the t-statistic.[18] The restricted ADL model selected by Autometrics coincides with the restricted model estimated by the sequential elimination of the regressors (table 3.4). Therefore, the same conclusions follow.

As another test of robustness, we estimate the ADL IS curve allowing for the explicit presence of the public sector in the model. We add lags 0–4 of the government expenditure gap and the public deficit gap (shown in figure 3.2) to the baseline model [equation (3.14)]. Table 3.6 displays the results of the estimated restricted ADL IS curve. The second lag of the government expenditure gap $\left(g_{t-2} - \overline{g}_{t-2}\right)$ and the third lag of the public deficit gap $\left(def_{t-3} - \overline{def}_{t-3}\right)$ are significative, in addition to variables that already appear in the baseline model [$\Delta x_{t-2}$ and $(r_t - \overline{r}_t)$]. However, the lag of the output gap ($x_{t-1}$) is not anymore significative, blurring the advantage of the KISC vis-à-vis the other models. Table 3.7(3.8) displays the estimated restricted ADL IS curve when only government

[18] Autometrics is a computer implementation of the general-to-specific modelling. It is part of the econometric software PcGive, version 13. For full details of Autometrics's algorithm, see Doornik and Hendry (2007) and Doornik (2009).

expenditure gap (public deficit gap) is added to the baseline model. In both cases the lagged output gap is now significative and its sign is in accordance with the KISC, but the interest rate gap is not significative. There is no clear evidence in favour of any of the models analysed.

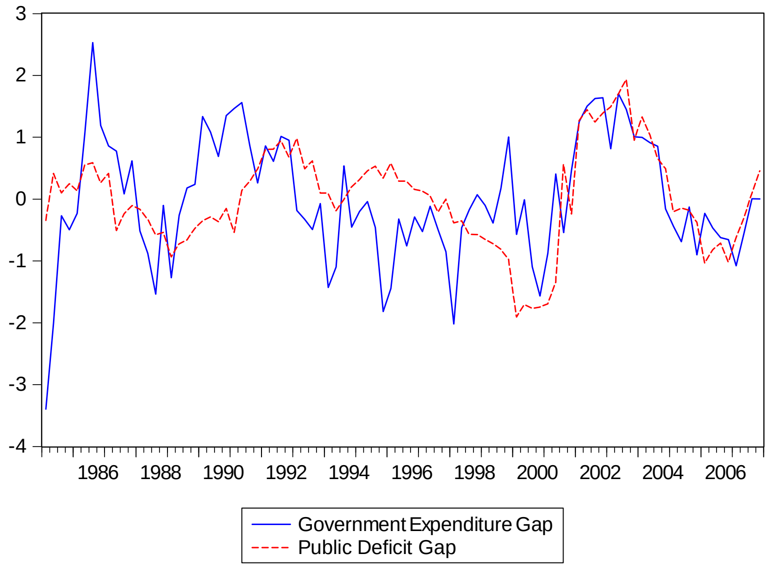



Figure 3.2: Public Deficit Gap and Government Expenditure Gap (U.S.)

| Null Hypothesis: Real Interest Rate Gap, Government Expenditure Gap, and Public Deficit Gap are exogenous | | | |
|---|---|---|---|
| | value | d.f. | Probability |
| Difference in J-statistics | 1.269 | 3 | 0.736 |

Table 3.5: Endogeneity Test for Real Interest Rate Gap, Government Expenditure Gap, Public Deficit Gap (U.S.)

| Dependent Variable: $\Delta x_t$ | | |
|---|---|---|
| Estimation Method: OLS | | |
| sample: 1985Q1–2007Q4 | | |
| Variable | Coefficient | Standard Error |
| $\Delta x_{t-2}$ | 0.278 | 0.094 |
| $\Delta x_{t-3}$ | -0.220 | 0.098 |
| $(r_t - \overline{r}_t)$ | -0.067 | 0.031 |
| $\left(g_{t-2} - \overline{g}_{t-2}\right)$ | -0.197 | 0.057 |
| $\left(def_{t-3} - \overline{def}_{t-3}\right)$ | 0.222 | 0.067 |
| Diagnostic Tests | | |
| AR 1-5 test: F(5,81) | 0.108 [0.990] | |
| ARCH 1-4 test: F(4,81) | 0.771 [0.547] | |
| Normality test: Chi^2(2) | 2.658 [0.264] | |
| Hetero test: F(6,82) | 1.673 [0.102] | |
| Hetero-X test: F(9,79) | 2.675 [0.001]** | |
| RESET23 test: F(2,84) | 1.363 [0.261] | |

The values in brackets denote probability values (p-values).
Significant diagnostic tests at 1% level are shown by two stars.

Table 3.6: Restricted ADL IS Curve when government expenditure gap and public deficit gap are added to the baseline model (U.S.)

| Dependent Variable: $\Delta x_t$ | | |
|---|---|---|
| Estimation Method: OLS | | |
| sample: 1985Q1–2007Q4 | | |
| Variable | Coefficient | Standard Error |
| $\Delta x_{t-1}$ | 0.250 | 0.101 |
| $\Delta x_{t-2}$ | 0.295 | 0.098 |
| $x_{t-1}$ | -0.092 | 0.032 |
| $(g_t - \overline{g}_t)$ | 0.182 | 0.073 |
| $\left(g_{t-1} - \overline{g}_{t-1}\right)$ | -0.203 | 0.066 |
| Diagnostic Tests | | |
| AR 1-5 test: F(5,79) | 0.326 [0.895] | |
| ARCH 1-4 test: F(4,81) | 0.207 [0.933] | |
| Normality test: Chi^2(2) | 3.440 [0.179] | |
| Hetero test: F(10,78) | 1.058 [0.404] | |
| Hetero-X test: F(20,68) | 0.842 [0.655] | |
| RESET23 test: F(2,82) | 2.116 [0.127] | |

The values in brackets denote probability values (p-values).
Significant diagnostic tests at 1% level are shown by two stars.

Table 3.7: Restricted ADL IS Curve when government expenditure gap is added to the baseline model (U.S.)

| Dependent Variable: $\Delta x_t$<br>Estimation Method: OLS<br>sample: 1985Q1–2007Q4 | | |
|---|---|---|
| Variable | Coefficient | Standard Error |
| $\Delta x_{t-2}$ | 0.366 | 0.096 |
| $x_{t-1}$ | -0.163 | 0.042 |
| $(def_t - \overline{def}_t)$ | -0.076 | 0.027 |
| Diagnostic Tests | | |
| AR 1-5 test: F(5,81) | 0.998 [0.423] | |
| ARCH 1-4 test: F(4,81) | 2.178 [0.078] | |
| Normality test: Chi^2(2) | 4.514 [0.104] | |
| Hetero test: F(6,82) | 0.846 [0.537] | |
| Hetero-X test: F(9,79) | 0.600 [0.793] | |
| RESET23 test: F(2,84) | 1.242 [0.294] | |

The values in brackets denote probability values (p-values).
Significant diagnostic tests at 1% level are shown by two stars.

Table 3.8: Restricted ADL IS Curve when public deficit gap is added to the baseline model (U.S.)

## 3.5 Empirical Evidence: Brazil (1996Q1 - 2012Q2)

We extend the unrestricted ADL IS curve [equation (3.14)] to introduce the exchange rate, an important open-economy modeling feature.

Including the exchange rate in the study of aggregate demand is important because the exchange rate allows additional channels for the transmission of monetary policy. In an open economy, the real exchange rate will affect the relative price between domestic and foreign goods, which, in turn, will affect both domestic and foreign demand for domestic goods.

In order to capture the importance of the exchange rate for the output dynamics we supplement the unrestricted ADL IS curve [equation (3.14)] with a 'level' term in the real exchange rate gap $(q_t - \overline{q}_t)$:

$$\Delta x_t = \sum_{s=1}^{n} \alpha_s \Delta x_{t-s} + \sum_{s=0}^{m} \beta_s \left(r_{t-s} - \overline{r}_{t-s}\right) + \gamma_1 x_{t-1} + \sum_{s=0}^{p} \delta_s \left(q_{t-s} - \overline{q}_{t-s}\right) + \varepsilon_t. \quad (3.15)$$

For Brazil the sample goes from 1996Q1 to 2012Q2. Figures 3.3–3.5 plot the output gap, the real interest rate gap, and the real exchange rate gap.[19]

Our empirical analysis starts with the estimation of the unrestricted open-economy ADL IS curve [equation (3.15)]. The endogeneity test displayed on table 3.9, indicates that both $(r_t - \overline{r}_t)$ and $(q_t - \overline{q}_t)$ can be treated as exogenous variables.[20] Then, we estimate the unrestricted ADL IS curve using OLS, setting $m = n = p = 4$ [see table 3.10].

[19] The Data Appendix gives details on the definitions of the variables used in the baseline model and the robustness analysis.

[20] Additional exogeneity tests (not displayed) suggest also that one-period lagged regressors are exogenous.

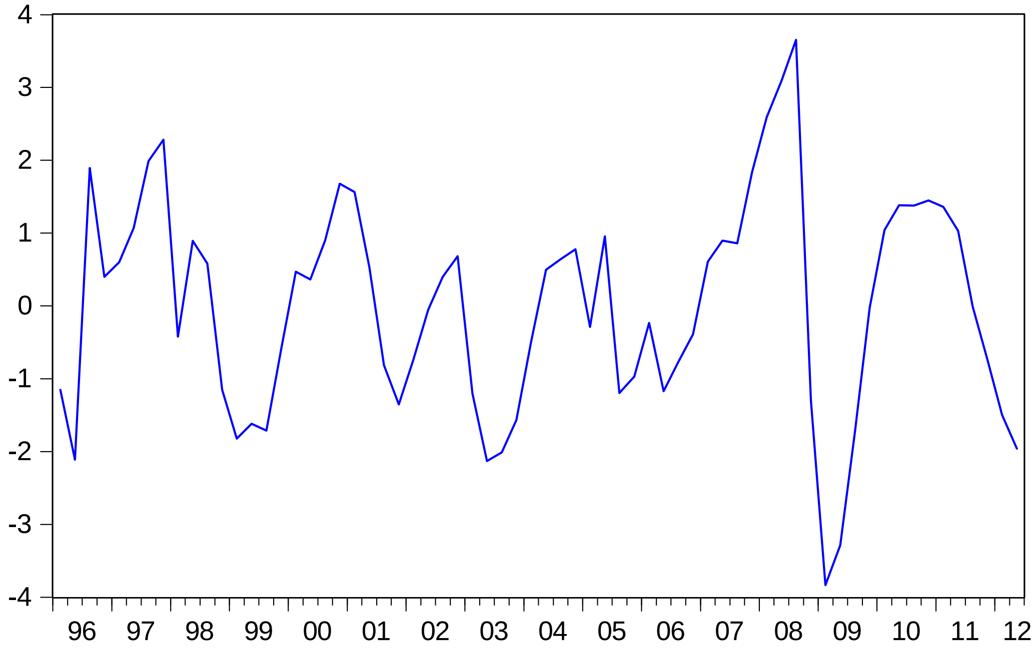



Figure 3.3: Ouput Gap (Brazil)

| Null Hypothesis: Real Interest Rate Gap and Real Exchange Rate Gap are exogenous | | | |
|---|---|---|---|
| | value | d.f. | Probability |
| Difference in J-statistics | 0.529 | 2 | 0.767 |

Table 3.9: Endogeneity Test for Real Interest Rate Gap and Real Exchange Rate Gap (Brazil)

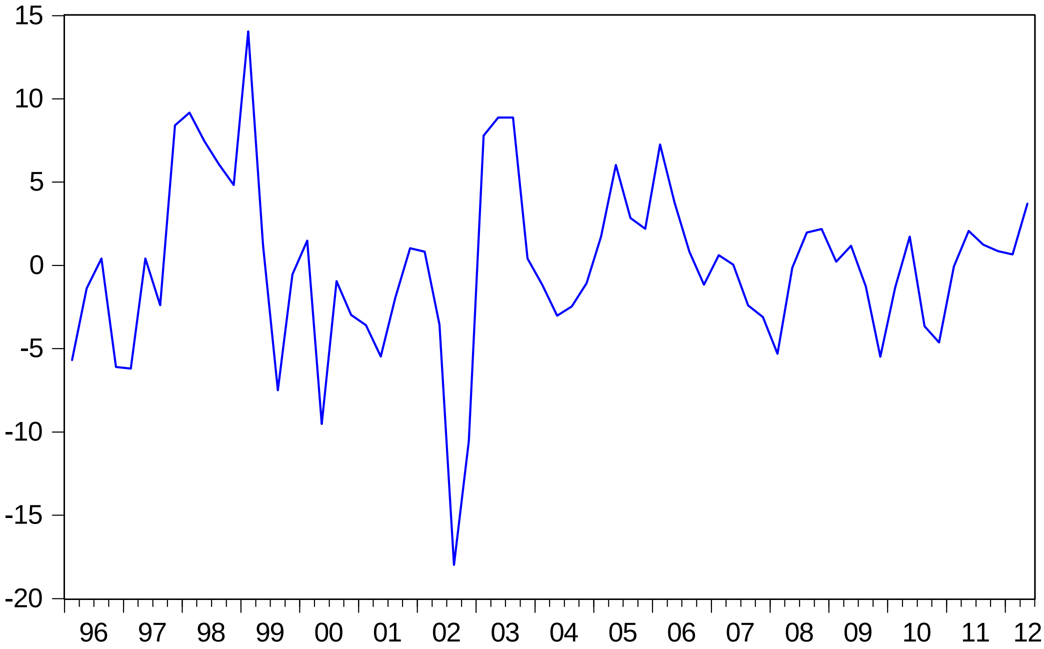



Figure 3.4: Real Interest Rate Gap (Brazil)

