Identifying the Monetary Transmission Mechanism using Structural Breaks[1]

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Abstract

We estimate the dynamics of a three equation model of the monetary transmission mechanism using data from 1970:Q1 to 1999:Q4. We find a significant break in the parameters of an estimated VAR in 1979:Q3 when Volcker took over as chairman of the Fed. We estimate a quasi-reduced form in which we allow for expectations to enter the private sector equations. We identify the private sector using the assumption that these equations do not break at the time of the change in policy. We present impulse response functions identified by our identification scheme and compare them with more traditional approaches.

1 Introduction

In this paper we introduce a method for estimating the structural equations of a rational expectations model by exploiting the assumption that economic policy is governed by a rule that may change infrequently and abruptly. Our idea is to use changes in the parameters of the policy rule as instruments to identify the parameters of structural equations that remain constant across different policy regimes. We apply our technique to the problem of identifying the monetary transmission mechanism.

For many economic data sets it is difficult or impossible to find extended periods over which the parameters of a vector autoregression remain constant. We believe that in a substantial number of examples of interest, it may be possible to make credible arguments that parameter change can be attributed to a specific cause. For example, the date of the break may coincide with an announcement by the policy authorities or with a change in administration. This is the situation that occurred in 1979 when Paul Volcker took over from G. William Miller as chairman of the Board of Governors of the Fed. Similar events occurred in 1973 with the collapse of the Bretton Woods system of fixed exchange rates and in 1999 when the European Monetary Union was created. In situations when structural change can credibly be attributed to a change in one structural equation, we show how to use this information to identify the remaining equations of the system.

2 Related Research

An early application of structural change to achieve identification is provided by Charles Bean [1]. The closest papers to the present one are by In Choi [8] and Roberto Rigobon [35], [36]. Choi uses structural breaks to show that a class of minimum distance estimators and a principal components instrumental variables estimator may lead to identification of a subset of parameters when classical identification procedures fail. Our work is closely related to this paper although Choi does not explicitly consider the case of rational expectations models. Rigobon exploits heteroskedasticity to identify the parameters of a structural model. Rigobon's work is also similar to our approach although we exploit changes in all of the parameters of the policy rule, not just a change in variance. In a related paper, Matt Klaeffling [24] estimates a class of models that includes future expectations, but the parameters of his model confound policy with structure and his approach cannot be applied to study the consequences of a change in the policy rule.

A number of recent papers analyze regime change in the context of U.S. monetary policy. Thomas Lubik and Frank Schorfheide [26], [27] develop a technique for differentiating between determinate and indeterminate policy regimes and apply their technique to U.S. monetary data. They place more structure on their model than we are prepared to do, although like Lubik and Schorfheide, our approach allows for both determinate and indeterminate equilibria. Eric Leeper and Tao Zha [25] have studied a class of "modest policy interventions" that enables them to make predictions of the effects of monetary policy changes within a given regime. We assume instead that the entire regime changes at a discrete point in time. Chris Sims and Tao Zha [41] study regime changes using a Markov switching model, and Timothy Cogley and Thomas Sargent [13] adopt a similar approach that allows for parameter drift. We discuss the relationship of our work to this literature in Section 7.4 in which we argue that the results of Sims and Zha, of heteroskedasticity across monetary policy regimes, might be interpretable in terms of structural changes in a monetary policy rule.

Our empirical application to structural changes in monetary policy is closely related to the large literature, surveyed by Lawrence Christiano, Martin Eichenbaum and Charles Evans [9], on the use of structural VAR's to estimate the monetary transmission mechanism. Recent prominent contributions to this literature include papers by Ben Bernanke and Ilian Mihov [3] and Lawrence Christiano, Martin Eichenbaum and Charles Evans [10]. Our work differs from these more conventional identification procedures in two ways. First, we include expectations in the private sector structure and second, we exploit a structural break to achieve identification rather than relying on assumptions about causal ordering or on the long-run effects of shocks.

Bernanke and Mihov estimate models of regime change and find that the best indicator of monetary policy stance differs across regime. Our own analysis complements their approach. Although we consider only a single indicator of monetary policy, the fed funds rate, our estimates are potentially of more use to the policy maker since we claim to identify parameters of the "deep structure" that are, conceptually, invariant to changes in the policy rule. This issue was first pointed to by John Keating [23] who argued that most standard SVAR identification schemes will yield inconsistent estimates since they confound structural parameters with expectational effects. Our estimates are not subject to the "Keating critique" since we explicitly account for expectations in our structural estimates.

3 Development of Notation

In this section we define a notation and we illustrate its generality by means of two examples. The general notation is useful for the discussion of identification that we take up in Section 4. The examples are useful since they flag some of the issues that will arise when we apply our technique to a practical problem: the identification of the monetary transmission mechanism.

3.1 A Class of Models

We consider a class of economic models with a *structural form* represented by the equation:

$$\mathbf{A}X_t = \mathbf{B}Z_{t-1} + U_t, \tag{1}$$

$$E_{t-1}\left[U_t U_{t+k}^T\right] = \begin{cases} \mathbf{\Omega} & k = 0, \\ 0 & k \neq 0, \end{cases} \tag{2}$$

$$rank\left(\mathbf{\Omega}\right) \leq n, \tag{3}$$

where X_t is a vector of endogenous variables, Z_{t-1} is a vector of exogenous or predetermined variables and U_t is a vector of random variables that are independent and identically distributed.[1] We further partition $\mathbf{\Omega}$ into submatrices $\mathbf{\Omega}_{VV}$, $\mathbf{\Omega}_{WV}$ and $\mathbf{\Omega}_{VW}$:

$$\mathbf{\Omega} = \begin{bmatrix} \mathbf{\Omega}_{VV} & \mathbf{0} & \mathbf{\Omega}_{VW} \\ \mathbf{0} & \mathbf{0} & \mathbf{0} \\ \mathbf{\Omega}_{WV} & \mathbf{0} & \mathbf{\Omega}_{WW} \end{bmatrix}. \tag{4}$$

[1] Uppercase letters denote vectors, lowercase letters represent scalars and boldface letters are matrices. We use superscript "T" for the transpose operator.

Our reasons for partitioning $\mathbf{\Omega}$ in this way will become clear when we discuss special cases of this notation in Sections 3.2 and 3.3.

We assume that the model has a *reduced form* which we write as

$$X_t = \mathbf{\Gamma} Z_{t-1} + \epsilon_t, \tag{5}$$

$$E_{t-1}\left[\epsilon_t \epsilon_t^T\right] = \mathbf{\Sigma}, \tag{6}$$

and we partition $\mathbf{\Sigma}$ as

$$\mathbf{\Sigma} = \begin{bmatrix} \mathbf{\Sigma}_{VV} & \mathbf{0} & \mathbf{\Sigma}_{VW} \\ \mathbf{0} & \mathbf{0} & \mathbf{0} \\ \mathbf{\Sigma}_{WV} & \mathbf{0} & \mathbf{\Sigma}_{WW} \end{bmatrix}.$$

We introduce the notation

$$\theta = vec\left([\mathbf{A}, \mathbf{B}, \mathbf{\Omega}]^T\right), \quad \phi = vec\left([\mathbf{\Gamma}, \mathbf{\Sigma}]^T\right),$$

to refer to the *structural parameters* $\theta \in \Theta$ and the *reduced form parameters*, $\phi \in \Phi$ where Θ and Φ are subsets of finite dimensional Euclidian space. Given this notation we make the following assumption.

Assumption 1 (Uniqueness) *Every structural model $\theta \in \Theta$ has a unique reduced form $\phi \in \Phi$. The mapping from the structure to the reduced form can be represented by a function $g : \Theta \to \Phi$ such that $\phi = g(\theta)$.*

For most situations, Assumption 1 is uncontroversial although there is an important case in which it has bite; this is the case of a rational expectations model with multiple indeterminate solutions. We deal with this issue in Subsection 3.2 in which we introduce the rational expectations model and explain how we propose to deal with possible non-uniqueness of the solution. In Subsection 3.3 we show how to write the rational expectations model in the form of an error-correction model. This extension allows us to deal with both stationary and non-stationary data.

3.2 The Linear Rational Expectations Model

Consider the following linear rational expectations model in a vector of variables $Y_t \in R^n$

$$\tilde{\mathbf{A}} Y_t = \tilde{\mathbf{F}} E_t\left[Y_{t+1}\right] + \sum_{j=1}^{k} \tilde{\mathbf{B}}_j Y_{t-j} + V_t, \tag{7}$$

$$E_{t-1}\left[V_t V_{t+k}^T\right] = \begin{cases} \mathbf{\Omega}_{VV} & k = 0, \\ 0 & k \neq 0, \end{cases}$$

and define

$$\mathbf{A} = \begin{bmatrix} \tilde{\mathbf{A}} & -\tilde{\mathbf{B}}_1 & \cdots & -\tilde{\mathbf{B}}_k & -\tilde{\mathbf{F}} \\ \mathbf{0} & \mathbf{I} & \cdots & \mathbf{0} & \mathbf{0} \\ \vdots & \vdots & \ddots & \vdots & \vdots \\ \mathbf{0} & \mathbf{0} & \cdots & \mathbf{I} & \mathbf{0} \\ \mathbf{I} & \mathbf{0} & \cdots & \mathbf{0} & \mathbf{0} \end{bmatrix}, \quad \mathbf{B} = \begin{bmatrix} \mathbf{0} & \mathbf{0} & \cdots & \mathbf{0} & \mathbf{0} & \mathbf{0} \\ \mathbf{I} & \mathbf{0} & \cdots & \mathbf{0} & \mathbf{0} & \mathbf{0} \\ \vdots & \vdots & \ddots & \vdots & \vdots & \vdots \\ \mathbf{0} & \mathbf{0} & \cdots & \mathbf{I} & \mathbf{0} & \mathbf{0} \\ \mathbf{0} & \mathbf{0} & \cdots & \mathbf{0} & \mathbf{0} & \mathbf{I} \end{bmatrix},$$

$$X_t = \begin{bmatrix} Y_t \\ Y_{t-1} \\ \vdots \\ Y_{t-k} \\ E_t\left(Y_{t+1}\right) \end{bmatrix}, \quad Z_{t-1} = \begin{bmatrix} Y_{t-1} \\ Y_{t-2} \\ \vdots \\ Y_{t-k-1} \\ E_{t-1}\left(Y_t\right) \end{bmatrix}, \quad U_t = \begin{bmatrix} V_t \\ 0 \\ \vdots \\ 0 \\ W_t \end{bmatrix},$$

$$W_t \equiv \left[Y_t - E_{t-1}\left(Y_t\right)\right],$$

$$E_{t-1}\left(W_t W_{t+k}^T\right) = \begin{cases} \mathbf{\Omega}_{WW} & k = 0, \\ 0 & k \neq 0. \end{cases}$$

Using these definitions the rational expectations model fits into the structure of Equation (1) although the matrix $\mathbf{A}$ may have reduced rank and some of the generalized eigenvalues of $\{\mathbf{A}, \mathbf{B}\}$ may be outside the unit circle.[2] Nevertheless, as long as the rational expectations model has at least one stationary solution Equation (5) will be well defined and it is always possible to find a function $g : \Theta \to \Phi$, by choosing values for the endogenous component of U_t (these are the variables W_t) that keep the vector X_t bounded. In our example, this solution would correspond to finding a matrix $\mathbf{M}$ such that when

$$W_t = \mathbf{M} V_t,$$

[2]The generalized eigenvalues of a pair of matrices $\{\mathbf{A}, \mathbf{B}\}$ are values of λ that solve the problem $|\mathbf{A} - \lambda \mathbf{B}| = 0$. For the equation $\mathbf{A} Y_t = \mathbf{B} Y_{t-1}$ the generalized eigenvalues of $\{\mathbf{A}, \mathbf{B}\}$ determine the asymptotic behavior of Y_t. Generalized eigenvalues are useful in computing solutions to rational expectations problems in which both $\mathbf{A}$ and $\mathbf{B}$ have reduced rank. See Golub and Van Loan [18] pages 375ff.

the influence of unstable roots of the system is eliminated.[3]

In the context of this example, Assumption 1 may appear restrictive since it is well known that linear rational expectations models may possess multiple solutions. We take the view that, if a given structure leads to multiple reduced forms, then the structural specification is incomplete. If there are multiple reduced forms then the model must be supplemented by a full specification of the process by which expectations are formed or by a selection mechanism such as the expectational stability criterion suggested by Evans and Honkapohja [15].[4]

3.3 An Error Correction Representation of the Rational Expectations Model

By writing the model in first differences, including lagged levels, and finding an appropriate transformation of the parameter matrices it is possible to rewrite the rational expectations model (7) in a form, similar to that of a standard error correction model;

$$\bar{\mathbf{A}}\Delta Y_t = \bar{\mathbf{F}}\Delta E_t\left[Y_{t+1}\right] + \sum_{j=1}^{k-1}\bar{\mathbf{B}}_j\Delta Y_{t-j} + \mathbf{\Pi} Y_{t-1} + V_t, \tag{8}$$

where Δ is the difference operator and the matrices $\bar{\mathbf{A}}, \bar{\mathbf{F}}, \bar{\mathbf{B}}_j$ and $\mathbf{\Pi}$ are defined as follows

$$\begin{aligned} \mathbf{\Pi} &= \mathbf{F}+\sum_{j=1}^{k}\mathbf{B}_j - \mathbf{A}, \qquad \bar{\mathbf{A}} = \mathbf{A} - \mathbf{F}, \\ \bar{\mathbf{B}}_j &= -\sum_{i=j+1}^{k}\mathbf{B}_i, \qquad \bar{\mathbf{F}} = \mathbf{F}. \end{aligned}$$

The error-correction representation of the system is useful because it allows stationary and non-stationary data to be treated in a unified way. Stationary, non-stationary and cointegrated data can all be modeled by Equation (8). If Y_t is stationary then $rank\left(\mathbf{\Pi}\right) = n$. In a non-stationary world, $rank\left(\mathbf{\Pi}\right) = 0$. In a

[3] There are many possible algorithms to compute the reduced form; we refer the reader to Sims [39] who provides a discussion of the relationship between θ and ϕ for the general linear rational expectations model.

[4] In the empirical application to U.S. monetary policy this issue does not arise since our estimates lead to a determinate equilibrium in both policy regimes. In a related paper, Beyer Farmer [5], we show how to handle the case of an indeterminate equilibrium.

non-stationary but cointegrated environment $\mathbf{\Pi}$ has reduced rank $(r \leq n)$ and it can be decomposed as

$$\mathbf{\Pi} = \boldsymbol{\alpha}\boldsymbol{\beta}^T,$$

where $\boldsymbol{\alpha}$ is an $n \times r$ matrix of loading factors and $\boldsymbol{\beta}^T$ is an $r \times n$ matrix of cointegrating vectors.

The error-correction model fits into the structure of Equation (1) with the following definitions of X_t, Z_{t-1} and the matrices $\mathbf{A}$ and $\mathbf{B}$:

$$X_t = \begin{bmatrix} \Delta Y_t \\ \Delta Y_{t-1} \\ \vdots \\ \Delta Y_{t-k+1} \\ E_t\left[Y_{t+1}\right] - Y_t \end{bmatrix}, \quad Z_{t-1} = \begin{bmatrix} \Delta Y_{t-1} \\ \Delta Y_{t-2} \\ \vdots \\ \Delta Y_{t-k} \\ E_{t-1}\left[Y_t\right] - Y_{t-1} \\ ci_{t-1} \end{bmatrix}, \quad ci_{t-1} \equiv \boldsymbol{\beta}^T Y_{t-1}, \tag{9}$$

$$\mathbf{A} = \begin{bmatrix} \bar{\mathbf{A}} & -\bar{\mathbf{B}}_1 & \cdots & -\bar{\mathbf{B}}_{k-1} & -\bar{\mathbf{F}} \\ \mathbf{0} & \mathbf{I} & \cdots & \mathbf{0} & \mathbf{0} \\ \vdots & \vdots & \ddots & \vdots & \vdots \\ \mathbf{0} & \mathbf{0} & \cdots & \mathbf{I} & \mathbf{0} \\ \mathbf{I} & \mathbf{0} & \cdots & \mathbf{0} & \mathbf{0} \end{bmatrix}, \quad \mathbf{B} = \begin{bmatrix} \mathbf{0} & \mathbf{0} & \cdots & \mathbf{0} & \mathbf{0} & \mathbf{0} & \boldsymbol{\alpha} \\ \mathbf{I} & \mathbf{0} & \cdots & \mathbf{0} & \mathbf{0} & \mathbf{0} & \mathbf{0} \\ \vdots & \vdots & \ddots & \vdots & \vdots & \vdots & \vdots \\ \mathbf{0} & \mathbf{0} & \cdots & \mathbf{I} & \mathbf{0} & \mathbf{0} & \mathbf{0} \\ \mathbf{0} & \mathbf{0} & \cdots & \mathbf{0} & \mathbf{0} & \mathbf{I} & \mathbf{0} \end{bmatrix}. \tag{10}$$

In our empirical application in Section 6, we will apply a model with the structure of Equation (8) to U.S. data on unemployment, inflation and the interest rate.

4 Optimal Policy and Identification: Why they are Related

This section is about identification and its importance to the goals of a policy maker. We explain why our approach is different from the literature on structural VARs initiated by Sims [40] and surveyed in Christiano Eichenbaum and Evans [9]. We will argue that although the structural VAR approach has many important strengths, it cannot be used to design an *optimal* policy rule since vector autoregressions are, by their nature, reduced form objects that confound the effects of private sector structural parameters with the coefficients of the policy maker's reaction function.[5]

4.1 Why Identification is an Issue for Policy

We define the goal of a policy maker to be the maximization of the discounted sum of a function $J(Y_t)$ that depends on the complete set of variables $Y_t \in R^n$. We partition Y_t into a set of state variables $Y_{1t} \in R^{n_1}$ and a set of control variables $Y_{2t} \in R^{n_2}$ where $n_1 + n_2 = n$ and we assume that the data is generated by a rational expectations model with a *true structure* parametrized by $\theta = vec\left([\mathbf{A}, \mathbf{B}, \mathbf{\Omega}]^T\right)$;

$$\begin{bmatrix} \mathbf{A}_{11} & \mathbf{A}_{12} \\ \mathbf{A}_{21} & \mathbf{A}_{22} \end{bmatrix} \begin{bmatrix} Y_{1t} \\ Y_{2t} \end{bmatrix} + \begin{bmatrix} \mathbf{A}_{13} & \mathbf{A}_{14} \\ \mathbf{A}_{23} & \mathbf{A}_{24} \end{bmatrix} \begin{bmatrix} E_t\left[Y_{1t+1}\right] \\ E_t\left[Y_{2t+1}\right] \end{bmatrix}$$
$$= \begin{bmatrix} \mathbf{B}_1 & 0 & 0 \\ \mathbf{B}_2 & 0 & 0 \end{bmatrix} \begin{bmatrix} Z_{t-1} \\ E_{t-1}\left[Y_{1t}\right] \\ E_{t-1}\left[Y_{2t}\right] \end{bmatrix} + \begin{bmatrix} V_{1t} \\ V_{2t} \end{bmatrix}, \qquad (11)$$

$$E\left[V_t V_t^T\right] = \mathbf{\Omega}_{VV}. \qquad (12)$$

Associated with the true structure, is the reduced form,

$$\overset{Y_t}{\begin{bmatrix} Y_{1t} \\ Y_{2t} \end{bmatrix}} = \overset{\mathbf{\Gamma}}{\begin{bmatrix} \mathbf{\Gamma}_{11} & \mathbf{\Gamma}_{12} \\ \mathbf{\Gamma}_{21} & \mathbf{\Gamma}_{22} \end{bmatrix}} Z_{t-1} + \overset{\epsilon_t}{\begin{bmatrix} \epsilon_{1t} \\ \epsilon_{2t} \end{bmatrix}}, \qquad (13)$$

parameterized by $\phi = g(\theta)$. We refer to the first n_1 equations of (11) as the *private-sector block* and the second n_2 equations as the *policy block*, for reasons described below.