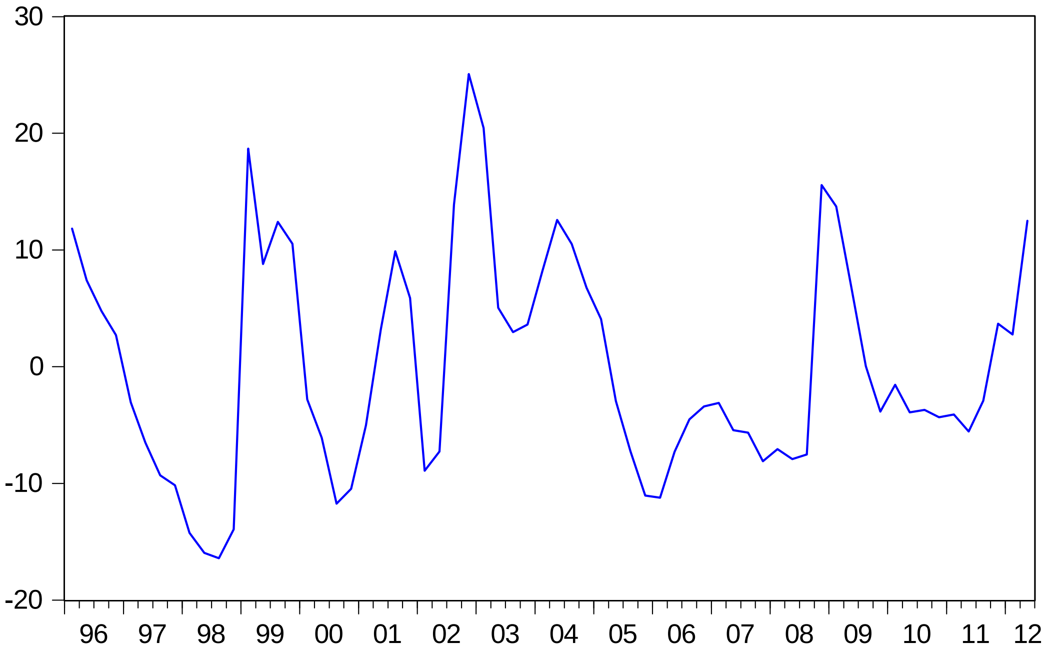



Figure 3.5: Real Exchange Rate Gap (Brazil)

| Dependent Variable: $\Delta x_t$ | | |
|---|---|---|
| Estimation Method: OLS | | |
| sample: 1996Q1–2012Q2 | | |
| Variable | Coefficient | Standard Error |
| $\Delta x_{t-1}$ | 0.255 | 0.147 |
| $\Delta x_{t-2}$ | 0.040 | 0.121 |
| $\Delta x_{t-3}$ | 0.191 | 0.103 |
| $(r_t - \overline{r}_t)$ | -0.033 | 0.025 |
| $(r_{t-1} - \overline{r}_{t-1})$ | -0.084 | 0.029 |
| $(r_{t-2} - \overline{r}_{t-2})$ | 0.017 | 0.032 |
| $(r_{t-3} - \overline{r}_{t-3})$ | -0.030 | 0.031 |
| $(r_{t-4} - \overline{r}_{t-4})$ | -0.048 | 0.028 |
| $x_{t-1}$ | -0.663 | 0.164 |
| $(q_t - \overline{q}_t)$ | -0.068 | 0.015 |
| $(q_{t-1} - \overline{q}_{t-1})$ | 0.010 | 0.023 |
| $(q_{t-2} - \overline{q}_{t-2})$ | -0.028 | 0.023 |
| $(q_{t-3} - \overline{q}_{t-3})$ | 0.025 | 0.023 |
| $(q_{t-4} - \overline{q}_{t-4})$ | 0.000 | 0.016 |
| Diagnostic Tests | | |
| AR 1-5 test: F(4,44) | 0.579 [0.678] | |
| ARCH 1-4 test: F(4,54) | 3.704 [0.009]** | |
| Normality test: Chi^2(2) | 3.059 [0.216] | |
| Hetero test: F(28,33) | 1.607 [0.095] | |
| RESET23 test: F(2,46) | 3.358 [0.043]* | |

The values in brackets denote probability values (p-values).
Significant diagnostic tests at 5% level are shown by one star.
Significant diagnostic tests at 1% level are shown by two stars.

Table 3.10: Unrestricted ADL IS Curve (Brazil)

Next, we reduce the unrestricted ADL IS curve by sequential elimination of regressors based on t-ratios of the parameters estimators. We sequentially delete those regressors with the smallest t-ratios and re-estimate the model until the t-ratios are all greater than the threshold $\gamma = 2$. The estimated restricted ADL model is displayed on table 3.11.

The selected model is composed by the first lags of the change in output gap ($\Delta x_{t-1}$), of the output gap ($x_{t-1}$), the real interest rate gap ($r_{t-1} - \overline{r}_{t-1}$), and the current value of the real exchange rate gap ($q_t - \overline{q}_t$). The diagnostic tests do not indicate any specification problems. Using equation (3.13) and comparing the values of Tables 3.1 and 3.11, we observe that none of the restrictions implied by the NKISC is accepted by the Brazilian data. The HISC gets only one restriction correct ($\alpha_1 > 0$). The KISC has two restrictions correct ($\alpha_1 > 0$, $\gamma_1 < 0$). Furthermore, the restrictions on the real interest rate gap implied by Brazilian data ($\beta_0 = 0$ and $\beta_1 < 0$) are consistent with a more backward-looking specification of the KISC, where the output gap depends on the lagged real interest rate gap, instead of the current rate.[21]

We conclude that only the KISC is consistent the evidence for Brazil from 1996Q1–2012Q2.

As the Brazilian economic policy suffered major changes in January 1999, when the

[21] Examples of such specification are Ball (1999) and Rudebusch and Svensson (1999).

exchange rate was allowed to float and in June 1999, when inflation targeting was implemented, we re-estimated the ADL model for the 1999Q3–2012Q2 subsample. Not only the restricted model is equal to the 1996Q1–2012Q2 sample, but the parameters estimates are quite stable. None of the conclusions changes.

| Dependent Variable: $\Delta x_t$<br>Estimation Method: OLS<br>sample: 1996Q1–2012Q2 | | |
|---|---|---|
| Variable | Coefficient | Standard Error |
| $\Delta x_{t-1}$ | 0.219 | 0.098 |
| $x_{t-1}$ | -0.658 | 0.088 |
| $(r_{t-1} - \overline{r}_{t-1})$ | -0.113 | 0.023 |
| $(q_t - \overline{q}_t)$ | -0.068 | 0.013 |
| Diagnostic Tests | | |
| AR 1-5 test: F(4,56) | 0.129 [0.971] | |
| ARCH 1-4 test: F(4,56) | 1.333 [0.269] | |
| Normality test: Chi^2(2) | 5.903 [0.052] | |
| Hetero test: F(8,55) | 0.789 [0.613] | |
| Hetero-X test: F(14,49) | 0.671 [0.790] | |
| RESET23 test: F(2,58) | 1.498 [0.232] | |

The values in brackets denote probability values (p-values).

Table 3.11: Restricted ADL IS Curve (Brazil)

*Robustness*

In order to test the robustness of the results, we use Autometrics to reduce the ADL IS curve model, instead of the sequential elimination of the regressors based on the t-statistic. The results (see table 3.12) bring the fourth lag of the real interest and real exchange rate as additional variables, and omits the first lag of the change in output gap, when compared to the results of the baseline model. However, the conclusion from the baseline model does not change: only the KISC is consistent with data.

| Dependent Variable: $\Delta x_t$ | | |
|---|---|---|
| Estimation Method: OLS | | |
| sample: 1996Q1–2012Q2 | | |
| Variable | Coefficient | Standard Error |
| $x_{t-1}$ | -0.484 | 0.082 |
| $(r_{t-1} - \overline{r}_{t-1})$ | -0.101 | 0.022 |
| $(r_{t-4} - \overline{r}_{t-4})$ | 0.053 | 0.022 |
| $(q_t - \overline{q}_t)$ | -0.067 | 0.012 |
| $(q_{t-4} - \overline{q}_{t-4})$ | 0.025 | 0.011 |
| Diagnostic Tests | | |
| AR 1-5 test: F(4,52) | 0.627 [0.645] | |
| ARCH 1-4 test: F(4,54) | 3.277 [0.017]* | |
| Normality test: Chi^2(2) | 1.882 [0.390] | |
| Hetero test: F(12,49) | 2.562 [0.010]* | |
| Hetero-X test: F(27,34) | 2.758 [0.002]** | |
| RESET23 test: F(2,54) | 2.824 [0.068] | |

The values in brackets denote probability values (p-values).
Significant diagnostic tests at 5% level are shown by one star.
Significant diagnostic tests at 1% level are shown by two stars.

Table 3.12: Restricted ADL IS Curve when Autometrics is employed (Brazil)

As another test of robustness we estimate the ADL IS curve allowing for the explicit presence of the public sector in the model. We add lags 0–4 of the government expenditure gap and the public deficit gap (shown in figure 3.6) to the baseline model [equation (3.15)]. The results are displayed on table 3.14. The restricted model brings the current government expenditure gap $[(g_t - \overline{g}_t)]$ and the fourth lag of the public deficit gap $[(def_{t-4} - \overline{def}_{t-4})]$, in addition to variables that already appear in the estimated baseline model: $x_{t-1}$, $(r_{t-1} - \overline{r}_{t-1})$, $(q_t - \overline{q}_t)$. The same conclusion holds: only the KISC is consistent with data.

| Null Hypothesis: Real Interest Rate Gap, Real Exchange Rate, Government Expenditure Gap, and Public Deficit Gap are exogenous | | | |
|---|---|---|---|
| | value | d.f. | Probability |
| Difference in J-statistics | 4.413 | 4 | 0.352 |

Table 3.13: Endogeneity Test for Real Interest Rate Gap, Real Exchange Rate, Government Expenditure Gap, Public Deficit Gap (Brazil)

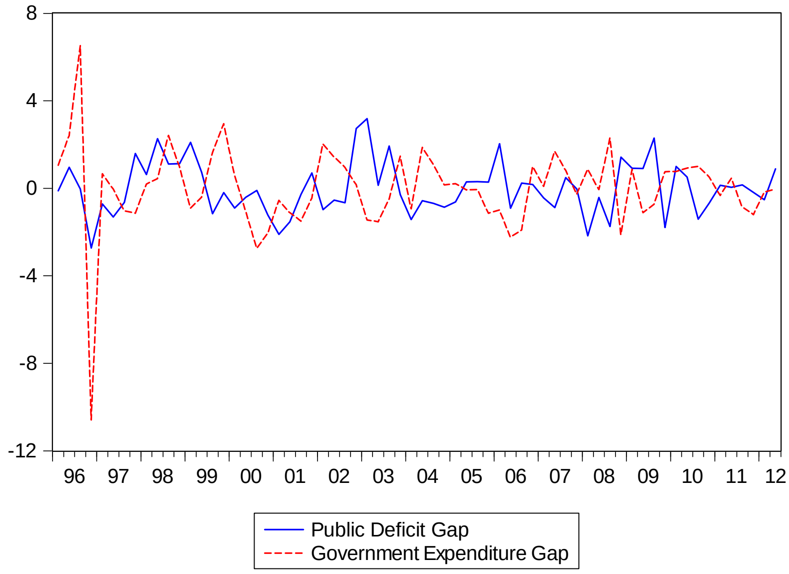



Figure 3.6: Public Deficit Gap and Government Expenditure Gap (Brazil)

| Dependent Variable: $\Delta x_t$ | | |
|---|---|---|
| Estimation Method: OLS | | |
| sample: 1996Q1–2012Q2 | | |
| Variable | Coefficient | Standard Error |
| $(r_{t-1} - \overline{r}_{t-1})$ | -0.118 | 0.020 |
| $x_{t-1}$ | -0.517 | 0.074 |
| $(q_t - \overline{q}_t)$ | -0.074 | 0.011 |
| $(g_t - \overline{g}_t)$ | 0.174 | 0.080 |
| $\left(def_{t-4} - \overline{def}_{t-4}\right)$ | 0.211 | 0.081 |
| Diagnostic Tests | | |
| AR 1-5 test: F(4,47) | 0.093549 [0.9840] | |
| ARCH 1-4 test: F(4,54) | 1.9623 [0.1134] | |
| Normality test: Chi^2(2) | 2.3373 [0.3108] | |
| Hetero test: F(22,39) | 1.4399 [0.1568] | |
| RESET23 test: F(2,49) | 2.3288 [0.1081] | |
| The values in brackets denote probability values (p-values). | | |

Table 3.14: Restricted ADL IS Curve when government expenditures gap and the public deficit gap are added to the baseline model (Brazil)

## 3.6 Concluding Remarks

In this paper we show that the NKISC and the HISC, despite being models based on microeconomic foundations with optimizing agents and rational expectations can be nevertheless considered implausible because their empirical implications fail to match those of the available data both for the U.S. and for Brazil. Conversely, the evidence seems to be consistent with the KISC which, in contrast with the NKISC and the HISC, is based on some *ad hoc* assumptions.

## 3.A Data Appendix

U.S. Data

The output gap ($x_t$) is given by 100 times the difference between the log of the quarterly real GDP and the log of the Congressional Budget Office's estimate of the potential real GDP.

The real interest rate gap ($r_t$) is calculated as the difference between the Federal Funds rate and next period actual inflation, detrended by the Hodrick-Prescott (HP) filter.

The public deficit gap ($def_t$) is calculated as (minus) the Net Government Saving over real GDP ratio, detrended by the HP filter.

The government expenditure gap ($g_t$) is given by the log of real government consumption expenditures & gross investment, detrended by the HP filter.

The capacity utilization gap is total industry capacity utilization, detrended by the HP filter.

Brazilian Data

The output gap ($x_t$) is given by 100 times the log of the quarterly real GDP seasonally adjusted, detrended by the HP filter.

The real interest rate gap ($r_t$) is calculated as the difference between the SELIC rate and next period actual inflation, detrended by the HP filter.

The real exchange rate gap ($q_t$) is calculated as 100 times the log of the quarterly average of monthly indexes of real effective exchange rates, detrended by the HP filter.

The public deficit gap ($def_t$) is calculated as the quarterly average of monthly values of the public sector borrowing requirements over GDP, seasonally adjusted and detrended by the HP filter.

The government expenditure gap ($g_t$) is given by the log of the government consumption index, detrended by the HP filter.

The capacity utilization gap is general installed capacity utilization, detrended by the HP filter.

# Chapter 4

# The Effects of Fiscal Policy and its Interactions with Monetary Policy in Brazil

## 4.1 Introduction

Recently, there has been a worldwide movement toward the adoption of a policy regime in which the central bank is assigned the task of achieving an inflation target. At the same time, the independence of central banks to pursue this goal has also increased, suggesting that the choice of monetary policy to achieve the inflation target is a problem that can, and in fact ought to be, separated from the choice of fiscal policy or any other public policy. As pointed out by Sims (1994), in a rational expectations, market-clearing equilibrium model with a costlessly-produced fiat money that is useful in transactions, the following is true:

1. the existence and uniqueness of the equilibrium price level cannot be determined from knowledge of monetary policy alone;
2. the determinacy of the price level under any policy depends on the public's beliefs about what the policy authority would do under conditions that are never observed in equilibrium.

Therefore, fiscal policy plays an important role, and the choice of monetary policy to achieve the inflation target should not be separated from the fiscal policy adopted by the government.

This paper aims to uncover some stylized facts related to the effects of fiscal policy shocks on the dynamics of the Brazilian economy and the interaction between fiscal and monetary policy in Brazil. To achieve our goal we use a structural vector autoregression (SVAR) model and the test proposed by Canzoneri, Cumby, and Diba (2001). The SVAR is identified by two alternative methodologies. The first methodology uses sign restrictions on impulse responses of the exogenous disturbances. The second methodology, developed by Lima, Maka, and Alves (2011) [LMA], combines sign restrictions with restrictions on the contemporaneous causal interrelationships among variables, derived by Directed Acyclic Graphs (DAGs). LMA analysis is concerned mainly with the identification of the effects of monetary policy and exchange rate shocks, so no attention was given to fiscal policy. In this paper, we extend LMA analysis introducing a set of fiscal variables (budget surplus and public sector liabilities) in their VAR model. The hybrid identification strategy pursued in this article consists of two steps. In the first step, we use DAGs to select over-identifying restrictions on the contemporaneous coefficients based on the conditional independence relations between the variables. These over-identifying restrictions allow us to identify monetary policy and demand shocks, and to restrict the covariance matrix of the reduced-form residuals. In the second step, maintaining restricted the covariance matrix of reduced-form residuals, we keep the identified monetary policy and demand shocks, and impose sign restrictions on the impulse response functions on other three shocks to identify the fiscal policy, supply, and exchange rate shocks. Analyzing the case of Brazil, we observe for both identification strategies that in response to positive ("contracionist") fiscal shocks there is a significative and long-lasting reduction in the price level, and a short-lived reduction on economic activity. There is no evidence of significative response of the exchange rate to fiscal innovations.