The concept of a true structure is only meaningful if we allow ourselves to contemplate changes in θ. If some subset of θ changes, perhaps as a consequence of the design of an optimal policy, one can ask the question: How do the reduced form parameters depend on the structural parameters?

Parallel to our definition of private sector and policy blocks of equations, we propose the following partition of the parameters into *structural private sector parameters* θ_1 and *structural policy parameters* θ_2 where $(\theta_1 \cup \theta_2) \in \Theta$, and Θ

[5] This is, of course, just a restatement of Lucas' "Critique of Econometric Policy Evaluation" [28].

is the set of structural parameters;[6]

$$\begin{aligned}
\theta_1 &= vec\left([\mathbf{A}_{11}, \mathbf{A}_{12}, \mathbf{A}_{13}, \mathbf{A}_{14}, \mathbf{B}_1, \mathbf{\Omega}_{11}]^T\right), \\
\theta_2 &= vec\left([\mathbf{A}_{21}, \mathbf{A}_{22}, \mathbf{A}_{23}, \mathbf{A}_{24}, \mathbf{B}_2, \mathbf{\Omega}_{21}, \mathbf{\Omega}_{22}]^T\right).
\end{aligned}$$

Given this notation we define the following policy problem.

Problem 1 (Optimal Policy)

$$\max_{\theta_2} E_1 \sum_{t=1}^{\infty} \beta^t J(Y_t)$$

such that

$$\begin{aligned}
Y_t &= \mathbf{\Gamma}(\theta_1; \theta_2) Z_{t-1} + \epsilon_t, \\
E\left(\epsilon_t \epsilon_t^T\right) &= \mathbf{\Sigma}(\theta_1; \theta_2),
\end{aligned}$$

$\beta \in (0,1)$ *and* $J(\cdot)$ *is a concave increasing function.*

The notation $\mathbf{\Gamma}(\theta_1; \theta_2)$ and $\mathbf{\Sigma}(\theta_1; \theta_2)$ is an explicit expression of the fact that $\mathbf{\Gamma}$ and $\mathbf{\Sigma}$ are elements of ϕ, the reduced form parameters, and these elements are functions of the structural private sector parameters θ_1 and the structural policy parameters θ_2.

Identification is not a purely econometric issue; it is central to the design of an optimal policy. In order to solve Problem 1 the policy maker must know the functions $\mathbf{\Gamma}(\theta_1; \theta_2)$ and $\mathbf{\Sigma}(\theta_1; \theta_2)$ that determine how $\{Y_t\}_{t=1}^{\infty}$ varies with θ_2. Since we assume that the function $g(\theta_1, \theta_2)$ is known it is sufficient, to solve Problem 1, that the policy maker knows θ_1. But, in order to implement an optimal policy, does he need to know all of θ_1 or is knowledge of some smaller subset of parameters sufficient? The following section clarifies this question.

4.2 What the Policy Maker Needs to Know

The insight of this section is based on the fact that many structural models have the same reduced form. We will show that, although the policy maker cannot identify the parameters θ_1 of the true structure, there exists a second structural

[6] The inclusion of $\mathbf{\Omega}_{21}$ in the set of policy parameters allows for the assumption that the policy maker may respond contemporaneously to private sector shocks. We take up this point again in Section 7.2 in which we apply our method to the practical problem of identifying a policy rule in the context of the U.S. monetary policy transmission mechanism.

model, *which is identified*, that we call the *quasi-reduced form*. We denote the parameters of the quasi-reduced form with the notation $\bar{\theta}$ and we will show that $\phi = g\left(\bar{\theta}_1, \bar{\theta}_2\right)$; in other words, the quasi-reduced form and the true structural model have the same reduced form.

To control a set of variables Y_t the policy maker must know how these variables behave as functions of the subset Y_{2t} that are directly under his control. The policy maker does not need to know how subsets of the private sector variables Y_{1t} interact with each other. However, Y_{1t} must be a non trivial function of Y_{2t} and the parameters of this function must be known. The following assumption guarantees that these conditions can be met, in principle, if the policy maker has access to data from at least two distinct policy regimes.

Assumption 2 (Rank Condition) *a) The matrices $\mathbf{A}_{11}$ and $\mathbf{A}_{22}$ each have full rank. b) The elements of the matrix $\bar{\mathbf{A}}_{12} = \mathbf{A}_{11}^{-1}\mathbf{A}_{12}$ are all non-zero.*

Assumption 2 enables us to define the *quasi-reduced form*. The quasi-reduced form is a structural model that has the same reduced form as the true model. It has fewer parameters than the true reduced form and, more importantly, these parameters can in principle be recovered by observing data from two distinct policy regimes.

Definition 1 (Quasi-Reduced Form) *The quasi-reduced form is defined by the equations*

$$\begin{bmatrix} \mathbf{I} & \bar{\mathbf{A}}_{12} \\ \bar{\mathbf{A}}_{21} & \mathbf{I} \end{bmatrix} \begin{bmatrix} Y_{1t} \\ Y_{2t} \end{bmatrix} + \begin{bmatrix} \bar{\mathbf{A}}_{13} & \bar{\mathbf{A}}_{14} \\ \bar{\mathbf{A}}_{23} & \bar{\mathbf{A}}_{24} \end{bmatrix} \begin{bmatrix} E_t\left[Y_{1t+1}\right] \\ E_t\left[Y_{2t+1}\right] \end{bmatrix} = \begin{bmatrix} \bar{\mathbf{B}}_1 \\ \bar{\mathbf{B}}_2 \end{bmatrix} \begin{bmatrix} Z_{t-1} \end{bmatrix} + \begin{bmatrix} \bar{V}_{1t} \\ \bar{V}_{2t} \end{bmatrix}. \tag{14}$$

It is equal to the true model, premultiplied by the matrix

$$\mathbf{Q} = \begin{bmatrix} \mathbf{A}_{11}^{-1} & \mathbf{0} \\ \mathbf{0} & \mathbf{A}_{22}^{-1} \end{bmatrix}.$$

Proposition 1 *To solve problem 1 it is necessary and sufficient that the policy maker knows the parameters of the quasi-reduced form.*

Proof. *We first establish that the quasi-reduced form (14) and the true structure (11) have the same reduced form. Let*

$$\bar{\theta}_1 = vec\left(\left[\bar{\mathbf{A}}_{11}, \bar{\mathbf{A}}_{12}, \bar{\mathbf{A}}_{13}, \bar{\mathbf{A}}_{14}, \bar{\mathbf{B}}_1, \bar{\mathbf{\Omega}}_{11}\right]^T\right)$$

be the private-sector parameters of the quasi-reduced form and define

$$\bar{\phi} = g\left(\bar{\theta}_1, \bar{\theta}_2\right),$$

to be the reduced form parameters of the solution to (14). By definition, Equation (14) is equal to the structural model (11) multiplied by the matrix $\mathbf{Q}^{-1}$*. Since* $\mathbf{Q}^{-1}$ *is non-singular Equations (14) and (11) define an equivalence class of structural models and since, from Assumption 1,* $\phi = g\left(\bar{\theta}_1, \bar{\theta}_2\right)$ *exists and is unique,* $\bar{\phi} = \phi$ *hence the true structure and the quasi-reduced form define the same data generation process for* $\{Y_t\}$*. Proposition 1 is immediate since to solve Problem 1 it is necessary and sufficient that the policy maker can change the elements of* Y_t *through control of the elements of* Y_{2t} *and the history of* Y_t*. This is guaranteed by Assumption 2 and the definition of the quasi-reduced form.* ■

Notice that there exists a large class of structural models that are obtained by premultiplying Equation (14) by *any* non-singular matrix. The quasi-reduced form is the unique member of this class that removes all contemporaneous endogenous variables from the private sector structural equations. The existence of contemporaneous structural dependencies is irrelevant to the policy problem since the policy maker needs to know only how Y_{1t} depends on Y_{2t} and not how the elements of Y_{1t} depend on each other. The following section establishes that the definition of the quasi-reduced form is important since its parameters can, in principle, be identified by observing the effects of regime change.

5 Identification and Structural Breaks

In this section we explain our main idea – the identification of deep private sector structural parameters by exploiting structural breaks in policy. We formalize this idea for the case of a single discrete break.[7]

[7] Although our method could easily be extended to the more general case of multiple breaks we have not dealt with this issue for two reasons. First, the econometrics of testing for multiple breaks in empirical models of simultaneous equations has not been extensively developed (although some progress has been made, e.g. Pierre Perron and Zhongjun Qu, [32]). Second, the possibility of multiple breaks raises the interesting, but conceptually difficult issue, of how to model expectations in a world in which policy may change in predictable ways. We believe our methods could be extended to provide identification in structural models of regime switching. To date, models of this kind are either based on reduced form VAR models, an

5.1 Some Notation to Describe a Break

To set up our procedure we assume that the econometrician has access to a sample of data for dates $t = 1, ...T$, generated by a true structural model with quasi-reduced form (14) parameterized by $\left(\bar{\theta}_1, \bar{\theta}_2\right)$. We assume further that the policy sector parameters $\bar{\theta}_2$ are known to have changed at date τ (where $0 < \tau < T$), whereas the private sector parameters $\bar{\theta}_1$ remained constant.

To identify the parameters $\bar{\theta}_1$ we first define a vector of *step dummy variables*

$$d_t(\tau) = \mathbf{1}(t > \tau)$$

where $\mathbf{1}(\cdot)$ is the indicator function, equal to 1 if $t > \tau$ and 0 otherwise. Similarly, we define vectors of *slope dummy variables*,

$$Y_t^S = d_t(\tau) Y_t, \quad Z_t^S = d_t(\tau) Z_t.$$

Parallel to our introduction of dummy variables we will also need notation that describes possible changes in *all* of the elements of $\bar{\theta}$, the parameters of the quasi-reduced form. Let $\bar{\theta}$ represent these parameters over the first sub-sample, i.e. for $t \leq \tau$, and let $\bar{\theta}^S$ be the value of $\bar{\theta}$ over the subsequent sub-sample, i.e. for $t > \tau$. Now define the change in the parameter $\bar{\theta}$ across sub-samples with the notation

$$\bar{\theta}^{\Delta} = \bar{\theta}^S - \bar{\theta}.$$

Next, we will use our notation to increase the dimension of the model by adding dummy variables to the quasi-reduced form. We refer to the following equation,

$$\begin{bmatrix} \mathbf{I} & \bar{\mathbf{A}}_{12} & \mathbf{0} & \bar{\mathbf{A}}_{12}^{\Delta} \\ \bar{\mathbf{A}}_{21} & \mathbf{I} & \bar{\mathbf{A}}_{21}^{\Delta} & \mathbf{0} \end{bmatrix} \begin{bmatrix} Y_{1t} \\ Y_{2t} \\ Y_{1t}^S \\ Y_{2t}^S \end{bmatrix} + \begin{bmatrix} \bar{\mathbf{A}}_{13} & \bar{\mathbf{A}}_{14} & \bar{\mathbf{A}}_{13}^{\Delta} & \bar{\mathbf{A}}_{14}^{\Delta} \\ \bar{\mathbf{A}}_{23} & \bar{\mathbf{A}}_{24} & \bar{\mathbf{A}}_{23}^{\Delta} & \bar{\mathbf{A}}_{24}^{\Delta} \end{bmatrix} \begin{bmatrix} E_t\left[Y_{1t+1}\right] \\ E_t\left[Y_{2t+1}\right] \\ E_t\left[Y_{1t+1}^S\right] \\ E_t\left[Y_{2t+1}^S\right] \end{bmatrix}$$

$$= \begin{bmatrix} \bar{\mathbf{B}}_1 & \bar{\mathbf{B}}_1^{\Delta} \\ \bar{\mathbf{B}}_2 & \bar{\mathbf{B}}_2^{\Delta} \end{bmatrix} \begin{bmatrix} Z_{t-1} \\ Z_{t-1}^S \end{bmatrix} + \begin{bmatrix} \bar{V}_{1t} \\ \bar{V}_{2t} \end{bmatrix}, \tag{15}$$

$$E\left[\bar{V}_t \bar{V}_t^T\right] = \mathbf{\Omega}_{\bar{V}\bar{V}},$$

example is the work of Sims and Zha [41], or they deal with only very simple models as a consequence of inherent non-linearities involved in Markov switching models, as in the work of Troy Davig and Eric Leeper [14].

as the *augmented quasi-reduced form.* Associated with Equation (15) is the *augmented reduced form*

$$\begin{bmatrix} Y_{1t} \\ Y_{2t} \\ E_t\left[Y_{1t+1}\right] \\ E_t\left[Y_{2t+1}\right] \end{bmatrix} = \begin{bmatrix} \mathbf{\Gamma}_1 & \mathbf{\Gamma}_1^{\Delta} \\ \mathbf{\Gamma}_2 & \mathbf{\Gamma}_2^{\Delta} \\ \mathbf{\Gamma}_3 & \mathbf{\Gamma}_3^{\Delta} \\ \mathbf{\Gamma}_4 & \mathbf{\Gamma}_4^{\Delta} \end{bmatrix} \begin{bmatrix} Z_{t-1} \\ Z_{t-1}^{S} \end{bmatrix} + \begin{bmatrix} \epsilon_{1t} \\ \epsilon_{2t} \end{bmatrix}, \tag{16}$$

$$E\left(\epsilon_t \epsilon_t^T\right) = \begin{bmatrix} \mathbf{\Sigma}_{VV} & \mathbf{\Sigma}_{VW} \\ \mathbf{\Sigma}_{WV} & \mathbf{\Sigma}_{WW} \end{bmatrix}.$$

Armed with this notation we are ready to discuss identification.

5.2 Assumptions Required to Identify Parameters

In this subsection we introduce two assumptions and show that they are sufficient to guarantee that structural breaks can be used to identify the private sector parameters of the quasi-reduced form. We begin by describing the exclusion restrictions that arise from the assumption that $\bar{\theta}_1$ remains constant across regimes. These restrictions take the form;

$$\begin{aligned} \bar{\mathbf{A}}_{12}^{\Delta} &= 0, \quad \bar{\mathbf{A}}_{13}^{\Delta} = 0, \\ \bar{\mathbf{A}}_{14}^{\Delta} &= 0, \quad \bar{\mathbf{B}}_{1}^{\Delta} = 0. \end{aligned}$$

Since the elements of Z_{t-1}^{S} appear in the reduced form but are excluded from the private sector block of the quasi-reduced form, they can be used as instruments to identify the coefficients of $\bar{\mathbf{A}}_{12}, \bar{\mathbf{A}}_{13}$ and $\bar{\mathbf{A}}_{14}$.

To ensure that the equations of the quasi-reduced form are separately identified we make the following assumptions.

Assumption 3 (Row Independence) *The rows of*

$$\begin{bmatrix} \bar{\mathbf{A}} & \bar{\mathbf{B}} \end{bmatrix},$$

are linearly independent.

Assumption 3 is simply a statement that the equations in block 1 are different from each other. We also need to assume that there are enough excluded exogenous or predetermined variables to provide instruments for the included endogenous variables in the structural equations.

Assumption 4 (Instrument Sufficiency) *(i) The number of predetermined and exogenous variables, m is greater than or equal to $n_2 + n$. (ii) The matrix of reduced form parameters*

$$\begin{bmatrix} \mathbf{\Gamma}_2^{\Delta} \\ \mathbf{\Gamma}_3^{\Delta} \\ \mathbf{\Gamma}_4^{\Delta} \end{bmatrix}$$

has full row rank.

Assumption 4 means that there is sufficient parameter variation at date τ to provide independent instruments for the n_2 endogenous variables Y_{2t} and the $n_1 + n_2$ endogenous expectations $E_t[Y_{1t+1}]$ and $E_t[Y_{2t+1}]$.

Proposition 2 *Under assumptions 3 and 4 the parameters $\bar{\mathbf{A}}_{12}, \bar{\mathbf{A}}_{13}$ and $\bar{\mathbf{A}}_{14}$ of the quasi-reduced form are identified.*

Proof. The proof follows from applying the rank and order conditions, equation by equation, to the structural block of Equation (15). The order condition is implied by 4i since there are m excluded predetermined and exogenous variables (the variables Z_{t-1}^{Δ}) in each equation and $n_2 + n$ included endogenous variables (the n_2 values of Y_{2t} and the n values $E_t[Y_{t+1}]$). The sets of included endogenous variables are the same for each equation of the structural private block since the quasi-reduced form excludes the endogenous variables Y_{1t} from each of these equations by definition. Condition 4ii is just the rank condition (see Hamilton [20] page 246 Proposition 9.1 Condition c). ■

6 An Application to U.S. Monetary Policy

In this section we apply our method to U.S. monetary policy by studying a version of the New-Keynesian model of the monetary transmission mechanism as made popular amongst others by Clarida-Galí-Gertler [11], [12] Fuhrer and Rudebusch [16], Galí-Gertler [17], and Rotemberg and Woodford [37].

6.1 Identifying Parameters in a Three Equation Structural Model

Our version of the New-Keynesian model contains two private sector equations and a single policy equation. Our variables are u_t, the quarterly unemployment rate, π_t the inflation rate and i_t, the federal funds rate. We measure unemployment as an annualized percentage rate, the inflation rate by the annualized rate of change of the quarterly GDP deflator and i_t by the federal funds rate. We use quarterly data from 1960Q1 to 1999Q4.[8]

Using our notation the model can be represented as follows;.

$$\begin{bmatrix} \mathbf{I} & \overline{A}_{12} & \overline{\mathbf{A}}_{13} & \overline{A}_{14} \\ \overline{A}_{21} & \mathbf{1} & \overline{A}_{23} & \overline{a}_{24} \end{bmatrix} \begin{bmatrix} Y_{1t} \\ i_t \\ E_t\left[Y_{t+1}\right] \\ E_t\left[i_{t+1}\right] \end{bmatrix} = \begin{bmatrix} \overline{\mathbf{B}}_{12} \\ \overline{\mathbf{B}}_{21} \end{bmatrix} \left[Z_{t-1}\right] + \begin{bmatrix} \bar{V}_{1t} \\ \bar{V}_{2t} \end{bmatrix}, \quad (17)$$

$$Y_{1t} = \left(u_t, \pi_t\right)^T \quad Z_{t-1} = \left(Y_{t-1}, ..Y_{t-k}, i_{t-1}, .., i_{t-3}, 1\right)^T.$$

The first block of Equation (17) consists of an unemployment equation and an inflation equation. These equations can be derived by solving the New-Keynesian IS curve and Phillips curve simultaneously for a quasi-reduced form in which unemployment does not enter the inflation equation and inflation does not enter the unemployment equation. The second block consists of a single policy equation that represents an interest-rate rule.