Dungey and Fry (2009) [DF] propose a different hybrid identification approach that combines traditional short-run restrictions, sign restrictions and long run restrictions. The hybrid methodology adopted here has some similarities with the one used by DF. However, we do not use long run restrictions and instead of traditional short-run restrictions, we use DAGs to impose restrictions on the contemporaneous causal interrelationships among variables. As for the sign restrictions, we use the QR decomposition to generate the candidate shocks, while DF use the Givens rotation.[1]

The motivation for our hybrid strategy comes from the fact that the DAG and sign restrictions approaches complement each other, so that their combination may be superior than each methodology taken isolated. While the DAG approach imposes restrictions that may identify exogenous shocks, the response of variables to these shocks may indicate that they are not the ones we are trying to identify. They may be linear combinations of the shocks we are interested on or parameter uncertainty may be responsible for the distortions in the responses. On the other side, Sign restrictions have economic justification but may not impose enough restrictions to identify the shocks (as described in the previous paragraph). We believe that a combination of the available methodologies increases the chance that all shocks of interest are identified.

According to the traditional monetarist view, a necessary and sufficient condition for achieving price stability is a fully credible commitment of the central bank to stable prices. This traditional analysis has been challenged by the Fiscal Theory of the Price Level (FTPL), which links price determination to the government present value budget constraint, i.e. the equality of the public debt with the present discounted value of future expected primary surpluses.[2] The key intuition of the FTPL is that, if current and future fiscal policies are set without concern about sustainability, the general price level will "jump" in order to fulfill the present value budget constraint. This idea contrasts with the conventional monetarist theory of price determination, according to which the stock of money (and thus the central bank) is the sole determinant of the price level and fiscal policy is (often implicitly) assumed to passively adjust primary surpluses to guarantee solvency of the government for any price level.[3] Such fiscal policy is called Ricardian.

[1] Fry and Pagan (2007) show that the QR decomposition and the Givens rotation are equivalent. However, as the model grows in size the QR decomposition is expected to be superior in terms of computational speed.

[2] For an introduction to the FTPL see Carlstrom and Fuerst (2000), Christiano and Fitzgerald (2000), and Canzoneri, Cumby, and Diba (2001).

[3] This view can be summarized in Milton Friedman's dictum that "inflation is always and everywhere a monetary phenomenon".

The FTPL reverse the argument above: if the fiscal authority chooses primary surpluses independently of government debt, then it is the price level that has to adjust to satisfy the present value government budget constraint. This alternative regime is called non-Ricardian.[4]

Two main features distinguish our work from the related empirical literature that tested the FTPL for Brazil. First, in contrast to the existing studies that applied Canzoneri, Cumby, and Diba (2001) [CCD] test for Brazil,[5] which restrict their analysis to a 2-3 variable closed economy VAR model usually containing only the primary surplus and government liabilities, our investigation involves much more variables (9), including key variables like the exchange rate and interest rates, that allow us to better evaluate the impact of fiscal policy shocks and its interaction with other economic variables. Second, the identification strategies adopted in this article, depart from the Cholesky decomposition usually followed in the literature, and represent an effort to overcome the limitations of the available identification methodologies. We test the assumption, held by the Fiscal Theory of the Price Level, that the policy regime is non-Ricardian (Fiscal Dominance), applying the test proposed by CCD that analyzes the response of public sector liabilities to primary surplus shocks. This response depends on the identification adopted. For the hybrid identification we find that it is not possible to distinguish empirically between Ricardian (Monetary Dominance) and non-Ricardian (Fiscal Dominance) regimes. However, using sign restrictions there is some evidence that the government followed a Ricardian (Monetary Dominance) regime from January 2000 to June 2008.[6]

We also check if the identified exogenous monetary policy shocks show a "stepping on a rake" effect (tighter monetary policy leads to a higher inflation rate in the long run), as described by Sims (2008) in a theoretical framework designed for understanding the effects of fiscal uncertainties on monetary policy. According to our results, there is no evidence that a tighter monetary policy would lead to higher inflation in the long run.

The article is organized as follows. Section 2 presents an overview of the FTPL. Section 3 describes the empirical model and the results. Its first part presents the hybrid identification procedure that combines short-run restrictions on the contemporaneous coefficients with sign restrictions on the impulse response functions. Its second part shows an alternative identification procedure based on sign restrictions only. Finally, Section 4 offers some concluding remarks.

## 4.2 An Overview of the Fiscal Theory of the Price Level

The government budget constraint is an accounting identity linking monetary and fiscal policies at each point in time and across time.[7] The government budget constraint

[4]Some authors refer to the Ricardian regime as "Monetary Dominance" and to the non-Ricardian regime as "Fiscal Dominance".

[5]See, for example, Tanner and Ramos (2002), Rocha and Silva (2004), and Fialho and Portugal (2005).

[6]The results of the studies that applied the CCD test for Brazil are mixed. Tanner and Ramos (2002) found evidence of non-Ricardian regime for the 1991-2000 period using monthly data. Rocha and Silva (2004) and Fialho and Portugal (2005) instead, found evidence of Ricardian regime for the 1966-2000 (with annual data) and 1995-2003 (with monthly data) period, respectively.

[7]The presentation of the FTPL presented in this section follows CCD closely.

for period $j$ can be written in nominal terms as[8]

$$B_t = (T_t - G_t) + (M_{t+1} - M_t) + B_{t+1}/(1 + i_t) \tag{4.1}$$

where $M_t$ and $B_t$ are the stocks of base money and government debt at the beginning of period $t$, $T_t - G_t$ is the primary surplus during period $t$, and $i_t$ is the interest rate for period $t$.

Expressing the budget constraint in terms of total government liabilities, $M + B$, and scaling the fiscal variables on GDP, we have that

$$\frac{M_t + B_t}{P_t y_t} = \left[\frac{T_t - G_t}{P_t y_t} + \left(\frac{M_{t+1}}{P_t y_t}\right)\left(\frac{i_t}{1 + i_t}\right)\right] + \left(\frac{y_{t+1}/y_t}{(1 + i_t)(P_t/P_{t+1})}\right)\left(\frac{M_{t+1} + B_{t+1}}{P_{t+1} y_{t+1}}\right) \tag{4.2}$$

Equation (4.2) can be written synthetically as

$$w_t = s_t + \alpha_t w_{t+1} \tag{4.3}$$

where $w_t$ is the liabilities to GDP ratio, $s_t$ is the surplus (including seigniorage) to GDP ratio, and $\alpha_t$ is the discount factor represented by the ratio of the real growth in GDP to the real interest rate.

Iterating equation (4.3) forward from the current period, $j$, and taking expectations conditional on information available in period $j$, we obtain the present value budget constraint

$$w_j = s_j + E_t \sum_{t=j+1}^{+\infty} \left(\Pi_{k=j}^{t-1} \alpha_k\right) s_t \tag{4.4}$$

The difference between the conventional view and the FTPL lies on the way in which the government's present value budget constraint [equation (4.4)] is satisfied. The conventional view holds that this equation is a constraint on the government's tax and expenditure policies. According to this view, when equation (4.4) is disturbed, the government must alter its expenditures or its taxes to restore equality. FTPL advocates, however, argue that present value budget constraint is not a constraint on policy, but instead it is an equilibrium condition: when something threatens to disturb the equation, the market-clearing mechanism moves the price level, $P$, to restore equality.

The policy regime is said to be *Ricardian* (R) if the sequence $\{s_t\}$ is chosen so that the intertemporal budget equation (4.4) is satisfied no matter what $P$ is realized. In contrast, if $\{s_t\}$ is chosen in a way that does not guarantee that equation (4.4) is satisfied for all possible prices, the policy regime is said to be *non-Ricardian* (NR). The assumption that the policy regime is non-Ricardian is what distinguishes the FTPL from the conventional view.

The Ricardian and non-Ricardian regimes are observationaly equivalent as they use the same equations to explain a given data set. It is not possible to test whether the government has chosen to follow a Ricardian or a non-Ricardian policy regime because the FTPL per se has no testable implications. The budget constraint (4.4) holds in equilibrium for both regimes. The issue is whether, in determining or adjusting towards equilibrium, the price level adjusts to expected future surpluses, or whether the path of

[8] We are assuming the government issues nominal liabilities ($M$ and $B$); while the nominal values of these liabilities are fixed at the beginning of the period, their real values depend on the price level.

surpluses adjusts in response to the price level. All we observe is an equilibrium; we do not observe who adjusted to bring about that equilibrium. However, one way of assessing the empirical value of the FTPL is viewing the non-Ricardian assumption as a starting point for a set of testable auxiliary assumptions that restrict the time series data, and then test those restrictions.

CCD proposed to differentiate between R and NR regimes studying the response of public liabilities to positive surplus shocks in a bivariate VAR. In an R regime, the surplus positive innovation pays off some of the debt, and $w_{t+1}$ falls. In a NR regime, given a positive $s_t$ innovation, there are three possibilities:

**(i)** corr $(s_t, s_{t+k}) = 0$ and corr $(s_t, \alpha_{t+k}) = 0; w_{t+1}$ constant

**(ii)** corr $(s_t, s_{t+k}) > 0$ and corr $(s_t, \alpha_{t+k}) > 0; w_{t+1}$ increases

**(iii)** corr $(s_t, s_{t+k}) < 0$ and corr $(s_t, \alpha_{t+k}) < 0; w_{t+1}$ decreases

In cases (i) and (ii) it should in principle be possible to differentiate between R and NR regimes. For example, the impulse response function from a VAR in $s_t$ and $w_t$ would tell us how $w_{t+1}$ responds to an innovation in $s_t$. If $w_{t+1}$ falls, we have an R regime; if it does not, we have an NR regime. However, in case (iii) $w_{t+1}$ would fall in either an R regime or an NR regime, and we have an identification problem.

## 4.3 Data, Model Specification and Estimation

The model is estimated using monthly data and it is composed of the following variables: the short-term interest rate (SELIC), the nominal exchange rate (EXCHRATE), the price index (IPCA), the medium-term interest rate (SWAP), output, a monetary aggregate (M1), public sector net liabilities over GDP (LIABILITIES), primary surplus over GDP (SURPLUS), a discount factor based on nominal GDP (TXDESCS), a constant, and seasonal dummies.[9] We use the primary surplus as a measure of fiscal stance to avoid the problem of separately identifying tax revenue and government expenditure exogenous innovations.

Our sample period starts on 2000:01 and goes until 2008:06. The lag length chosen is six months. The model identifies five independent sources of exogenous disturbances: fiscal policy, monetary policy, demand, supply, and exchange rate shocks.

We use the Gibbs Sampling algorithm developed by Waggoner and Zha (2002) to estimate the model. A detailed description of the application of the methodology in our case is described in the Technical Appendix.

## 4.4 Model Identification

Over the last years there has been a growing interest on graphical models and in particular on those based on DAGs as a general framework to describe and infer causal relations, exploring the connection between causal structure and probability distributions. These methods have been used in a variety of fields but are unfamiliar to most economists. Swanson and Granger (1997) were the first to apply graphical models to identify contemporaneous causal order of a SVAR, although they restrict the admissible structures to

[9] A detailed description of the data and its sources can be found in the Data Appendix.

causal chains. Bessler and Lee (2002) use error correction and DAGs to study both lagged and contemporaneous relations in late 19th and early 20th century U.S. data. Demiralp and Hoover (2003) evaluate the PC algorithm employed by TETRAD in a Monte Carlo study and conclude that it is an effective tool of selecting the contemporaneous causal order of SVARs. Awokuse and Bessler (2003) use DAGs to provide over-identifying restrictions on the innovations from a VAR and compare their results with the ones of Sims (1996). Monetta (2004) use DAGs and the data set of Bernanke and Mihov (1998) to identify the monetary policy shocks and their macroeconomic effects in the U.S..

### 4.4.1 The hybrid approach

The hybrid identification strategy pursued in this article consists of two steps. In the first step, we use DAGs to select over-identifying restrictions on the contemporaneous coefficients based on the conditional independence relations between the variables. These over-identifying restrictions allow us to identify monetary policy and demand shocks, and to restrict the covariance matrix of the reduced-form residuals. In the second step, maintaining restricted the covariance matrix of reduced-form residuals, we keep the identified monetary policy and demand shocks, and impose sign restrictions on the impulse response functions of the other three shocks to identify fiscal policy, supply, and exchange rate shocks.

Step 1: Selection of the Over-Identifying Restrictions to Identify Monetary Policy Shocks[10]

Spirtes, Glymour, and Scheines (2000) [SGS] developed algorithms for inferring causal relations from data that are embodied in a computer program used in this article, called TETRAD.[11] The program assumes a multivariate normal distribution and takes as input the covariance matrix of the variables of the model,[12] converting it into a correlation matrix and performing hypothesis tests in which the null hypothesis is a zero partial correlation.

Conditional independence is a key notion in multivariate analyses such as graphical modelling, where two vertices are connected if and only if the corresponding variables are not conditionally independent. To confirm the conditional independence, it is a common practice to check whether or not the partial correlation is close enough to zero. This is done because it is assumed that zero partial correlation suggests that the variables are conditionally independent, or nearly so. Under the assumption of multivariate normality, a test of zero correlation or zero partial correlation is also a test of independence or conditional independence. Moreover, if $X$, $Y$ and $Z$ are normally distributed, the partial correlation coefficient $\rho_{XYZ}$ is zero if and only if $X$ is independent of $Y$ conditional on $Z$.

TETRAD begins with a 'saturated' causal graph, where any pair of nodes (variables) is joined by an undirected edge.[13] If the null hypothesis of zero partial correlation cannot be rejected – at, say, the 5% level, using Fisher's $z$ test – the edge is deleted.[14] After

[10] For an introduction on how to use DAGs to identify VARs, see Lima, Maka, and Céspedes (2008).

[11] The program is available for download at www.phil.cmu.edu/projects/tetrad/index.html. We used TETRAD III in this paper.

[12] In our application, the input of TETRAD is the covariance matrix of the reduced form VAR residuals.

[13] An edge in a graph can be either directed (marked by a single arrowhead on the edge) or undirected (unmarked). Arrows represent causal relationships: if there is an arrow pointing from $X_i$ to $X_j$ it means that $X_i$ has a direct causal effect on $X_j$.

[14] In the case of the normal distribution, the partial correlation coincides with the conditional correlation, which is another measure of conditional independence of two random variables. See Baba, Shibata,

examining all pair of vertices, TETRAD move on to triples, and so forth, orienting the edges left in the graph through the connection between probabilistic independence and graph theory. The final output of TETRAD is a set of observationally equivalent DAGs containing the proposed causal structure(s) of the model.