We operationalize our method in this system by assuming a single known breakpoint. To this end, we set up the following augmented quasi-reduced form;

$$\begin{bmatrix} \mathbf{I} & \bar{A}_{12} & \mathbf{0} & \bar{A}_{12}^{\Delta} \\ \bar{A}_{21} & 1 & \bar{A}_{21}^{\Delta} & \mathbf{0} \end{bmatrix} \begin{bmatrix} Y_{1t} \\ i_t \\ Y_{1t}^S \\ i_t^S \end{bmatrix} + \begin{bmatrix} \bar{\mathbf{A}}_{13} & \bar{A}_{14} & \bar{\mathbf{A}}_{13}^{\Delta} & \bar{A}_{14}^{\Delta} \\ \bar{A}_{23} & \bar{a}_{24} & \bar{A}_{23}^{\Delta} & \bar{a}_{24}^{\Delta} \end{bmatrix} \begin{bmatrix} E_t\left[Y_{1t+1}\right] \\ E_t\left[i_{t+1}\right] \\ E_t\left[Y_{1t+1}^S\right] \\ E_t\left[i_{t+1}^S\right] \end{bmatrix}$$
$$= \begin{bmatrix} \bar{\mathbf{B}}_1 & \bar{\mathbf{B}}_1^{\Delta} \\ \bar{\mathbf{B}}_2 & \bar{\mathbf{B}}_2^{\Delta} \end{bmatrix} \begin{bmatrix} Z_{t-1} \\ Z_{t-1}^S \end{bmatrix} + \begin{bmatrix} \bar{V}_{1t} \\ \bar{V}_{2t} \end{bmatrix}. \quad (18)$$

To identify the private sector structure we impose zero restrictions on the coefficients of the shift variables i_{1t}^S, $E_t\left[Y_{1t+1}^S\right]$, $E_t\left[i_{t+1}^S\right]$ and Z_{t-1}^S in the private sector block. This is equivalent to the assumption that the parameters of the private sector are invariant to changes in the policy rule. Since we allow future expectation variables to enter these equations, our assumption does not violate Lucas' [28] critique of econometric policy evaluation.

[8] The New-Keynesian model is sometimes written in terms of the output gap, inflation and the interest rate. We chose instead to represent the level of economic activity by unemployment. This choice is consistent with our earlier work (Beyer-Farmer [4]) and it has the added advantage of sidestepping the difficult issue of how to detrend GDP.

No such assumption is available to allow us to identify the policy rule. Instead, we experimented with alternative identification schemes. In our benchmark model we excluded current and expected future endogenous variables from the policy rule. Since we did not restrict the covariance matrix of the shocks to be diagonal, this scheme implicitly allows the policy authorities to react contemporaneously to all current information. We elaborate on this idea further in section 7.4 in which we discuss our estimates of the variance-covariance matrix of the structural shocks for two different policy regimes.

6.2 Dealing with Persistence in the Data

Unemployment, the nominal interest rate and inflation are all highly persistent processes. In earlier work (see Beyer-Farmer [4]) we found that these data can be modeled as a cointegrated VAR. There we found evidence for two cointegrating vectors and we adopt this specification in our current structural model. We chose to maintain this representation in our current study by transforming the New-Keynesian model to an error correction representation with the structure of Equation (8).

Our identified model has the following form

$$\begin{bmatrix} \mathbf{I} & \bar{A}_{12} & \mathbf{0} & \mathbf{0} \\ \mathbf{0} & 1 & \mathbf{0} & \mathbf{0} \end{bmatrix} \begin{bmatrix} \Delta Y_{1t} \\ \Delta i_t \\ \Delta Y_{1t}^S \\ \Delta i_t^S \end{bmatrix} + \begin{bmatrix} \bar{\mathbf{A}}_{13} & \bar{A}_{14} & \mathbf{0} & \mathbf{0} \\ \mathbf{0} & \mathbf{0} & \mathbf{0} & \mathbf{0} \end{bmatrix} \begin{bmatrix} E_t\left[\Delta Y_{1t+1}\right] \\ E_t\left[\Delta i_{t+1}\right] \\ E_t\left[\Delta Y_{1t+1}^S\right] \\ E_t\left[\Delta i_{t+1}^S\right] \end{bmatrix}$$
$$= \begin{bmatrix} \bar{\mathbf{B}}_1 & \boldsymbol{\alpha}_1 & \mathbf{0} & \mathbf{0} \\ \bar{\mathbf{B}}_2 & \alpha_2 & \bar{\mathbf{B}}_2^{\Delta} & \alpha_{\mathbf{2}}^{\boldsymbol{\Delta}} \end{bmatrix} \begin{bmatrix} Z_{t-1} \\ Z_{t-1}^S \end{bmatrix} + \begin{bmatrix} \bar{V}_{1t} \\ \bar{V}_{2t} \end{bmatrix}, \tag{19}$$

where Z_{t-1} and Z_{t-1}^Scontain two lags of ΔY_{1t} and Δi_t and the cointegrating vectors ci_{t-1} and ci_{t-1}^S. The matrices $\bar{\mathbf{B}}_1, \bar{\mathbf{B}}_2$ and $\bar{\mathbf{B}}_2^{\Delta}$ contain the loading factors $\boldsymbol{\alpha}_1, \alpha_2$ and α_2^{Δ} as in Equation (9). To be consistent with our identification assumption (and explained more fully in Appendix B) the 2×2 matrix of loading factors $\boldsymbol{\alpha}_1^{\Delta}$ is restricted to be zero. Notice that, although we favor a cointegrated representation of the data, our method can equally well be applied to the case in which they are modeled as stationary but highly persistent. Nothing of substance in our identification method hinges on this issue.[9]

[9] If the data are non-stationary and we incorrectly assume a stationary model our infer-

7 Empirical Results

Here we report the results of an empirical exercise on U.S. data. We begin by discussing our choice of sample period. This was governed by our theoretical structure that assumes the existence of a single structural break in policy. We then discuss our parameter estimates and compare the implications of our results for the determinacy properties of the data before and after 1979. Finally, we construct estimates of the structural shocks and present estimates of the impulse response functions implied by our model under two different policy regimes.

7.1 Choosing the Sample Period

Our goal in choosing a sample was to find the longest possible data series that contains a single structural break. Our choice was guided in part by anecdotal evidence and in part by formal and informal statistical tests. Anecdotal evidence comes from previous literature that associates regimes with the Chairmanship of the Fed. To supplement this evidence we chose to split our sample only when it was corroborated by formal statistical evidence for a break. We present the most important results in this Section and we refer the reader to Appendix A for a more detailed description.

Our decision to restrict ourselves to the case of a single break led us to discard the initial part of our data. Informal tests suggested two breaks, one in 1970 and one in 1979. Given this informal evidence we ran formal tests for a single break over the sub-samples 1960 to 1979 and 1970 to 1999. These formal tests confirmed the choice of 1970 as a suitable starting date and 1979 as an appropriate choice of the break date.

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ences will be invalid. If they are stationary, and we incorrectly assume a co-integrated model, the standard errors implied by our error correction representation will still be asymptotically correct. A consequence for identification is that if the data were stationary we would be able to use all level variables in Y_{t-1} and Y_{t-1}^S as instruments. Since we favor an interpretation in which there are two cointegrating vectors we limit ourselves to the stationary linear combinations ci_{t-1}. Moreover, since non-stationary data may take values in the real line we experimented with a version of our model in which we transformed the data by taking logistic and logarithmic transformations. We found all of our results to be robust to these transformations.

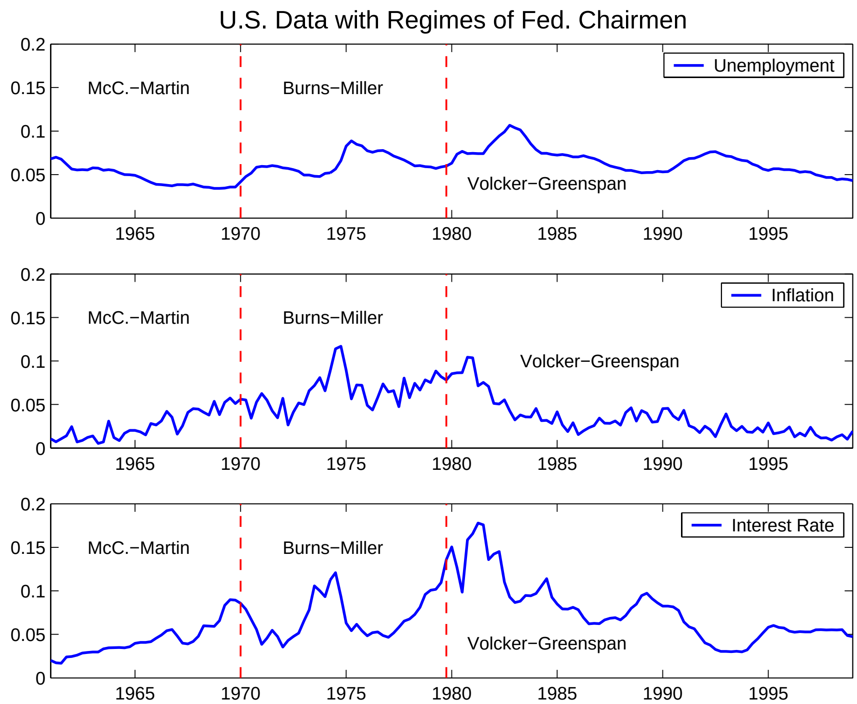

Figure 1

Figure 1 graphs our data, broken into three distinct sub-periods consisting of the dates when the Chairmanship of the Fed was in the hands of Mc. C. Martin, Burns-Miller and Volcker-Greenspan. Since the tenure of Miller was only six months, we combined it with the Burns chairmanship. We combined the Volcker and Greenspan years because in our data analysis we found that this subperiod is consistent with a single stable parameter model.[10]

Based on preliminary data analysis (see Appendix A), we made a provisional choice of 1970Q1-1999Q4 for our sample and conducted a formal test for a break based on the work of Jushuan Bai, Robin Lumsdaine and James Stock (BLS, [2]). We used the framework for integrated data as described in Section 3 pp 402–408 of their paper. This allows for breaks in the short term coefficients and in the cointegrating vectors. The BLS point estimate of the break date coincides with the maximum of the Likelihood function after estimating Equation (8) sequentially for all possible break dates using a DOLS regression as described in Stock and Watson [42].