Robins et al. (2003) showed that the asymptotically consistent procedures of SGS are pointwise consistent, but not uniform consistent.[15] Furthermore, they also showed that there exists no causality test, based on associations of non-experimental data under the conditions assumed by SGS, which is uniform consistent. Therefore, for any finite sample, it is impossible to guarantee that the results of the SGS causality tests (or any other causality test) will converge to the asymptotic results.

Under the SGS model, it is sufficient to have a sample covariance between two variables, say, $v_1$ and $v_2$, exactly equal to zero to deduce that $v_1$ is not a cause of $v_2$. However, if the sample correlation between $v_1$ and $v_2$ is not exactly zero (as will almost always happens in finite samples) and the true model is unknown, as Robins et al. (2003) have shown, the acceptance or rejection of the null hypothesis of zero partial correlation is not unequivocally tied to the absence of causality. In other words we don't know, in any finite sample, how close to zero a partial correlation has to be to indicate non-causality. When the sample correlation is not exactly zero, it is not possible to determine which significance level should be used to test for zero partial correlation when attempting to test for the presence of causality. The "significance level", used by Tetrad, cannot be interpreted as the probability of type I error for the pattern output, but merely as a parameter of search. The higher is this parameter of the search, the smaller is the absolute value of the partial correlation that is taken as an indication of absence of causality. Intuitively, we are assuming that small partial correlations indicate small direct causal effect but we don't know how small the absolute value of the correlation has to be to obtain the correct causal inferences for the sample data we are using. Nevertheless, we can test the sensibility of the impulse response function of the model to different discrete values of this "parameter".[16]

Applying the software TETRAD on a 20% "significance level"[17] (our search parameter) and imposing the restriction that the SWAP rate[18] affects contemporaneously the SELIC rate set by the Central Bank we obtain a graphical representation of the DAG containing the contemporaneous causal ordering of the variables, displayed on Figure 4.1.[19,20]

It is interesting to notice that the introduction of fiscal variables and discount factor

and Sibuya (2004) for further details.

[15] A pointwise consistent test is guaranteed to avoid incorrect decision if the sample size can be increased indefinitely. However, pointwise consistency is only a guarantee about what happens in the limit, not at any finite sample size. A stronger form of consistency, uniform consistency, guarantees that it is possible to bound the decisions error rates with a finite number of observations.

[16] This is a bit more data oriented than the usual procedure of changing the order of the Cholesky decomposition of reduced form VAR residuals to identify the model.

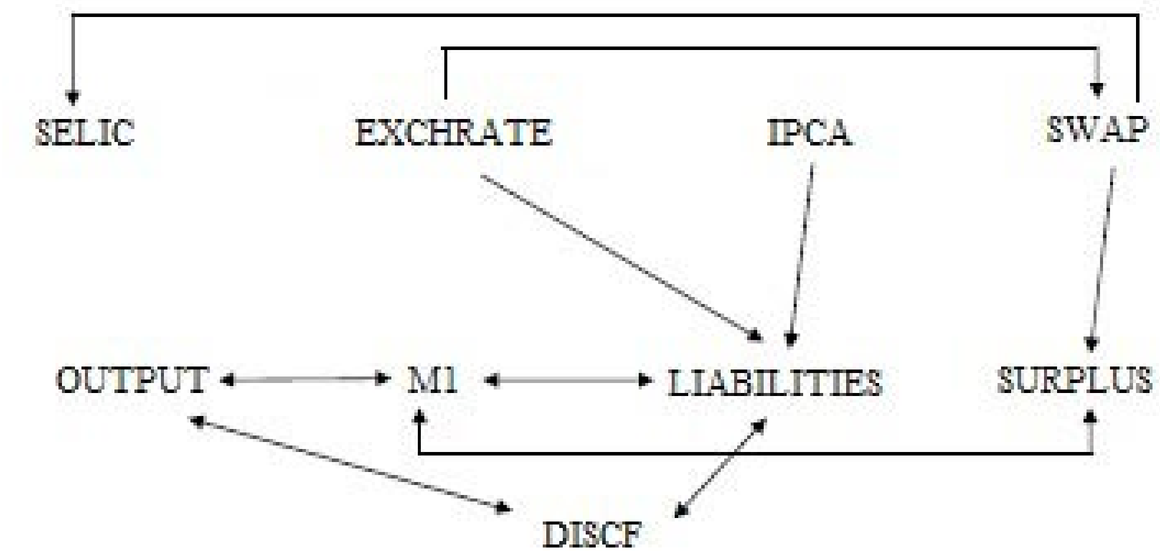
[17] We tested different discrete values for this parameter in the neighborhood of the chosen level (20%) and the model's impulse response function didn't change much.

[18] The 180 days SWAP rate is partially affected by the expectations of future SELIC rates.

[19] If we do not assume that the central bank takes into account the SWAP rate when setting the SELIC rate, we observe that the price level temporarily increases in response to a positive SELIC shock, a result known in the literature as the "price puzzle".

[20] In reality TETRAD puts an undirected edge between exchange rate and swap, meaning that there is causality in one of the two directions, but not on both. In what follows we restrict our attention on the causality going from exchange rate to the swap rate. However the results discussed next doesn't change much when the alternative causal ordering is used.

Figure 4.1: Contemporaneous causal ordering based on DAGs




changes completely the contemporaneous ordering obtained by LMA (see Figure 1 of their article). According to Figure 4.1, none of the policy variables affect contemporaneously the price level. The SELIC rate does not affect any variable contemporaneously, while the stock of money (M1) has an effect over the level of economic activity. LIABILITIES and SURPLUS have a contemporaneous effect only over M1 and the discount factor.

The causal ordering between the variables of the VAR can be represented by matrix A that establishes a relationship between reduced form and structural form residuals. The DAG pictured on Figure 1 can be represented by the following matrix:

$$A = \begin{bmatrix} A_{11} & 0 & 0 & A_{14} & 0 & 0 & 0 & 0 & 0 \\ 0 & A_{22} & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & A_{33} & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & A_{42} & 0 & A_{44} & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & A_{55} & A_{56} & 0 & 0 & A_{59} \\ 0 & 0 & 0 & 0 & A_{65} & A_{66} & A_{67} & A_{68} & 0 \\ 0 & A_{72} & A_{73} & 0 & 0 & A_{76} & A_{77} & 0 & A_{79} \\ 0 & 0 & 0 & A_{84} & 0 & A_{86} & 0 & A_{88} & 0 \\ 0 & 0 & 0 & 0 & A_{95} & 0 & A_{97} & 0 & A_{99} \end{bmatrix}$$

where $A_{ij}$ are parameters to be estimated and the vector of endogenous variables that multiplies A is given by [SELIC, exchange rate, IPCA, SWAP, output, M1, LIABILITIES, SURPLUS, TXDESCS].

The contemporaneous causal ordering resulting from the application of DAGs implies restrictions on the covariance matrix of the reduced-form residuals, meaning that we now have an overidentified model. Structural VAR models that are overidentified can be consistently estimated only by Bayesian estimation methods that introduce these restrictions on the covariance matrix of reduced form residuals. These restrictions are considered when Bayesian estimation methods are applied to the parameters of a structural VAR (and not to the parameters of a reduced form VAR). The method developed by Sims and Zha (1998), and adopted in this article, is one of these methods.

Using the contemporaneous causal ordering of Figure **??** to identify the SVAR, we obtained the impulse response functions of economic variables to exogenous and independent shocks, displayed on Figure 4.2. We identify SELIC shocks as monetary policy shocks and output shocks as demand shocks, leaving fiscal policy shocks to be identified by sign restrictions in the next step, when we identify also exchange rate and supply shocks in order to better identify fiscal policy disturbances.

According to Figure 4.2, after a positive monetary policy (SELIC) innovation that correspond to an increase in the SELIC rate, the stock of M1 falls and output decreases temporarily, taking near 12 months to recover. The direction of the exchange rate response is not clear, but it is more likely that it will depreciate slightly in the short-run. The price level goes down, but it takes near six months until the price level starts to fall despite the contraction of economic activity. The public sector net liabilities temporarily increase, probably as a result of a fall in the primary surplus and the larger interest payments. In response to a positive demand (output) innovation we observe an increase in prices and a possible exchange rate appreciation.

### Step 2: Imposing Sign Restrictions to Identify Fiscal Policy, Supply and Exchange Rate Shocks

Having identified monetary policy shocks and demand shocks, and restricted the covariance matrix of the reduced-form residuals using the contemporaneous causal order suggested by TETRAD for the consolidated public sector, now we impose sign restrictions on the remaining impulse response functions in order to identify fiscal policy, supply and exchange rate shocks. The fiscal policy sign restrictions are based on the model developed by Sims (2008), while the supply and exchange rate restrictions are based on LMA and can be justified by the short-run dynamics of a stochastic open-economy macroeconomic model. Table 4.1 summarizes the sign restrictions on the IRFs used to identify the fiscal policy, supply, and exchange rate shocks.[21] The sign restrictions are supposed to hold for two months.

According to Table 4.1, a positive ("contracionist") fiscal shock does not reduce the primary surplus, does not increase the SELIC rate, the price level, output, the SWAP rate. A positive supply shock implies that prices do not increase, while output and primary surplus do not go down. An unexpected depreciation of the nominal exchange rate is supposed to imply changes in the same direction of the real exchange rate, and that the short-term interest rate, prices, output, and the surplus do not go down after the exchange rate shock.

The IRFs that result from the imposition of sign restrictions are presented in figures 4.3-4.4, showing the median as well as the 68% probability bands for a horizon of 24 and 60 months following the shocks, respectively.

In response to positive ("contracionist") fiscal shocks we observe a significative and long-lasting reduction in the price level and a short-lived reduction on economic activity. There is no evidence of significative response of the exchange rate to fiscal innovations. The primary surplus increases but the direction of the response of public liabilities to

[21] The (log) real exchange rate is defined as $q_t = s_t + p_t^* - p_t$, where $s_t$ is the (log of ) nominal exchange rate, $p_t(p_t^*)$ is the (log of) domestic (foreign) price level. We assume that the foreign price level is constant, so that a restriction on the real exchange rate translates into a restriction on $s_t - p_t$.

Figure 4.2: IRFs with 68% probability bands, using the contemporaneous causal ordering of Figure 1 to identify the SVAR (24 months ahead)

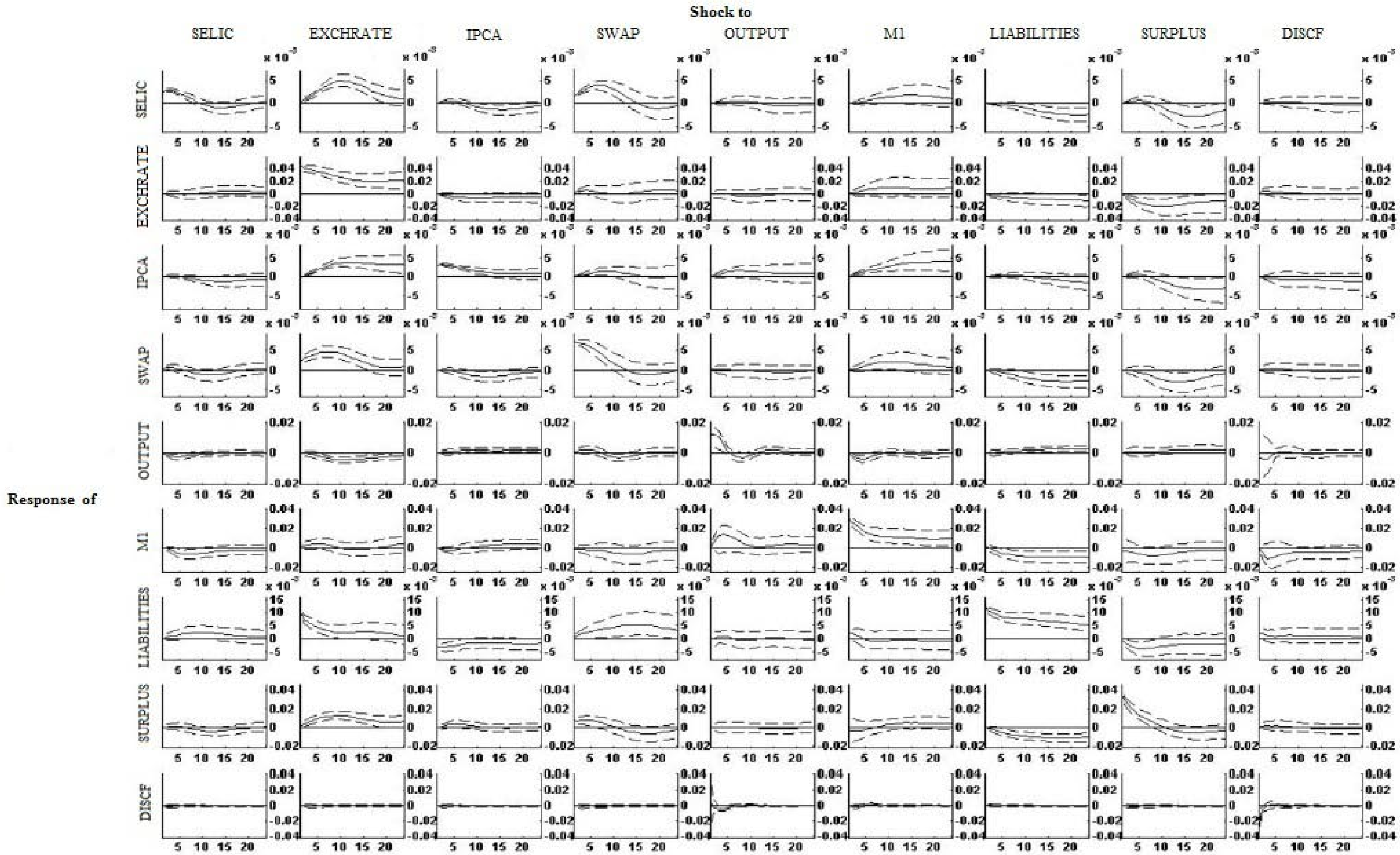


Table 4.1: Sign restrictions used to identify the SVAR model

| Response of / Type of shock | SELIC | IPCA | Output | SWAP | Real exchange rate | Surplus |
|---|---|---|---|---|---|---|
| Fiscal Policy | $\leq 0$ | $\leq 0$ | $\leq 0$ | $\leq 0$ | | $\geq 0$ |
| Supply | | $\leq 0$ | $\geq 0$ | | | $\geq 0$ |
| Exchange rate | $\geq 0$ | $\geq 0$ | $\geq 0$ | | $\geq 0$ | $\geq 0$ |

A blank entry indicates that no restrictions have been imposed.

positive fiscal policy innovations is not clear. Therefore, applying CCD's test to the results of the hybrid identification we are unable to distinguish empirically between Ricardian and non-Ricardian regimes. Monetary policy does not control the long run rate of inflation, as shown by response of prices to SELIC shocks 60 months ahead (Figure 4.4). However, there is no evidence of what Sims (2008) calls "step on a rake" effect, where increases in the interest rate increase, rather than decrease, the inflation rate.[22]

### 4.4.2 The sign restrictions approach

We consider now an alternative identification where we impose sign restrictions on the IRFs to all shocks, including monetary policy and demand shocks. We maintain the previous restrictions summarized in Table 1, and impose additional restrictions on monetary policy and demand shocks based on LMA and Sims (2008). We assume that in response to a "contractionary" monetary policy shock, interest rates does not fall, and that output, prices, M1, and the real exchange rate do not increase. We assume further that positive demand shocks do not decrease the SELIC rate, the price level, output, the primary surplus, and do not imply a depreciation of the real exchange rate. Table 4.2 shows the sign restrictions on the IRFs used to identify the fiscal policy, supply, exchange rate, monetary policy, and demand shocks. We impose the sign restrictions for a two months window.