[10] During the period 1979-1982 the Fed. practiced a monetary targeting approach and the data for this period displays considerable volatility. We experimented by cutting this data from our second sub-sample, but found that our results were not substantially altered.

Figure 2 graphs the recursive likelihood as a function of the break date, k. After applying the BLS procedure to the sample 1970Q1-1999Q4 we found a point estimate for k of 1979Q3 and a 90% confidence interval of 21 quarters on either side. Since this is a relatively wide confidence region we conducted a sensitivity analysis when estimating our model by checking the results for a range of alternative dates and found that our results are robust to the exact choice of break date.[11]

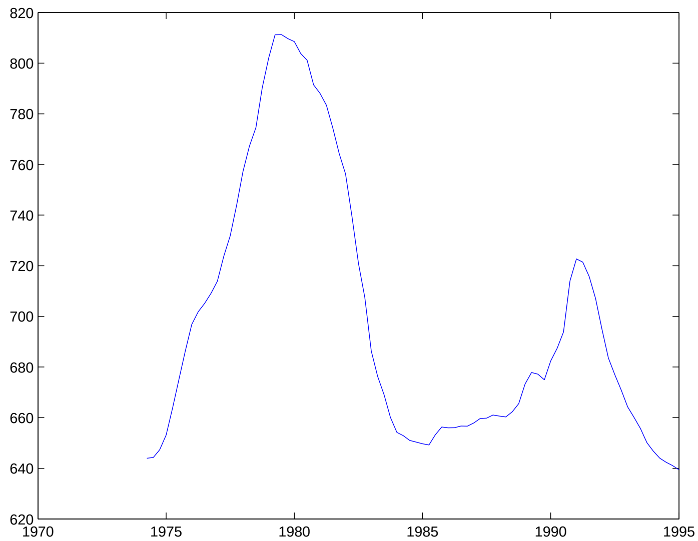

Figure 2: Recursive Likelihood function, trimming region 0.15.

To confirm the starting period for our sample, we conducted the BLS test for the period 1960Q1 to 1978Q1. Our results confirm the finding of our preliminary tests (see Appendix A) with a point estimate of the first break in 1969Q1 and a 90% percent confidence interval of 17 quarters. Although this point estimate is quite imprecise we chose 1970Q1 as our initial starting point because it coincides with the date when when Arthur Burns took over from Mc.Cheney Martin as Chairman of the Board of Governors of the Fed.

7.2 Parameter Estimates

In this section we present the results of our estimates and we discuss their implications for the monetary transmission mechanism.

[11] Robustness results are available from the authors on request. Although there is some indication of a secondary peak in the likelihood function in the early 1990's the evidence that we report in Appendix A supports the view that the data from 1979 through 1999 is well approximated by a single stable parameter model.

Parameter Estimates of the Benchmark Model						
	Unem. Eq.		Infl. Eqn.		Int. Rate. Eqn.	
(Coeff.)	First Regime	Param. Change	First Regime	Param. Change	First Regime	Param. Change
i_t	**−0.11****	–	0.5	–	–	–
	(0.05)		(0.34)			
$E[\Delta u_{t+1}]$	**0.45*****	–	1.12	–	–	–
	(0.17)		(0.84)			
$E[\Delta \pi_{t+1}]$	**0.11***	–	0.36	–	–	–
	(0.05)		(0.28)			
$E[\Delta i_{t+1}]$	**0.12***	–	**−0.81***	–	–	–
	(0.067)		(0.42)			
Δu_{t-1}	0.11***	–	−0.32	–	−0.57	**−1.91****
	(0.14)		(0.0.81)		(0.46)	(0.92)
$\Delta \pi_{t-1}$	−0.01	–	−0.12	–	**−0.23***	−0.09
	(0.03)		(0.22)		(0.14)	(0.23)
Δi_{t-1}	**0.06***	–	−0.28	–	0.14	−0.35**
	(0.03)		(0.27)		(0.18)	(0.24)
Δu_{t-2}	0.13	–	0.29	–	−0.75	**1.59****
	(0.12)		(0.59)		(0.52)	(0.82)
$\Delta \pi_{t-2}$	−0.03	–	−1.59	–	−0.09	0.13
	(0.02)		(0.12)		(0.08)	(0.23)
Δi_{t-2}	−0.029	–	0.10	–	**−0.40****	0.17
	(0.03)		(0.34)		(0.20)	(0.21)
ci_{1t-1}	**−0.038***	–	0.12	–	**−0.37****	0.17
	(0.02)		(0.12)		(0.18)	(0.22)
ci_{2t-1}	–	–	–	–	−0.12	−0.05
					(0.16)	(0.17)
const.	**0.002***	–	−0.00	–	0.01**	−0.00
	(0.001)		(0.00)		(0.00)	(0.01)

Bold numbers indicate statistical significance: a single star indicates 10% level – double star indicates 5% significance – three stars is 1% significance.

J-statistic = 11.2302 Prob[Chi-sq.(12) > J] = 0.5093

Table 1

Tables 1 and 2 compare estimates of the benchmark model, represented by Equation (19), with those of an unrestricted vector error correction model. The instruments for the structural estimation consisted of first and second lags of ΔY_t and ΔY_t^S and four estimated cointegrating vectors $ci_{t-1}, ci_{t-2}, ci_{t-1}^S$ and ci_{t-2}^S. The unemployment and inflation equations each contain four endogenous variables; these are the contemporaneous value of the fed funds rate and the 3×1 vector of date t expectations of Y_{t+1}.

Parameter Estimates of the VECM						
	Unem. Eqn.		Infl. Eqn.		Int. Rate. Eqn.	
(Coeff.)	First Regime	Param. Change	First Regime	Param. Change	First Regime	Param. Change
Δu_{t-1}	**0.32****	0.25	−0.25	−1.11	**−1.18***	−1.16
	(0.19)	(0.24)	(0.76)	(0.90)	(0.19)	(0.78)
$\Delta \pi_{t-1}$	−0.03	0.06	0.28	**−0.66****	−0.23	−0.23
	(0.05)	(0.07)	(0.29)	(0.32)	(0.19)	(0.19)
Δi_{t-1}	−0.11	0.12	0.16	−0.24	0.27	−0.50
	(0.08)	(0.09)	(0.39)	(0.41)	(0.19)	(0.31)
Δu_{t-2}	−0.09	0.25	0.25	0.64	−1.00	1.14
	(0.23)	(0.31)	(0.72)	(0.90)	(0.19)	(0.79)
$\Delta \pi_{t-2}$	**−0.03***	−0.022	0.24	**−0.52****	−0.088	0.24
	(0.02)	(0.06)	(0.19)	(0.23)	(0.19)	(0.16)
Δi_{t-2}	−0.01	0.05	−0.19	0.09	**−0.61****	0.24
	(0.09)	(0.10)	(0.43)	(0.48)	(0.19)	(0.30)
ci_{1t-1}	**−0.17****	0.10	**0.90*****	**−1.008****	**−0.34****	0.10
	(0.09)	(0.09)	(0.28)	(0.30)	(0.19)	(0.20)
ci_{2t-1}	0.03	−0.01	**0.91*****	**−0.87****	−0.11	−0.11
	(0.08)	(0.08)	(0.29)	(0.30)	(0.19)	(0.22)
const.	0.00	−0.00	**−0.04*****	**0.05****	0.01	−0.00
	(0.00)	(0.00)	(0.01)	(0.01)	(0.19)	(0.01)

Bold numbers indicate statistical significance: a single star indicates 10% level – double star indicates 5% significance – three stars is 1% significance.
Model is just identified.

Table 2

We estimated the model by GMM. For the Benchmark model we identified the coefficients on endogenous variables in the private sector structure by excluding twenty step and slope dummies from the unemployment and inflation equations.[12] We identified the policy rule by excluding all contemporaneous and expected future variables.

Since we have twenty exclusion restrictions and eight endogenous coefficients, Hansen's $J - statistic$ has a chi-squared distribution with twelve degrees of freedom. This statistic has a value of 11.23 which has a $p - value$ of 0.51. We conclude that the overidentifying restrictions are not rejected. However, the values of individual coefficients are imprecisely estimated.

The $F - statistic$ for the exclusion of the instruments in the first stage regression of the fed funds rate has a value of 3.14 (all variables in levels) and

[12] These include two lags of ΔY_t^S (twelve exclusion restrictions) the cointegrating vector ci_{1t-1} (two exclusion restrictions) the changes in the cointegrating vectors ci_{1t-1}^S and ci_{2t-1}^S (four exclusion restrictions) and the step dummy (two exclusion restrictions).

3.45 for our error correction specification. These values point strongly to the presence of weak instruments (see Stock, Wright and Yogo [43]). It might be possible to increase the strength of the instruments by considering multiple breaks and we plan to explore this avenue in future research. As a practical consequence, although we are not able to reject the assumption that there exists a stable private sector structure, we are also not able to pin down the value of the private sector parameters with any great precision.

For comparison purposes, Table 2 reports estimates from the VECM, estimated with the same method as the benchmark model and exactly identified by the exclusion restrictions that all three equations contain only lagged right-hand-side variables.

7.3 Determinacy of Equilibrium

There has been active debate in the literature over the determinacy of equilibrium under alternative policy rules. In their seminal work, Clarida-Galí-Gertler [11] (CGG) found evidence that the Burns-Miller regime was characterized by a passive policy rule in which the Fed responded weakly to expected future inflation. When policy is passive the theoretical model admits the existence of multiple sunspot equilibria. CGG interpreted the high variance of inflation and the output gap during this period as support for the view that a passive policy allowed non-fundamental uncertainty to exert an independent influence on economic activity.