Table 4.2: Sign restrictions used to identify the SVAR model

| Response of / Type of shock | SELIC | IPCA | Output | SWAP | Real exchange rate | Surplus | M1 |
|---|---|---|---|---|---|---|---|
| Fiscal Policy | $\leq 0$ | $\leq 0$ | $\leq 0$ | $\leq 0$ | | $\geq 0$ | |
| Supply | | $\leq 0$ | $\geq 0$ | | | $\geq 0$ | |
| Exchange Rate | $\geq 0$ | $\geq 0$ | $\geq 0$ | | $\geq 0$ | $\geq 0$ | |
| Monetary Policy | $\geq 0$ | $\leq 0$ | $\leq 0$ | | $\leq 0$ | | $\leq 0$ |
| Demand | $\geq 0$ | $\geq 0$ | $\geq 0$ | | $\leq 0$ | $\geq 0$ | |

A blank entry indicates that no restrictions have been imposed.

For the IRFs based on the alternative identification that uses only sign restrictions to

[22] Loyo (1999) refers to this situation as "tight money paradox".

Figure 4.3: IRFs based on the hybrid identification (24 months ahead), with 68% probability

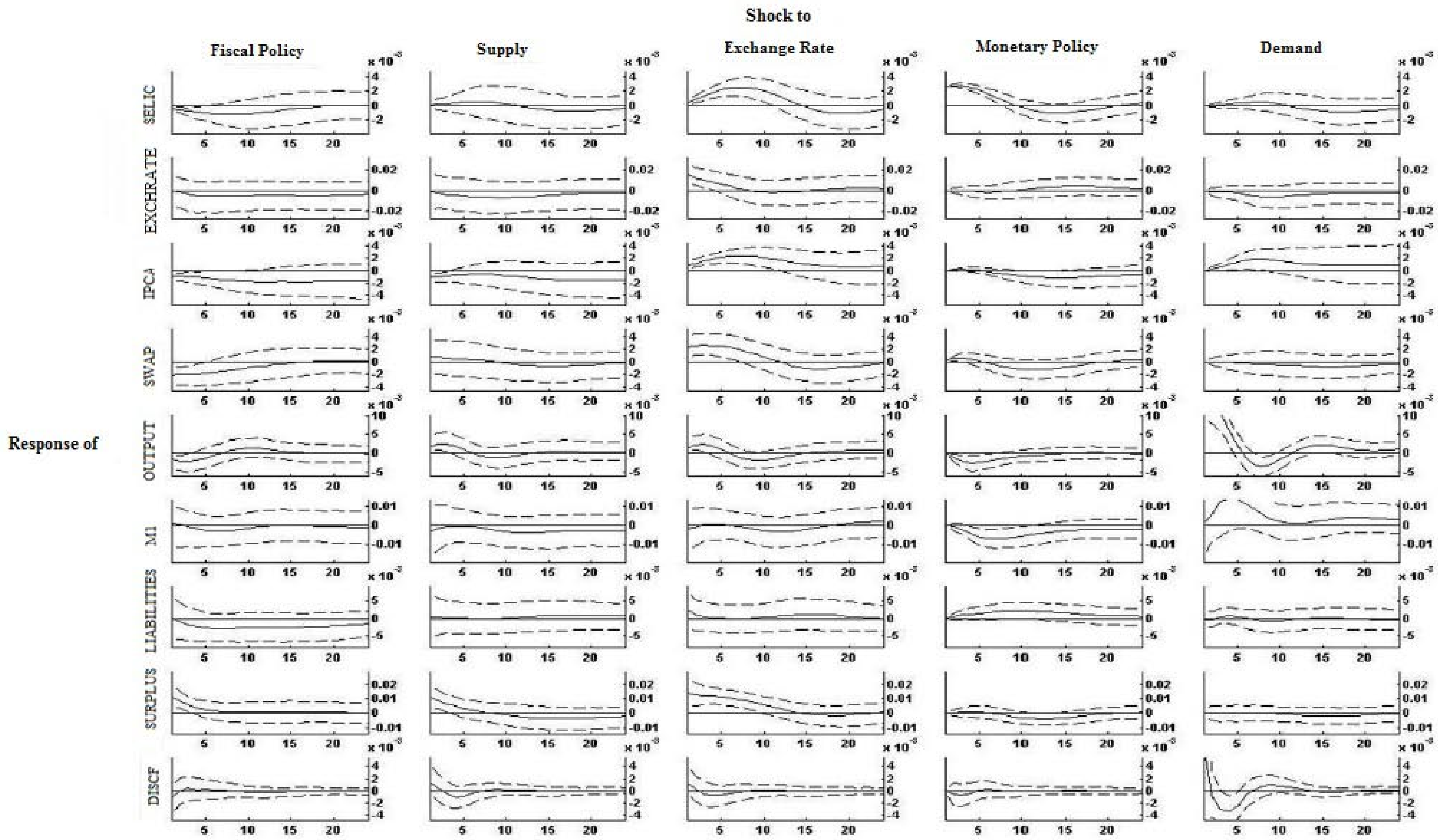
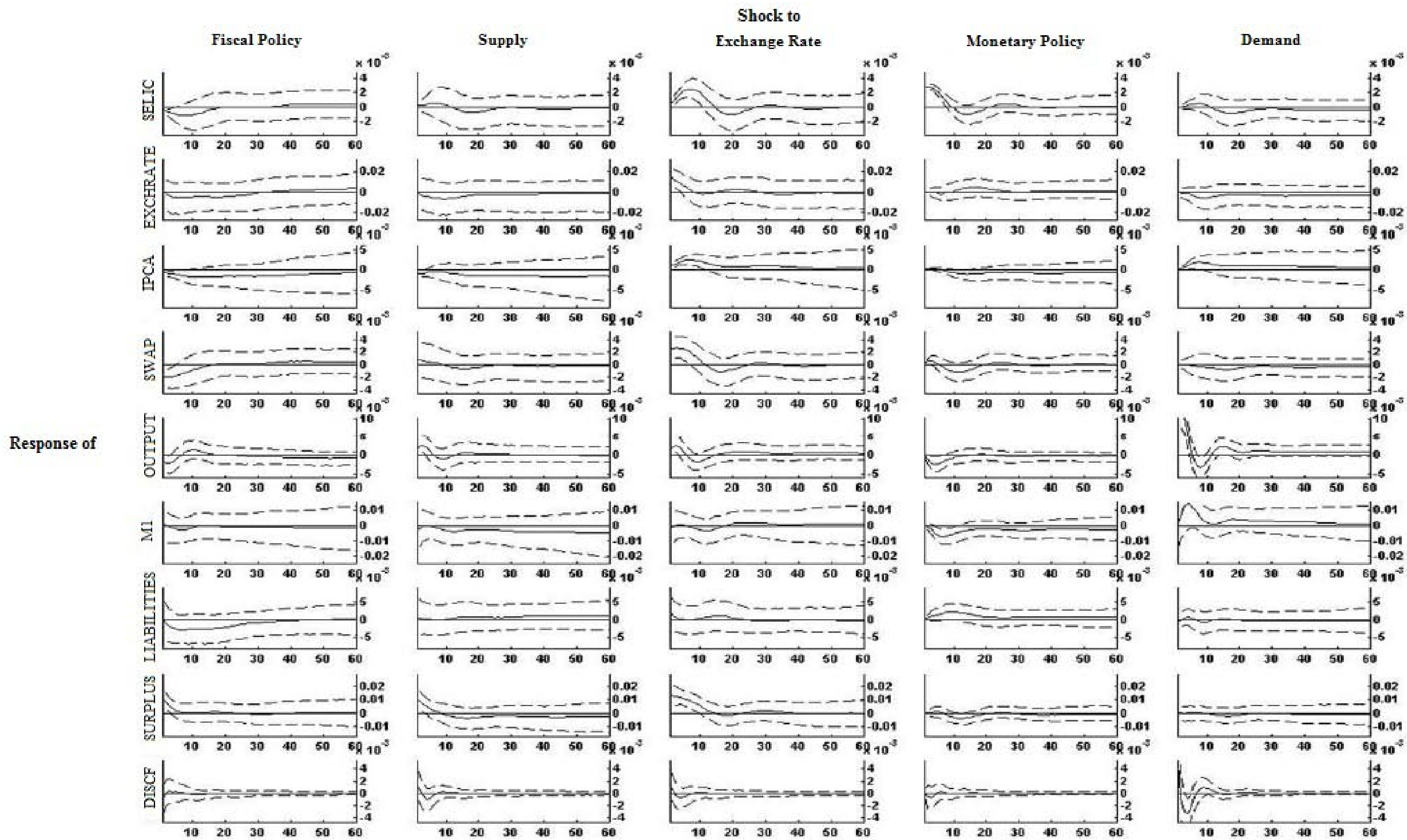


Figure 4.4: IRFs based on the hybrid identification (60 months ahead), with 68% probability bands



identify all shocks are presented on figures 4.5-4.6, showing the median as well as the 68% probability bands for a horizon of 24 and 60 months following the shocks, respectively. The main differences with respect of the IRFs based on the hybrid identification rely on the responses to fiscal and monetary policy shocks. Using sign restrictions only, we observe a reduction on government's liabilities in response to fiscal shocks, which according to CCD's test is evidence of a Ricardian regime. Monetary policy now has an important role as a source of short-run fluctuations on output, prices, and the exchange rate. Monetary policy now controls the long run rate of inflation and there is still there is no evidence of the "step on a rake" effect.

Figure 4.5: IRFs based on the sign restrictions identification (24 months ahead), with 68% probability bands

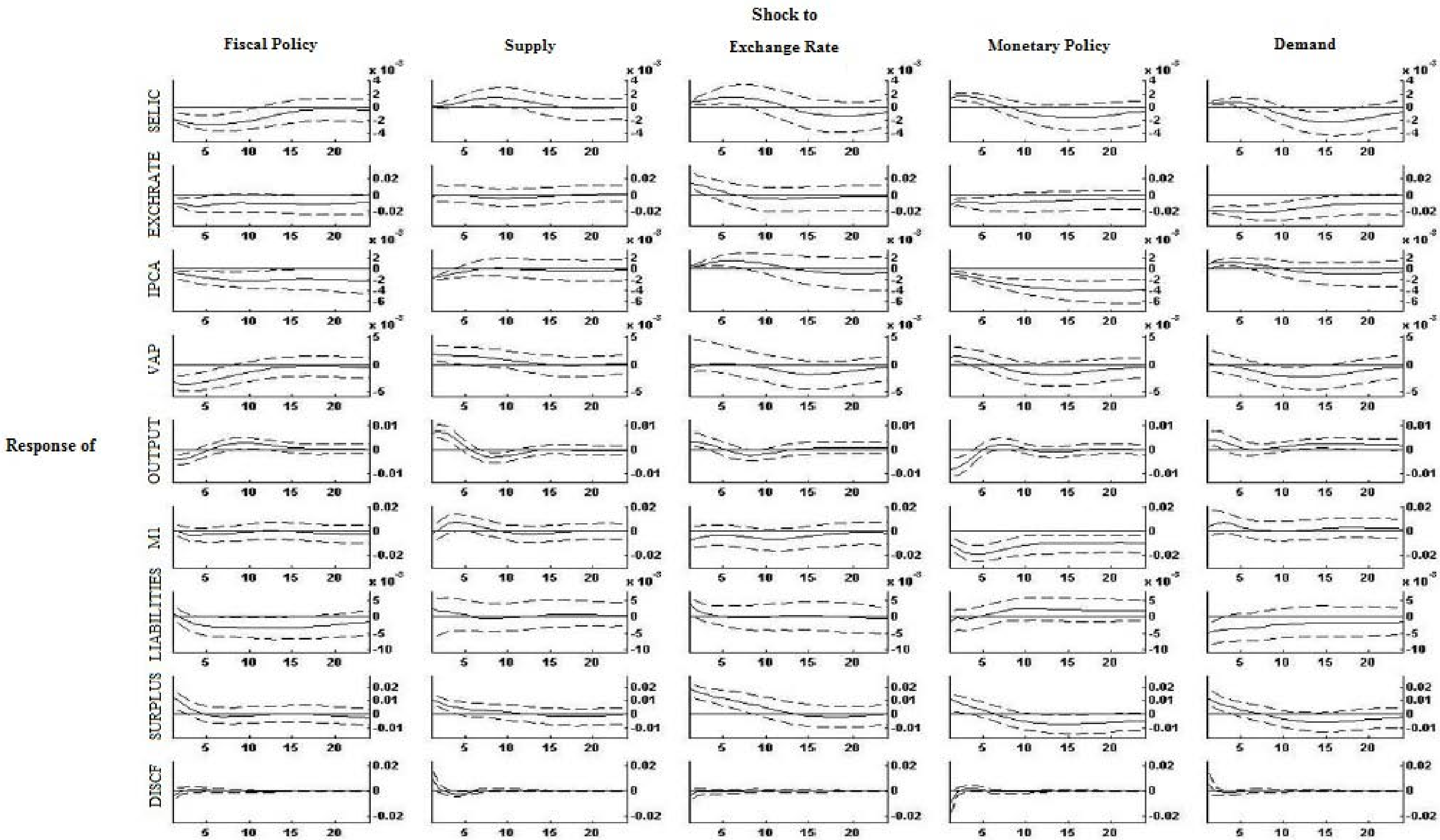
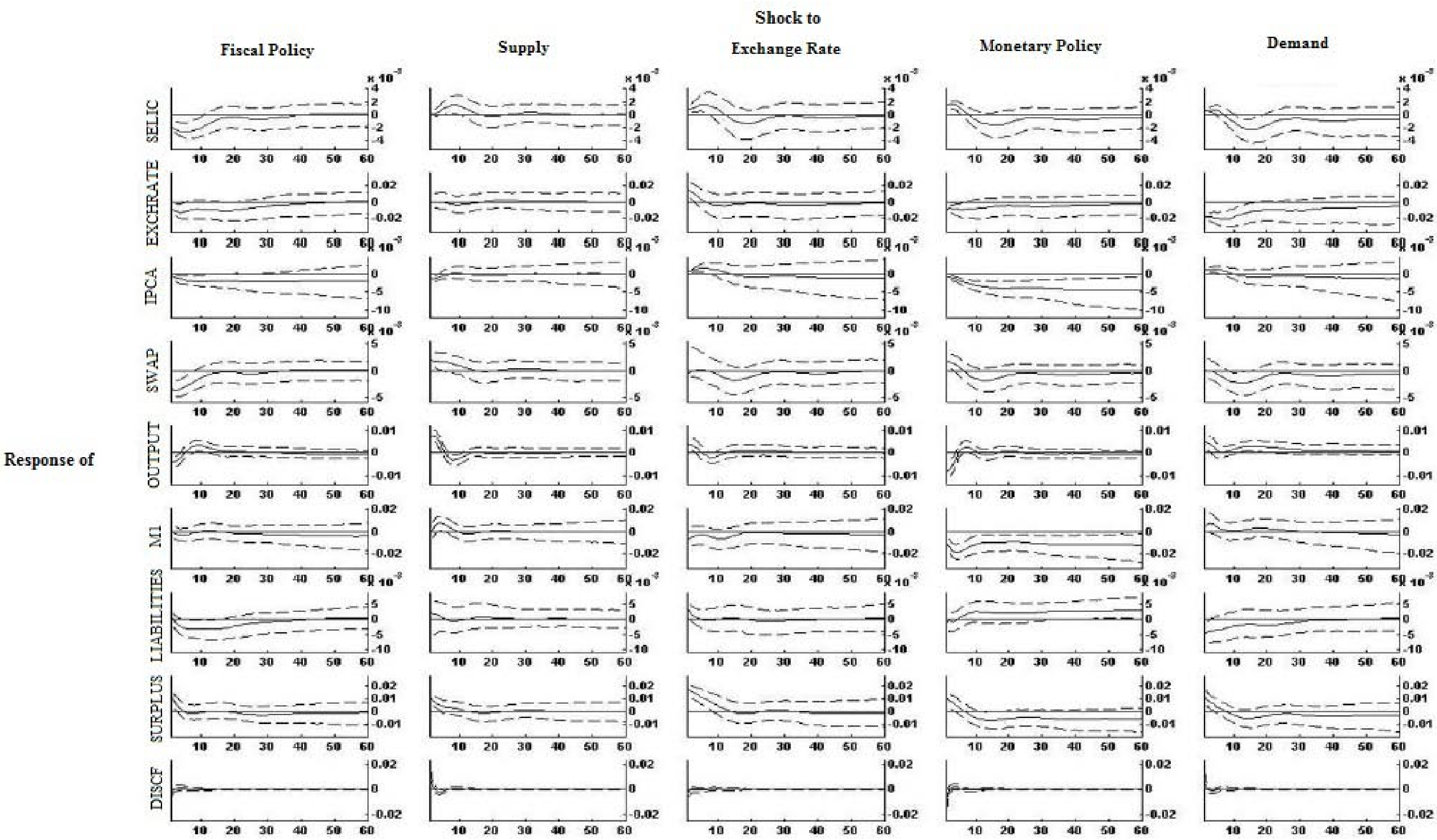


Figure 4.6: IRFs based on the sign restrictions identification (60 months ahead), with 68% probability bands



## 4.5 Concluding Remarks

While there is an agreement between most economists regarding the effects of monetary policy shocks, the empirical literature has struggled so far to provide robust stylized facts on the effects of fiscal policy shocks. In particular, there is no agreement on even the qualitative effects of fiscal policy shocks on macroeconomic variables.

This paper analyzed the effects of fiscal policy shocks and the interaction between fiscal and monetary policy. To achieve our goals we use a structural vector autoregression (SVAR) model and the test proposed by Canzoneri, Cumby, and Diba (2001). The SVAR is identified by two alternative methodologies. The first methodology used sign restrictions on impulse responses of the exogenous disturbances. The second methodology (hybrid) combined sign restrictions with restrictions on the contemporaneous causal interrelationships among variables, derived by Directed Acyclic Graphs (DAGs). Analyzing the case of Brazil, we observed for both identification strategies that in response to positive ("contracionist") fiscal shocks there is a significative and long-lasting reduction in the price level, and a short-lived reduction on economic activity. There is no evidence of significative response of the exchange rate to fiscal innovations.

Monetary and fiscal policy have two main objectives: controlling inflation and stabilizing the ratio of government debt to GDP. "Controlling inflation" means avoiding deviations of inflation from target and "stabilizing government debt" means maintaining the value of the ratio of the debt to GDP and preventing it from growing unsustainably. The conventional assignment gives monetary policy responsibility for controlling inflation and fiscal policy the role of stabilizing government debt (Monetary Dominance) ratio. In this case, since fiscal policy is assigned to stabilize debt, monetary policy is free to target inflation. However, the assignments can be reversed: fiscal policy can determine inflation, while monetary policy prevents debt from becoming unstable. This second regime can arise in crises or states of fiscal stress, and is the distinguishing assumption held by the Fiscal Theory of the Price Level (FTPL) [Fiscal Dominance]. The FTPL is a specific case of monetary-fiscal interaction and it challenges conventional-purely monetary-explanations of price level determination.

We tested for Brazil the assumption, held by the Fiscal Theory of the Price Level, that the Brazilian policy regime is non-Ricardian (Fiscal Dominance), applying the test proposed by CCD that analyzes the response of public sector liabilities to primary surplus shocks. This response depends on the identification adopted. For the hybrid identification we found that it is not possible to distinguish empirically between Ricardian (Monetary Dominance) and non-Ricardian (Fiscal Dominance) regimes. However, using sign restrictions there is some evidence that the government followed a Ricardian (Monetary Dominance) regime from January 2000 to June 2008.

We also checked if the identified exogenous monetary policy shocks show a "stepping on a rake" effect (tighter monetary policy leads to a higher inflation rate in the long run), as described by Sims (2008) in a theoretical framework designed for understanding the effects of fiscal uncertainties on monetary policy. According to our results, there is no evidence whatsoever that a tighter monetary policy would lead to higher inflation in the long run.

## 4.A Data Appendix

Short-term interest rate (SELIC): SELIC interest rate – adjusted average rate of daily financing guaranteed by federal government securities, calculated in the Special Settlement and Custody System (SELIC) and published by the Central Bank of Brazil (BCB) – annualized rate.

Nominal exchange rate: R$ / US$ – end of period buying rate – source: BCB.

Price index (IPCA): IPCA price index – source: IBGE.

Medium-term interest rate (SWAP): 180 days SWAP rate (PRE $\times$ CDI) – source: Brazilian Mercantile & Futures Exchange – annualized rate.

Output: the industrial production index – three month moving average – source: IBGE.

Monetary Aggregate: M1 – working days average – source: BCB.

Surplus: primary surplus of the consolidated public sector (includes central government, state and municipal governments, and public enterprises), as a ratio of the GDP – 12 months accumulated – source: BCB.

Public Sector Net Liabilities: consolidated public sector debt plus monetary base, as a ratio of the GDP – 12 months accumulated – source: BCB.

Discount factor $= \frac{Y_{t+1}/Y_t}{(1+i_t^*)}$, where $Y$ is monthly GDP reported by the BCB and $i^*$ is calculated as (nominal) interest payments – excluding the effect of exchange rate fluctuation – over public sector borrowing requirements.

# Technical Appendix

Let $y_t$ be the data vector – there are 9 variables in the model, therefore $y_t$ has dimension $n \times 1 (n = 9)$ for each period $t$:

$$y_t = [y_{1t} y_{2t} \cdots y_{nt}]$$

where

$y_{1t} = \log$ (Gross annualized Selic interest rate),

$y_{2t} = \log$ (Nominal exchange rate(R$/US$)),

$y_{3t} = \log$ (IPCA index),

$y_{4t} = \log$ (180 days Swap rate (PRE $\times$ CDI – annualized considering 252 working days)),

$y_{5t} = \log$ (Industrial Production Index),

$y_{6t} = \log$ (M1),

$y_{7t} = \log$ (government liabilities as ratio of the GDP),

$y_{8t} = \log$ (primary surplus as ratio of the GDP) and

$y_{9t} = \log$ (Discount factor).

The structural VAR model has the general form:

$$y_t' A_{t=1}'^{p} y_{t+1}' A_t' + z_t' D' + \varepsilon_t', \quad \text{for} \quad t = 1, \cdots, T \tag{4.5}$$

where

$y_t$ is an $n \times 1$ column vector of endogenous variables at time $t$,

$A$ and $A_t$ are $n \times n$ parameter matrices,

$D$ is an $n \times h$ parameter matrix,

$z_t$ is an $h \times 1$ column vector of seasonal dummies and constant term at time $t$,

$\varepsilon_t$ is an $n \times 1$ column vector of structural disturbances at time $t$;

$p$ is the lag length, and $T$ is the sample size ($p = 6$ and $T = 103$).

The parameters of individual equations in (1) correspond to the columns of $A'$, $A_t'$ and $D'$.

The structural disturbances have a Gaussian distribution with $E(\varepsilon_t | y_t, \ldots, y_{t-1}, z_1, \ldots, z_T) = 0_{n\times 1}$ and $E(\varepsilon_t \varepsilon_t' | y_1, \ldots, y_{t-1}, z_1, \ldots, z_T) = I_{n\times n}$ and then are normalized to have an identity covariance matrix. Right multiplying the structural form (1) by $(A'^{-1}$, we will obtain the usual representation of a reduced-form VAR with the reduced-form variance matrix being $\Omega = (A'^{-1}$

Unlike typical unrestricted VAR models, $\Omega$ will be restricted when the contemporaneous parameter matrix $A$ is overidentified.

The structural VAR models (4.5) can be rewritten in the compact form:

$$y_t'A' = x_t'F' + \varepsilon_t'$$

where

$$x'_{t_{1\times k}} = \left[y'_{t-1}\cdots y'_{t-p}z'_t\right], \qquad F_{n\times k} = [A_1 \cdots A_p D]$$

and $k = np+h$. We will refer to $F'$ as lagged parameters even though $F'$ may also contain exogenous parameters.

For $1 \leq i \leq n$ let $a_i$ be the $i$'th column of $A'$, let $f_i$ be the $i$'th column of $F'$ and let $T_i$ be an $n \times n$ matrix of rank $q_i$. The linear restrictions of interest can be summarized as follows:

$$T_i a_i = 0, \qquad i = 1, \cdots, n \tag{4.6}$$

The restrictions given by (4.6) are said to be non-degenerate if there exists at least one non-singular matrix $A'$ satisfying them. In this paper, all restrictions are assumed to be non-degenerate.

When VAR models are large and degrees of freedom are low, the likelihood function itself can be ill behaved and there is the well-known tendency of estimates to become unreliable. To deal with these problems, Litterman (1986) introduces a widely used Bayesian prior distribution for reduced-form models to down-weight models with large coeficients on distant lags and explosive dynamics. Sims and Zha (1998a) incorporate Litterman's idea in the structural framework by specifying the prior distribution of $a_i$ and $f_i$ as

$$a_i \sim N\left(0, \bar{S}_t\right) \text{ and } f_i|a_i \sim N\left(\bar{P}_i a_i, \bar{H}_i\right) \tag{4.7}$$

where $\bar{H}_i$ is defined as an $k \times k$ diagonal, symmetric and positive definite (SPD) matrix:

$$\bar{H}_i = \begin{bmatrix} \begin{matrix} \frac{\lambda_0\lambda_1}{\sigma_i} & 0 & \cdots & 0 \\ 0 & \frac{\lambda_0\lambda_1}{\sigma_i} & 0 & \vdots \\ \vdots & 0 & \ddots & 0 \\ 0 & \cdots & 0 & \frac{\lambda_0\lambda_1}{\sigma_i 6^{\lambda_3}} \\ & 54\times 54 & & \end{matrix} & 0_{54\times 12} \\ 0_{12\times 54} & \begin{matrix} \frac{\lambda_0\lambda_4}{\sigma_i} & 0 & \cdots & 0 \\ \frac{\lambda_0\lambda_4}{\sigma_i} & 0 & \vdots & \\ \vdots & 0 & \ddots & 0 \\ 0 & \cdots & 0 & \frac{\lambda_0\lambda_4}{\sigma_i} \\ & 12\times 12 & & \end{matrix} \end{bmatrix}_{(66\times 66)}$$

The standard deviation of the conditional prior of $f_i$ (subset of parameters of equation 1) for the coefficient on lag $l$ of the variable $j$, is given by

$$\frac{\lambda_0\lambda_1}{\sigma_i l^{\lambda_3}}$$

where the hyperparameter $\lambda_0$ controls the tightness of beliefs on $A'$; $\lambda_1$ controls what Litterman called overall tightness of beliefs around the random walk prior; $\lambda_3$ controls the rate at which prior variance shrinks for increasing lag length; $\lambda_4$ is the tightness for the constant term and seasonal dummies, i.e., for the last 12 rows of each column of $F'$. We give it a conditional prior mean of zero and a standard deviation controlled by $\lambda_0\lambda_4$.

The vector of parameters $\sigma_1, \ldots, \sigma_n$ (one for each equation) are scale factors, allowing for the fact that the units of measurement or scale of variation may not be uniform across variables. The scale factors are taken as the sample standard deviations of residuals from univariate autoregressive models, with lag length $p$, fit to the individual series in the sample.

The diagonal matrix $S_i$ is an $n \times n$ positive semidefinite matrix, the individual elements in the $i$'th column of $A'$ are assumed independent, with prior standard deviations set to $\lambda_0/\hat{\sigma}_i$ (parameters defined above):

$$\bar{S}_{i_{(n\times n)}} = \begin{bmatrix} \frac{\lambda_0}{\hat{\sigma}_1} & 0 & 0 & \cdots & 0 \\ 0 & \frac{\lambda_0}{\hat{\sigma}_2} & 0 & \cdots & \vdots \\ \vdots & 0 & \frac{\lambda_0}{\hat{\sigma}_3} & & \\ & \vdots & & \ddots & 0 \\ 0 & 0 & \cdots & 0 & \frac{\lambda_0}{\hat{\sigma}_n} \end{bmatrix}$$

We use the following values for the hyperparameters:

| Hyperparameter | Value |
|---|---|
| $\lambda_0$ | 0.5 |
| $\lambda_1$ | 0.25 |
| $\lambda_3$ | 1 |
| $\lambda_4$ | 0.5 |

$\bar{P}_i$is a $k \times n$ matrix defined as:

$$\bar{P}_i = \begin{bmatrix} I_{9\times 9} \\ 0_{57\times 9} \end{bmatrix}$$

The prior form summarized above represents a class of existing Bayesian priors that have been widely used for structural VAR models. Combining the prior form (4.7) with the restriction (4.6), we wish to obtain the functional form of the conditional prior distribution:

$$q\left(a_i, f_i | T_i a_i = 0\right) \tag{4.8}$$

In our case, the following matrices are the restricted $A$ and $A'$ matrices obtained by the application of the TETRAD software, together with the assumption that the swap rate affects the selic rate contemporaneously (swap $\rightarrow$ selic):

$$A = \begin{bmatrix} A_{11} & 0 & 0 & A_{14} & 0 & 0 & 0 & 0 & 0 \\ 0 & A_{22} & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & A_{33} & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & A_{42} & 0 & A_{44} & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & A_{55} & A_{56} & 0 & 0 & A_{59} \\ 0 & 0 & 0 & 0 & A_{65} & A_{66} & A_{67} & A_{68} & 0 \\ 0 & A_{72} & A_{73} & 0 & 0 & A_{76} & A_{77} & 0 & A_{79} \\ 0 & 0 & 0 & A_{84} & 0 & A_{86} & 0 & A_{88} & 0 \\ 0 & 0 & 0 & 0 & A_{95} & 0 & A_{97} & 0 & A_{99} \end{bmatrix}$$

$$A' = \begin{bmatrix} A_{11} & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & A_{22} & 0 & A_{42} & 0 & 0 & A_{72} & 0 & 0 \\ 0 & 0 & A_{33} & 0 & 0 & 0 & A_{73} & 0 & 0 \\ A_{14} & 0 & 0 & A_{44} & 0 & 0 & 0 & A_{84} & 0 \\ 0 & 0 & 0 & 0 & A_{55} & A_{65} & 0 & 0 & A_{95} \\ 0 & 0 & 0 & 0 & A_{56} & A_{66} & A_{76} & A_{86} & 0 \\ 0 & 0 & 0 & 0 & 0 & A_{67} & A_{77} & 0 & A_{97} \\ 0 & 0 & 0 & 0 & 0 & A_{68} & 0 & A_{88} & 0 \\ 0 & 0 & 0 & 0 & A_{59} & 0 & A_{79} & 0 & A_{99} \end{bmatrix}$$