Under Volcker-Greenspan, CGG estimated a Fed policy rule that was active with a strong interest rate response to expected inflation. This active policy was associated with lower variance of inflation and unemployment in the post 1979 period. The Clarida-Galí-Gertler results have been subsequently confirmed by a number of studies, most notably by Lubik and Schorfheide [26] who constructed the posterior-odds ratio for determinacy versus indeterminacy in a three equation model.

In contrast to this literature, our point estimates of the structural model are consistent with determinacy in both sample periods. We conducted a monte-carlo simulation in which we took 1000 draws from the estimated variance co-variance matrix of the structural parameters and for each draw we computed the determinacy properties of the equilibrium. For the Burns-Miller regime 53.7%

of the draws were determinate, consistent with the point estimates, whereas under Volcker Greenspan we found that 75.5% of the draws were in the determinacy region. The main reason for the differences of our results from those of Lubik-Schorfheide is in our different identification assumptions. Whereas Lubik-Schorfheide imposed strong priors on the New-Keynesian structure, our identification scheme is much weaker and relies on the assumption of private-sector constancy across a structural break.

7.4 The Characteristics of Structural Shocks

In this subsection we report the characteristics of our estimates of the variance covariance matrices of the structural shocks for each regime.

Since we did not restrict the covariance matrix of the structural residuals we must take a stand on how to interpret non-zero elements of the estimated covariance matrix. Consider first, the possibility that the covariance of inflation and unemployment, $\sigma_{\pi u}$, is estimated to be non-zero. Since we have normalized the quasi-reduced form to exclude contemporaneous effects of inflation and unemployment on each other, we would expect the existence of contemporaneous links in the true structure to show up as a non zero value in our estimate of this parameter. The existence of a stable private sector structure implies that $\sigma_{\pi u}$ should be stable across regimes. Although, in principle, we could have imposed the stability hypothesis as an additional moment restriction in GMM, in practice we allowed our estimate of $\sigma_{\pi u}$ to break. As reported below, our estimates of $\sigma_{\pi u}$ are not significantly different in the two regimes which we take as additional evidence in support of the existence of a stable private sector structure.

Consider next the possibility that structural shocks to unemployment and inflation may be correlated with structural shocks to the fed funds rate. Recall that, to identify policy, we excluded contemporaneous and expected future values of all variables from the policy rule. In contrast, our private sector equations allow for contemporaneous effects of the interest rate on both unemployment and inflation. These identifying assumptions imply that if unemployment and inflation shocks are correlated with the interest rate then we should interpret this correlation as evidence of contemporaneous responses of the fed to private sector innovations. Further, if those covariances change from one period to the next, we take that as evidence of changes in the policy rule across regimes.

In Figure 3 we graph the residuals from our benchmark model alongside residuals for the more conventional VECM identification and in Tables 4 and 5 we present estimates of a Choleski decomposition of the variance covariance matrices of the shocks.

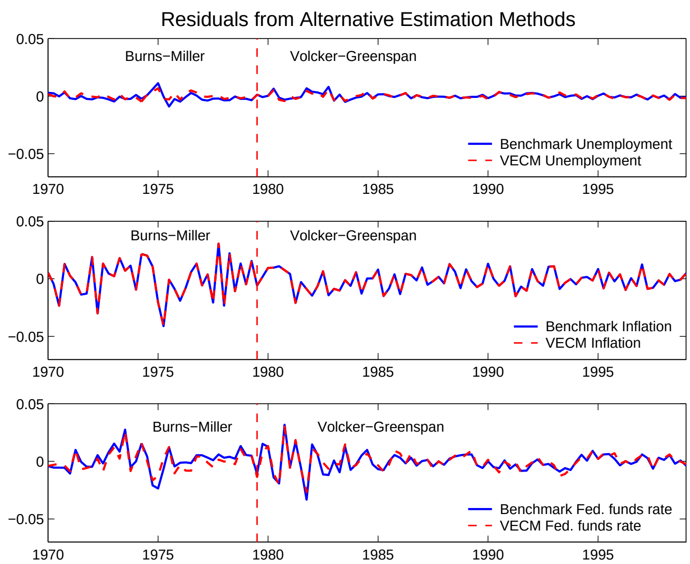

Figure 3

Table 3 reports estimates of $\mathbf{P}_{\bar{V}\bar{V}} \equiv (\mathbf{\Omega}_{\bar{V}\bar{V}})^{1/2}$ for the Benchmark model where

$$\mathbf{\Omega}_{\bar{V}\bar{V}} = \begin{bmatrix} \mathbf{\Omega}_{11} & \Omega_{21}^{T} \\ \Omega_{21} & \omega_{22} \end{bmatrix}.$$

Standard errors of the elements of $\mathbf{P}_{\bar{V}\bar{V}}$ (computed by Monte-Carlo simulation) are reported in parentheses. The method we used to estimate $\mathbf{\Omega}_{\bar{V}\bar{V}}$ is documented in Appendix C and described more completely in Beyer-Farmer [5].

Recall that variation in $\mathbf{\Omega}_{\bar{V}\bar{V}}$ across regimes should be associated solely with changes in the elements of Ω_{21}, (the covariances of fed funds rate shocks with fundamental unemployment and inflation shocks) and ω_{22} the (variance of fed funds rate shocks). Consistent with this assumption, our estimates of $\mathbf{\Omega}_{11}$ are very close across sub-periods.

Estimated Standard Errors of Structural Shocks (Benchmark Model)
(Choleski Decomposition of Estimated VCV Matrix)

	Burns-Miller			Volcker-Greenspan		
	u	π	i	u	π	i
u	3.7			**2.6*****		
	(4.8)			(0.4)		
π	1.4	**9.1***		−4.5	**8.3****	
	(25.4)	(5.7)		(3.3)	(3.0)	
i	**−6.8****	8.6	4.3	−1.5	−2.0	**6.9*****
	(2.4)	(5.5)	(3.6)	(2.1)	(2.4)	(1.3)

Bold numbers indicate statistical significance: One two or three stars indicate values that are one, two or three times the standard error of the relevant coefficient.
All figures are $\times 10^{-3}$.

Table 3

The Choleski decomposition of $\mathbf{\Omega}_{11}$ can be written as follows

$$\mathbf{\Omega}_{11} = \overbrace{\begin{bmatrix} \sigma_{uu} & 0 \\ \sigma_{\pi u} & \sigma_{\pi\pi} \end{bmatrix}}^{\mathbf{P}_{11}} \overbrace{\begin{bmatrix} \sigma_{uu} & \sigma_{\pi u} \\ 0 & \sigma_{\pi\pi} \end{bmatrix}}^{\mathbf{P}_{11}^{T}}.$$

In the first two rows of Table 3 we report the estimates of σ_{uu} , $\sigma_{\pi u}$ and $\sigma_{\pi\pi}$ for the Burns-Miller regime (left panel) and the Volcker-Greenspan regime (right panel). Our point estimates of σ_{uu} are equal to 3.7 and 2.6 (with standard deviations of 4.8 and 0.4) for the two sub-periods and point estimates of $\sigma_{\pi u}$ are 1.4 and -4.5 (with standard deviations of 25.4 and 3.3). The parameter $\sigma_{\pi\pi}$ is estimated at 9.1 (with a standard deviation of 5.7) and 8.3 (with a standard deviation of 3.0). We did not conduct a formal test for equality the two sets of parameters although informally we note that in all cases these estimates are less than two standard deviations apart. We take this as further evidence in favor of our assumption of a constant private sector structure.

To summarize, the parameter constancy assumptions of the benchmark model are broadly consistent with our estimates. Quantitatively, we found that the private sector was hit by an unemployment shock with a standard deviation of roughly 3.0 and an uncorrelated inflation shock approximately three times larger.

The last row of Table 4 reports estimates of $P_{21} = (\Omega_{21})^{1/2}$ and p_{22};

$$\begin{bmatrix} P_{21} & p_{22} \end{bmatrix} = \begin{bmatrix} \sigma_{iu} & \sigma_{i\pi} & \sigma_{ii} \end{bmatrix},$$

defined as the square root of the variance of the fed funds rate, p_{22} and the square root of its covariances with unemployment and inflation, P_{21}. We did not include contemporaneous values of unemployment and inflation in the policy rule. Consequently, if the Fed was responding contemporaneously to private sector unemployment and inflation shocks this contemporaneous response should show up as non-zero elements of P_{21}. If the Fed response changed across regimes, we should see changes in our estimates of these elements as reflected in the last row of Table 3.

Under the Burns-Miller regime the interest rate shock is negatively correlated with the unemployment shock; the parameter σ_{iu} equals -6.8 and is significant; under Volcker-Greenspan this parameter has a point estimate of $= -1.5$ and is statistically insignificant. Athanasios Orphanides and John C. Williams [31] have argued that the main change in policy that occurred in 1979 was a reduction in activist attempts to soften real shocks to unemployment by easing monetary conditions and the evidence we find for changes in our covariance estimates are consistent with this position.

We find no evidence of contemporaneous effects of inflation on unemployment in the policy rule for either regime. Although our point estimate of $\sigma_{i\pi}$ is equal to 8.6 under Burns-Miller and -2.0 under Volcker-Greenspan in both cases, these estimates are statistically insignificant.

For comparison, Table 4 reports the Choleski decomposition of the variance covariance matrix of shocks to a conventional VECM.[13] Here, the identification of shocks is through the causal ordering of the variables and we find strong evidence of changes in all elements of $\mathbf{P}_{\bar{V}\bar{V}}$.