Then, we can obtain the $T_i'$s matrices which satisfy the constraints for each column $i$ of $A'$:

$$T_{i_{q_i \times n}} a_{i_{n \times 1}} = 0_{q_i \times 1}$$

Each matrix $T_i$ reproduces the restrictions present in column $i$ of $A'$, given by TETRAD. All element of $T_i$ off the diagonal are zero. At the diagonal, there are zeros in the position of free parameters and ones in the position of parameters restricted to be equal zero. Therefore, for example

$$T_1 = \begin{bmatrix} 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 1 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 1 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 1 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 1 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 1 \end{bmatrix} \text{ and } T_9 = \begin{bmatrix} 1 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 1 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 1 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 1 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \end{bmatrix}$$

Let $U_i$ be an $n \times q_i$ matrix whose columns form an orthonormal basis for the null space of $T_i$. The column $a_i$ will satisfy the restriction (4.6) if and only if there exist a $q_i \times 1$ vector $b_i$ ($q_i$ = number of free parameter at column $i$ of matrix $A'$) such that

$$a_i = U_i b_i \tag{4.9}$$

The column vector $b_i$ contains the free parameters of column $i$ of matrix $A'$ given by TETRAD. For this matrix $A'$ the $U_i'$s are given by,

$$U' = \begin{bmatrix} U_1' \\ U_2' \\ U_3' \\ U_4' \\ U_5' \\ U_6' \\ U_7' \\ U_8' \\ U_9' \end{bmatrix} = \begin{bmatrix} 1 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 1 & 0 & 0 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 1 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 1 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 1 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 1 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 1 \\ 0 & 0 & 0 & 0 & 1 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 1 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 1 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 0 & 1 & 0 \\ 0 & 1 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 1 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 1 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 1 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 1 \\ 0 & 0 & 0 & 1 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 1 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 0 & 1 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 1 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 1 \end{bmatrix}$$

For example,

$$a_4 = \begin{bmatrix} 0 \\ A_{42} \\ 0 \\ A_{44} \\ 0 \\ 0 \\ 0 \\ 0 \\ 0 \end{bmatrix} = U_4 b_4 = \begin{bmatrix} 0 & 0 \\ 1 & 0 \\ 0 & 0 \\ 0 & 1 \\ 0 & 0 \\ 0 & 0 \\ 0 & 0 \\ 0 & 0 \\ 0 & 0 \end{bmatrix} \begin{bmatrix} A_{42} \\ A_{44} \end{bmatrix}$$

The distributions of $b_i$ and $f_i$ are given by

$$b_i \sim N\left(0, \tilde{S}_i\right) \text{ and } f_i | b_i \sim N\left(\tilde{P}_i b_i, \tilde{H}_i\right) \tag{4.10}$$

where

$$\tilde{H}_i = \bar{H}_i, \qquad \tilde{P}_i = \bar{P}_i U_i, \qquad \text{and} \qquad \tilde{S}_i = \left(U_i' \bar{S}_i^{-1} U_i\right)^{-1}$$

Note that $\tilde{S}_i$ is a $q_i \times q_i$ positive semidefinite matrix, $\tilde{H}_i$ is an $r_i \times r_i$ positive semidefinite matrix, and $\tilde{P}_i$ is a $r_i \times q_i$ matrix. It can be verified that the prior distribution (4.10) for bi is equivalent to the prior distribution (4.8) for $a_i$. For the most part of this paper, we work directly with bi with the understanding that the original parameters ai can be easily recovered via the linear transformations $U_i$.

Let $b = [b_1' \dots b_n']'$, $f = [f_1' \dots f_n']'$, $X = [x_1 \dots x_T]'$, and $Y = [y_1 \dots y_T]'$. The likelihood function for $b$ and $f$ $(l((b, f)|X, Y))$ is proportional to

$$|\det [U_1 b_1| \cdots |U_n b_n]|^T \exp\left(-\frac{1}{2}\sum_{i=1}^{n} b_i' U_i' Y' Y U_i b_i - 2 f_i' X' Y U_i b_i + f_i' X' X f_i\right) \tag{4.11}$$

Combining the priors on $b$ and $f$ given by (4.10) with the likelihood function given by (7) leads to the following joint posterior probability distribution function for $b$ and $f$:

$$p(b_1, \cdots, b_n|X, Y)\, \Pi_{i=1}^{n} p(f_i|b_i, X, Y)$$

where

$$p(b_1, \dots, b_n|X, Y) \propto |\det [U_1 b_1| \cdots |U_n b_n]|^T \exp\left(-\frac{T}{2}\sum_{i=1}^{n} b_i' S_i^{-1} b_i\right) \tag{4.12}$$

$$p(f_i|b_i, X, Y) = \phi(P_i b_i, H_i) \tag{4.13}$$

with

$$H_i = \left(X'X + \tilde{H}_i^{-1}\right)^{-1}$$

$$S_i = \left(\frac{1}{T}\left(U_i' Y' Y U_i + \tilde{S}_i^{-1} + \tilde{P}_i' \tilde{H}_i^{-1} \tilde{P}_i - P_i' H_i^{-1} P_i\right)\right)^{-1}$$

Since (8) has an unknown distribution, we must take draws from the posterior distribution of $b$ by Gibbs Sampling and, and given each draw of $b$, take draws of $f$ from the Gaussian conditional distribution (4.13). The notation $\phi(P_i b_i, H_i)$ in (4.13) denotes the Gaussian density with mean $P_i b_i$ and covariance matrix $H_i$.

In many works with VARs, only the likelihood function (i.e., proportional to the posterior density under a flat prior for $b$ and $f$) is considered. Because (4.11) is the same as (4.12) and (4.13) when the prior variances (diagonal elements in $\tilde{S}_i$ and $\tilde{H}_i$) approach infinity, the posterior density specified in (4.12) and (4.13) includes the likelihood as a special case.

To obtain small-sample inferences of $b$ and $f$ or for functions of them (e.g., impulse responses), it is necessary to simulate the joint posterior distribution of $b$ and $f$. This simulation involves two consecutive steps. First, simulate draws of $b$ from the marginal posterior distribution (4.12). Second, given each draw of $b$, simulate draws of $f$ from the conditional posterior distribution (4.13). The second step is straightforward because it requires draws only from a multivariate normal distribution. The first step, as mentioned earlier, can be challenging when linear restrictions on $A$ imply a restricted reduced-form covariance matrix.

The following algorithm was designed to obtain a sample of the impulse response functions, which satisfy the sign restrictions.

**Algorithm:** The following steps compose the algorithm for simulating draws from the posterior distribution of $b$, $f$ and, given these draws, draws of the impulse responses that satisfy the sign restrictions.

1. [1.]

2. Get the values at the peak of the posterior density function.

3. For $s = 1, \ldots, N_1$ and given $b_1^{(s-1)}$ obtain $b_1^{(s)}, \ldots, b_n^{(s)}$ by
a. simulating $b_1^{(s)}$ from the distribution $b_1|b_2^{(s-1)}, \ldots, b_n^{(s-1)}$,
b. simulating $b_2^{(s)}$ from $b_2|b_1^{(s)}, b_3^{(s-1)}, \ldots, b_n^{(s-1)}$,
$\vdots$
c. simulating $b_n^{(s)}$ from $b_n|b_1^{(s)}, \ldots, b_{n-1}^{(s)}$.

4. Keep $b_1^{(N_1)}, \ldots, b_n^{(N_1)}$.

5. For $s = N_1 + 1, N_2$ and given $b_1^{(s-1)}, \ldots, b_n^{(s-1)}$, obtain $b_1^{(s)}, \ldots, b_n^{(s)}$ by
d. simulating $b_1^{(s)}$ from the distribution $b_1|b_2^{(s-1)}, \ldots, b_n^{(s-1)}$,
e. simulating $b_2^{(s)}$ from $b_2|b_1^{(s)}, b_3^{(s-1)}, \ldots, b_n^{(s-1)}$,
$\vdots$
f. simulating $b_n^{(s}$ from $b_n|b_1^{(s)}, \ldots, b_{n-1}^{(s)}$.
g. Given $b_1^{(s)}, \ldots, b_n^{(s)}$ simulate $f_1^{(s)}, \ldots, f_n^{(s)}$ from the conditional normal distribution described in equation (9).
h. Given $b_1^{(s)}, \ldots, b_n^{(s)}$ and $f_1^{(s)}, \ldots, f_n^{(s)}$ obtain $A^{(s)}$ and $B^{(s)} = F^{(s)}A^{(s)-1}$ ($A$ and $F$ were described previously – $B$ contains the reduced form parameters).
i. Draw an independent standard normal $n \times n$ matrix $\tilde{X}$ and let $\tilde{X} = \tilde{Q}\tilde{R}$ be the $QR$ decomposition of $\tilde{X}$ with the diagonal $\tilde{R}$ normalized to be positive.
j. Let $P = \tilde{Q}$ and generate the impulse responses $IRF^{(s)}$ from $A^{(s)}P$ and $B^{(s)}P = F^{(s)}A^{(s)-1}P$.
k. If $IRF^{(s)}$ satisfies the sign restrictions keep it, otherwise discard it.
l. If the number of accepted $IRF$ is equal to 1000 stop.

6. Collect all the $IRF$ that were not discarded in step 4.

In step 2 and 4 of the Algorithm, all simulations are carried out according Theorem 2 of Waggoner and Zha (2002). The central result of Theorem 2 states that drawing from the distribution of $b_i$ conditional on $b_1, \ldots, b_{i-1}, b_n$ is equivalent to drawing from a multivariate Gaussian distribution and a special univariate distribution.

For a fixed $i*$, where $1 \leq i* \leq n$. Let $w$ be an non-zero $n \times 1$ vector perpendicular to each vector in $\{U_i b_i | i \neq i*\}$. Since the restrictions are assumed to be non-degenerate, the $n-1$ vectors $U_i b_i$ for $i \neq i*$ will almost surely be linearly independent and $U'_{i*}w$ will be non-zero. Define $w_1 = T'_{i*}U'_{i*}w/||T'_{i*}U'_{i*}w||$, where $T_{i*}$ is a $q_{i*} \times q_{i*}$ matrix such that $T_{i*}T'_{i*} = S_{i*}$, and choose $w_2, \ldots, w_{q*}$ so that $w_1, w_2, \ldots, w_{q*}$ form an orthonormal basis for $R^{q*}$. Then the random vector $b_i$ conditional on $b_1, \ldots, b_{i-1}, b_{i+1}, \ldots, b_n$ can be represented as

$$b_i = \beta_1 U_i' T_i^{-1} w_1 + \sum_{j=2}^{qi} \beta_j U_i' T_i^{-1} w_j$$

The random variable $\beta_j$, for $2 \leq j \leq q_i$, is normally distributed with mean zero and variance $1/T$ and is straightforward to simulate. The density function for $\beta_1$, the special univariate distribution, is proportional to $|\beta_1|^T \exp(-T\beta_1^2/2)$. Waggoner and Zha (2002) show how to simulate from this latter distribution.

**Hybrid Identification**[23]

Suppose we want to keep the identification of the first shock obtained by TETRAD (the monetary policy shock). Then we have to modify matrix $P$ employed in step 4-j of the previous algorithm. It will take the hybrid form:

$$P = \tilde{Q} = \begin{bmatrix} 1 & 0 & 0 & \cdots & 0 \\ 0 & \tilde{Q}_{22} & \tilde{Q}_{32} & \cdots & \tilde{Q}_{92} \\ 0 & \tilde{Q}_{23} & \tilde{Q}_{33} & \cdots & \tilde{Q}_{93} \\ \vdots & \vdots & \vdots & \ddots & \vdots \\ 0 & \tilde{Q}_{29} & \tilde{Q}_{39} & \cdots & \tilde{Q}_{99} \end{bmatrix}$$

where the submatrix,

$$Q_s = \begin{bmatrix} \tilde{Q}_{22} & \tilde{Q}_{32} & \cdots & \tilde{Q}_{92} \\ \tilde{Q}_{23} & \tilde{Q}_{33} & \cdots & \tilde{Q}_{93} \\ \vdots & \vdots & \ddots & \vdots \\ \tilde{Q}_{29} & \tilde{Q}_{39} & \cdots & \tilde{Q}_{99} \end{bmatrix}$$

is obtained by a draw of an independent standard normal $(n-1) \times (n-1)$ matrix $\tilde{X}$, and $Qs$ is obtained by the $QR$ decomposition of $\tilde{X}(\tilde{X} = Q_s\tilde{R}$, with the diagonal $\tilde{R}$ normalized to be positive).

[23] A discussion of the differences between our hybrid identification methodology and that of Dungey and Fry (2009) is presented on Section 1.

# Chapter 5

# Conclusions

We showed that the New Keynesian macroeconomic model of economic fluctuations, despite being based on microeconomic foundations with optimizing agents and rational expectations, can be nevertheless considered implausible because their empirical implications fail to match those of the available data. The evidence seems to be consistent with the (backward-looking) Keynesian model of economic fluctuations which is often criticized as "starting from curves" rather than starting from deep parameters, based on tastes and technology, from which behavioral relationships can be derived.

There is no consensus on two key Phillips curves issues. First, are inflation expectations forward-looking or backward-looking? If expectations are forward-looking, future events (including changes in monetary policy) can influence the current inflation rate. If, instead, expectations are backward-looking, inflation has inertia. Such inertia affects the design of monetary policy. Second, what is a proper measure of inflationary pressures—the output gap or the real marginal cost?

In this dissertation we tested Phillips curves using an ecompassing framework. We conclude that the NKPC does not provide a useful description of the inflation process in the cases of the U.S. and Brazil. The evidence rejects the restrictions implied by the NKPC, the HPC, and SIPC, but does not reject those of the APC. Regarding the two fundamental Phillips curves issues, our results suggest that inflation expectations are backward-looking and that the output gap is a better measure of inflationary pressures than the real marginal cost.

While Phillips curves has been subject of extensive empirical investigation, few studies have analysed IS curves.We tested IS curves using also an encompassing framework. The evidence is consistent with the KISC but not with the NKISC or the HISC.

Monetary and fiscal policy have two main objectives: controlling inflation and stabilizing the ratio of government debt to GDP. "Controlling inflation" means avoiding deviations of inflation from target and "stabilizing government debt" means maintaining the value of the ratio of the debt to GDP and preventing it from growing unsustainably. The conventional assignment gives monetary policy responsibility for controlling inflation and fiscal policy the role of stabilizing government debt (Monetary Dominance) ratio. In this case, since fiscal policy is assigned to stabilize debt, monetary policy is free to target inflation. However, the assignments can be reversed: fiscal policy can determine inflation, while monetary policy prevents debt from becoming unstable. This second regime can arise in crises or states of fiscal stress, and it is the distinguishing assumption held by the Fiscal Theory of the Price Level (FTPL) [Fiscal Dominance]. The FTPL is a specific case of monetary-fiscal interaction and it challenges conventional-purely monetary-explanations of price level determination.

We tested for Brazil the assumption, held by the Fiscal Theory of the Price Level, that the Brazilian policy regime is non-Ricardian (Fiscal Dominance), applying the test proposed by Canzoneri, Cumby, and Diba (2001) that analyzes the response of public sector liabilities to primary surplus shocks. This response depends on the identification adopted. For the hybrid identification we found that it is not possible to distinguish empirically between Ricardian (Monetary Dominance) and non-Ricardian (Fiscal Dominance) regimes. However, using sign restrictions there is some evidence that the government followed a Ricardian (Monetary Dominance) regime from January 2000 to June 2008.

# Bibliography


[1] Areosa, W. and Medeiros, M. (2007). "Inflation Dynamics in Brazil: The Case of a Small Open Economy". *Brazilian Review of Econometrics* 27(1), 131–166.

[2] Baba, K., Shibata, R., and Sibuya, M. (2004). "Partial Correlation and Conditional Correlation as Measures of Conditional Independence". *Australian & New Zealand Journal of Statistics* 46(4): 657-664.

[3] Ball, L. (1999). "Policy Rules for Open Economies". In John Taylor (ed.), *Monetary Policy Rules.* University of Chicago Press.

[4] Ball, L. and Mankiw, N.G. (2002). "The NAIRU in Theory and Practice". *Journal of Economic Perspectives, 16, Fall, 115–136.*

[5] Ball, L. and Mazumber, S. (2011). "Inflation Dynamics and the Great Recession". *Brookings Papers on Economic Activity*, Spring, 337–381.

[6] Barbosa, F. (2013a). *Teoria Macroeconômica.* EPGE/FGV. Unpublished.

[7] Barbosa, F. (2013b). Taking Stock of Keynesian Models of Fluctuations. Mimeo. EPGE/FGV.

[8] Bastos, E. (2012). "Distribuição Funcional da Renda no Brasil: Estimativas Anuais e Construção de Uma Série Trimestral". IPEA Texto para Discussão no. 1702.

[9] Baum, C., Schaffer, M., and Stillman, S. (2003). "Instrumental Variables and GMM: Estimation and Testing". *Stata Journal* 3(1), 1–31.

[10] Baum, C., Schaffer, M., and Stillman, S. (2007). "Enhanced Routines for Instrumental Variables/Generalized Method of Moments Estimation and Testing". *Stata Journal* 7(4), 465–506.

[11] Blanchard, O. and Galí, J. (2007). "Real Wage Rigidities and the New Keynesian Model". *Journal of Money, Credit and Banking*, 39(1), 35–65.

[12] Brüggemann, R., Lütkepohl, H. (2001). "Lag Selection in Subset VAR Models with an Application to a U.S. Monetary System", in R. Friedmann, L. Knüppel and H. Lütkepohl (eds), *Econometric Studies: A Festschrift in Honour of Joachim Frohn*, 107–128.

[13] Bryan, M. and Cecchetti, S. (1994). "Measuring Core Inflation". In *Monetary Policy*, edited by N. Gregory Mankiw. University of Chicago Press.

[14] Caldara, D. and Kamps, C. (2008). “What Are the Effects of Fiscal Policy Shocks? A VAR-Based Comparative Analysis”. European Central Bank Working Paper no. 877.

[15] Calvo, G. A. (1983). “Staggered Prices in a Utility-Maximizing Framework”. *Journal of Monetary Economics*, 12, 383–398.

[16] Campos, J., Ericsson, N., and Hendry, D. (2005). "General-to-specific Modeling: An Overview and Selected Bibliography". Board of Governors of the Federal Reserve System. International Finance Discussion Papers Number 838.

[17] Campos, C. and Nakane, M. (2003). "Phillips Curve and the Effects of Nominal Shocks in Open Economies: The Role of Price Setting. Mimeo.

[18] Canzoneri, M., Cumby, R., and Diba, B. (2001). “Is the Price Level Determined by the Needs of Fiscal Solvency?”. *American Economic Review* 91(5): 1221-1238.

[19] Carlstrom, C. and Fuerst, T. (2000). “The Fiscal Theory of the Price Level”. Federal Reserve Bank of Cleveland, *Economic Review*, 36(1): 22-32.

[20] Carvalho, R. (2008). *A Equação de Euler e a IS: Estimações para o Brasil.* Master’s Thesis. EPGE/FGV. Unpublished.

[21] Christiano, L. and Fitzgerald, T. (2000). “Understanding the Fiscal Theory of the Price Level”. NBER Working Paper no. 7668.

[22] Clarida, R., Galí, J., and Gertler, M. (1999). "The Science of Monetary Policy: A New Keynesian Perspective". Journal of Economic Literature, Vol. XXXVII, 1661–1707.

[23] Cysne, R. (1985). "A Relação de Phillips no Brasil: 1964-66 x 1980-84". *Revista Brasileira de Economia*, 39(4), 401–422.

[24] Dennis, R. (2009). "Consumption Habits in a New Keynesian Business Cycle Model". *Journal of Money, Credit and Banking*, 41(5), 1015–1030.

[25] Doornik, J. (2009). “Autometrics”. In J. L. Castle and N. Shephard (eds.), *The Methodology and Practice of Econometrics: A Festschrift in Honour of David F. Hendry.* Oxford University Press, 88–121.

[26] Doornik, J., and Hendry, D. (2009). *PcGive 13.* Timberlake Consultants Ltd.

[27] Dungey, M.. and Fry, R. (2009). “The Identification of Fiscal and Monetary Policy in a Structural VAR”. *Economic Modelling*, doi:10.1016/j.econmod.2009.05.001.

[28] Estrella, A., Fuhrer, J. (2002). "Dynamic Inconsistencies: Counterfactual Implications of a Class of Rational Expectations Models". *American Economic Review* 92, 1013–1028.

[29] Fasolo, A. and Portugal, M. (2004). "Imperfect Rationality and Inflationary Inertia: a New Estimation of the Phillips Curve for Brazil". *Estudos Econômicos*, 34(4), 725–776.

[30] Fialho, M. and Portugal, M. (2005). "Monetary Policy and Fiscal Policy Interactions in Brazil: an Application of the Fiscal Theory of the Price Level". *Estudos Econômicos* 35(4): 657-685.

[31] Friedman, M. (1968). "The Role of Monetary Policy". *American Economic Review*, 58, 1–17.

[32] Fry, R. and Pagan, A. (2007). "Some Issues in Using Sign Restrictions for Identifying Structural VARs". NCER Working Paper no.14.

[33] Fuhrer, J. (2000). "Habit Formation in Consumption and its Implications for Monetary Policy Models". *American Economic Review* 90, 367–390.

[34] Fuhrer, J., Rudebusch, G. (2004). "Estimating the Euler Equation for Output". *Journal of Monetary Economics* 51, 1133–1153.

[35] Galí, J. (2008). *Monetary Policy, Inflation, and the Bussiness Cycle*. Princeton University Press.

[36] Galí, J. and Gertler, M. (1999). "Inflation Dynamics: A Structural Econometric Analysis". *Journal of Monetary Economics*, 44, 195–222.

[37] Galí, J., Gertler, M., and Lópes-Salido, J. (2001). "European Inflation Dynamics". *European Economic Review*, 45(7), 1237–1270.

[38] Galí, J. and Monacelli, T. (2005). "Monetary Policy and Exchange Rate Volatility in a Small Open Economy". *Review of Economic Studies*, 72(3), 707–734.

[39] Gordon, R. (1982). "Inflation, Flexible Exchange Rates, and the Natural Rate of Unemployment." In *Workers, Jobs, and Inflation*, edited by Martin Neil Baily. Brookings.

[40] Hamilton, J. (1995). *Time Series Analysis*. Princeton University Press.

[41] Hammond, G. (2010). "State of the Art of Inflation Targeting". Bank of England. CCBS Handbook No. 29.

[42] Holland, M. and Mori, R. (2010). "Dinâmica da Inflação no Brasil e os Efeitos Globais". Revista Economia, 13(3), 649–670.

[43] Kiley, M. (2007). "A Quantitative Comparison of Sticky-Price and Sticky-Information Models of Price Setting". *Journal of Money, Credit and Banking*, 39(1), 101–125.

[44] King, R. and Watson, M. (2012). "Inflation and Unit Labor Cost". *Journal of Money, Credit and Banking* , 44(2), 111–149.

[45] Leeper, E. (2010). "Monetary Science, Fiscal Alchemy". Paper presented at the Jackson Hole Symposium.

[46] Lima, E. (2003). "The NAIRU, Unemployment and the Rate of Inflation in Brazil". *Revista Brasileira de Economia*, 57(4), 899–930.

[47] Lima, E., Maka, A., and Alves, P. (2011). "Monetary Policy and Exchange Rate Shocks in Brazil: Sign Restrictions versus A New Hybrid Identification Approach". *Brazilian Review of Econometrics*, 31(1), 97–136.

[48] Lima, E. and Brito, L. (2009). "Inflação e Nível de Atividade no Brasil: Estimativas via Curva de Phillips". In *Setor Externo e Inflação,* edited by Renault Michel and Leonardo Carvalho. IPEA.

[49] Lima, E., Maka, A., and Céspedes, B. (2008). "Monetary Policy, Inflation, and the Level of Economic Activity in Brazil After the Real Plan: Stylized Facts from SVAR Models". *Revista Brasileira de Economia*, 62(2): 123–150.

[50] Loyo, E. (1999). "Tigh Money Paradox on the Loose: A Fiscalist Hyperinflation". Mimeo. Kennedy School of Government, Harvard University.

[51] Mankiw, N. G. and Reis, R. (2002). "Sticky Information vs. Sticky Prices: A Proposal to Replace the New Keynesian Phillips Curve". *Quarterly Journal of Economics*, CXVII, 1295–1328.

[52] Mazali, J. and Divino, A. (2010). "Real Wage Rigidity and the New Phillips Curve: The Brazilian Case". *Revista Brasileira de Economia*, 64(3), 291–306.

[53] McCallum, B. (1989). *Monetary Economics: Theory and Policy.* Macmillan.

[54] McCallum, B. and Nelson, E. (2000). "Monetary policy for an open economy: an alternative framework with optimising agents and sticky prices". *Oxford Review of Economic Policy,*16, 4, 74–91.

[55] Mendonça, M., Sachsida, A., and Medrano, L. (2012). "Inflação Versus Desemprego: Novas Evidências para o Brasil". *Economia Aplicada*, 16(3), 471-494.

[56] Minella, A., Freitas, P., Goldfajn, I., and Muinhos, M. (2003). "Inflation Targeting in Brazil: Constructing Credibility Under Exchange Rate Volatility". *Journal of International Money and Finance*, 22(7), 1015–1040.

[57] Monacelli, T. (2005). "Monetary Policy in a Low Pass-Through Environment". *Journal of Money Credit and Banking*, 37(6), 1047–1066.

[58] Nason, J. and Smith, G. (2008a). "The New Keynesian Phillips Curve: Lessons From Single-Equation Econometric Estimation". *Economic Quarterly,* 94(4), 361–395.

[59] Patterson, K. (2000). *An Introduction to Applied Econometrics: a Time Series Approach.* St. Martin's Press.

[60] Peersman G. and Smets, F. (1999). "The Taylor Rule: A Useful Monetary Policy Benchmark for the Euro Area?". *International Finance*, 1, 85-116.

[61] Robins, J., Sheines, R., Spirtes, P., and Wasserman, L. (2003). "Uniform Consitency in Causal Inference". Biometrika 90(3): 491-515.

[62] Rocha, F. and Silva, E. (2004). "Teoria Fiscal do Nível e Preços: Um Teste para a Economia Brasileira no Período 1966-2000". *Pesquisa e Planejamento Econômico* 34(3): 419-435.

[63] Rotemberg, J., Woodford, M. (1999). "Interest Rate Rules in an Estimated Sticky Price Model". In: Taylor, J. (Ed.), *Monetary Policy Rules.* University of Chicago Press, 57–119.

[64] Rudd, J. and Whelan, K. (2007). “Modelling Inflation Dynamics: A Critical Survey of Recent Research”. *Journal of Money Credit and Banking*, 39, 155–170.

[65] Rudd, J. and Whelan, K. (2006). “Can Rational Expectations Sticky-Price Models Explain Inflation Dynamics?”. *American Economic Review*, 96 (1), 303-320.

[66] Rudebusch, G. and Svensson, L. (1999). "Policy Rules for Inflation Targeting". In John Taylor (ed.), *Monetary Policy Rules.* University of Chicago Press.

[67] Santos, F. and Holland, M. (2008). "Estimando a Demanda Agregada no Brasil: o Papel dos Fatores Externos". Proceedings of the 36th Brazilian Economic Association Annual Meeting.

[68] Schwartzman, F. (2006). "Estimativa de Curva de Phillips Para o Brasil Com Preços Desagregados". *Economia Aplicada*, 10(1), 137–155.

[69] Sims, C. (2008). “Stepping on a Rake: The Role of Fiscal Policy in the Inflation of the 1970’s”. Mimeo. Princeton University. Paper available at http://sims.princeton.edu/yftp/FiscalTheoryGreatInflation/BOJpresentation.pdf .

[70] Sims, C. (1994). “A Simple Model for Study of the Determination of the Price Level and the Interaction of Monetary and Fiscal Policy”. *Economic Theory* 4: 381-399.

[71] Sims, C. and Zha, T. (1998). “Bayesian Methods for Dynamic Multivariate Models”. *International Economic Review*, 39: 949-968.

[72] Spirtes, P., Glymour, C., and Scheines, R. (2000). *Causation, Prediction, and Search* (second edtion). MIT Press.

[73] Stock, J., and Watson, M. (2007). *Introduction to Econometrics* (second edition). Addison-Wesley.

[74] Stock, J., and Watson, M. (2002). "Has the Business Cycle Changed and Why?". NBER Working Paper No. 9127.

[75] Tanner, E. and Ramos, A. (2002). “Fiscal Sustainability and Monetary versus Fiscal Dominance: Evidence from Brazil, 1991-2000”. IMF Working Paper WP/02/05.

[76] Walsh, C. (2010). *Monetary Theory and Policy* (third edition). MIT Press

[77] Woodford, M. (2003). *Interest and Prices: Foundations of a Theory of Monetary Policy.* Princeton University Press.

[78] Zhang, J. (2002). Consistency in Causal Inference under a Variety of Assumptions. Masters Thesis. Department of Philosophy, Carnegie Mellon University.

[79] Zhang, J. and Spirtes, P. (2003). “Strong Faithfulness and Uniform Consistency in Causal Inference”. Proceedings of the 19th Annual Conference on Uncertainty in Artificial Intelligence. Morgan Kaufmann Publishers: 632-639.