[13] As with the benchmark model this is equivalent to estimation in levels in which we impose the restriction that the model parameter estimates are consistent with the existence of a single common trend and two cointegrating vectors.

Estimated Standard Errors of the Shocks to a VECM
(Choleski Decomposition of Estimated VCV Matrix)

	Burns-Miller			Volcker-Greenspan		
	u	π	i	u	π	i
u	**2.7***** (0.6)			**2.0***** (0.4)		
π	−0.6 (3.1)	**10.9***** (2.6)		**−2.4*** (1.6)	7.5*** (1.1)	
i	**−6.4***** (1.8)	**2.0***** (1.7)	**5.2***** (1.4)	**−3.0**** (1.5)	0.4 (1.3)	**6.4***** (1.7)

Bold numbers indicate statistical significance: One two or three stars indicate values that are one, two or three times the standard error of the relevant coefficient.
All figures are $\times 10^{-3}$.

Table 4

In related work, Sims and Zha [41] estimated a reduced form Markov switching model in which they allowed for multiple breaks in the variance-covariance matrix of the shocks. They were able to account for most of the parameter variation in a structural VAR by switches in the parameters of the variance-covariance matrix across regimes. Although our estimates are not strictly comparable, since Sims-Zha allowed for multiple switches between regimes rather than a single break, our findings are suggestive that the Sims-Zha reduced form estimates might be explainable by a structural model in which reduced form heteroskedasticity is explained by structural breaks in the parameters of the policy rule of the Fed.

7.5 Impulse Response Functions in the Benchmark Model

In this section we sketch the method used to compute impulse response functions to structural shocks (we refer the reader to Beyer-Farmer [5] for the full details). The main difference of our approach from that of a conventional structural VAR is that we add an additional step to compute the reduced form parameters from estimates of the structure.

First we formed estimates of the structural parameters for each regime. Given these estimates we were able to express the reduced form of the model in levels,

$$Y_t = \sum_{j=1}^{k} \hat{\Gamma}_j Y_{t-j} + \hat{\mathbf{\Psi}}_{\bar{V}} \bar{V}_t \tag{20}$$

where the regime dependent matrices $\mathbf{\Gamma}_j$ and $\mathbf{\Psi}_{\bar{V}}$ were found as functions of the estimated structural parameters $\hat{\mathbf{A}}$ and $\hat{\mathbf{B}}$ using a version of Sims [39] algorithm `Gensys`.

The period zero effects of the $i'th$ shock on the $j'th$ equation were found by hitting Equation j of the system by the $j'th$ impact vector multiplied by a shock equal to the standard deviation of the $i'th$ component of $\hat{\mathbf{P}}_{\bar{V}\bar{V}}$. In matrix notation, the $j'th$ impact vector is given by the $j'th$ row of $\hat{\mathbf{\Psi}}_{\bar{V}}\hat{\mathbf{P}}_{\bar{V}\bar{V}}$ and the shock is a vector D_i with zeros in position $j \neq i$ and $[\hat{\mathbf{p}}_{\bar{V}\bar{V}}]_{ii}$ in position i. The $i'th$ vector of impact effects is equal to

$$Y_{0i} = \hat{\mathbf{\Psi}}_{\bar{V}}\hat{\mathbf{P}}_{\bar{V}\bar{V}}D_i.$$

To compute the standard error bounds we took 1000 draws from a normal distribution with mean $\hat{\theta}$ and variance covariance matrix $\Sigma_{\hat{\theta}}$ where $\hat{\theta}$ is the GMM point estimates of the structural parameters and $\Sigma_{\hat{\theta}}$ is the GMM estimate of their VCV matrix. We repeated the computations of the impulse response estimates for every draw and took the $97.5th$ and $2.5th$ percentile of these draws for each date t. We discarded all draws for which the determinacy properties were different from the point estimate, hence, our confidence bounds in each case are conditional on determinacy of the equilibrium. This left 537 good draws in the case of Burns-Miller and 755 in the case of Volcker-Greenspan.

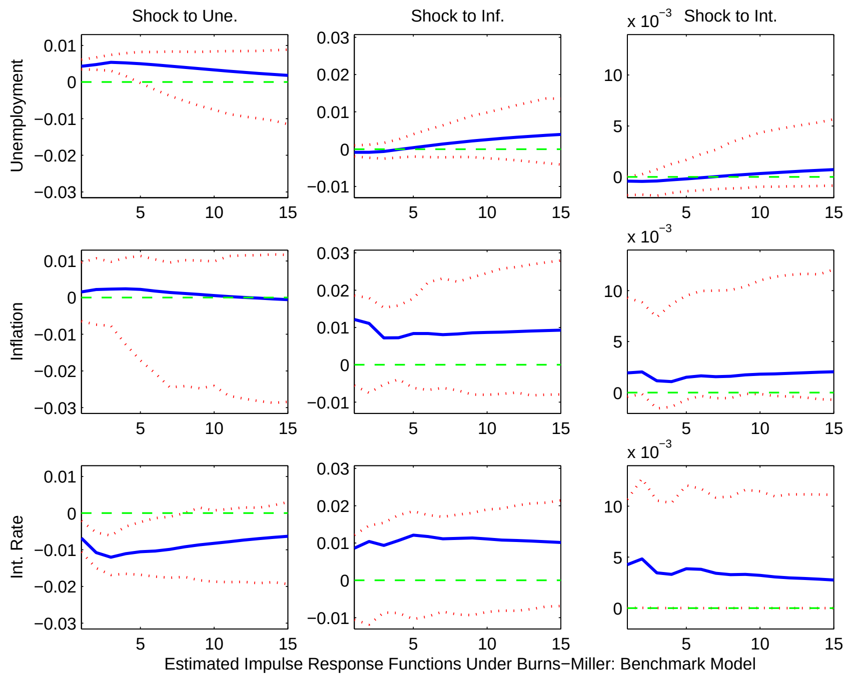

Figure 4

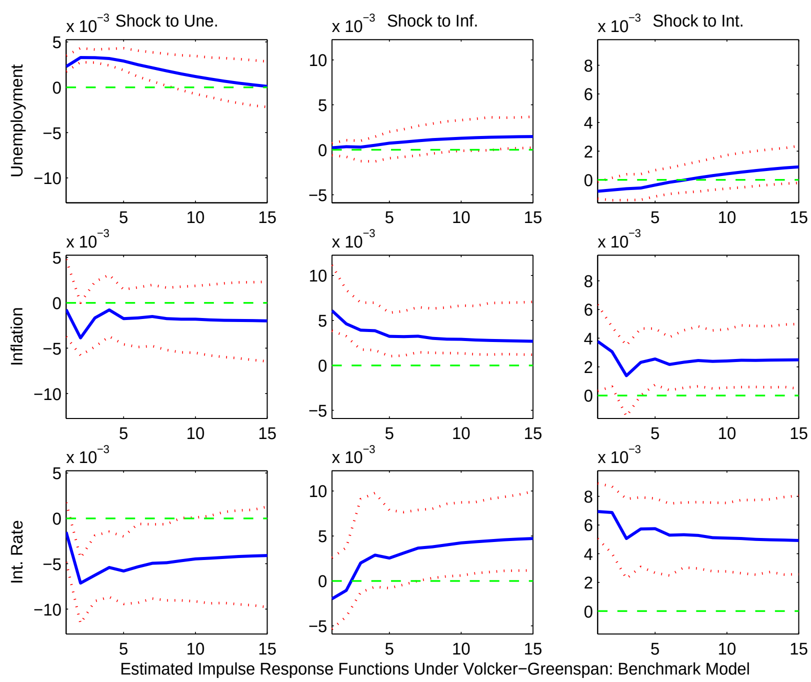

Figure 5

The four graphs in the top left of Figures 4 and 5 illustrate the impulse responses of unemployment and inflation to the two private sector shocks for the two different regimes. Under both regimes we find that the structural shock to unemployment has an impact effect on unemployment of roughly 40 basis points. The effect of the shock builds slightly over five quarters and dissipates over fifteen quarters. This shock has a negligible effect on inflation.

The structural shock to the inflation equation has an impact effect on inflation of approximately 100 basis points under Burns-Miller and 50 basis points under Volcker-Greenspan. It has little or no effect on unemployment on impact, (in either regime) and its effect on unemployment is insignificantly different from zero over a fifteen quarter horizon. The inflation shock, under both regimes, has highly persistent effects on inflation which show little sign of dying out after fifteen quarters. However, the point estimates of the fifteen period effect under Volcker-Greenspan are around 20 basis points as opposed to 90 basis points under Burns-Miller.

The bottom rows of Figures 4 and 5 illustrate our estimates of the response of the fed funds rate to the unemployment and inflation shocks. It is here that we should look for evidence of difference in policy rules across regimes. The bottom left panel of Figure 4 shows that a one standard deviation shock to unemployment was met, under Burns-Miller, by a 100 basis point drop in the nominal interest rate, one quarter later. This effect was highly persistent and the forecast of the interest rate effect of an unemployment shock is still approximately 80 basis points fifteen quarters later. Under Volcker-Greenspan, the effect of an unemployment shock has a similar dynamic path but is roughly half the size with an impact effect of 50 basis points suggesting that the Fed responded less aggressively in the later period to recessions.

The center bottom panel of Figures 4 and 5 illustrates the effects on the interest rate of an inflation shock. Under Burns-Miller our point estimate suggests that a 100 basis point shock to inflation feeds immediately into the fed funds rate and the effect is permanent. Under Volcker-Greenspan our estimates suggest that the impact effect of inflation is negligible but the long run effect is approximately 20 basis points, consistent with the permanent increase in inflation. Both of these graphs are consistent with long-run super-neutrality as the permanent increase in inflation associated with an inflation shock is met by a permanent increase in the fed funds rate of equal magnitude.

The final column of Figures 4 and 5 illustrates the effect of a fed funds rate shock on unemployment and inflation. The bottom graphs in these columns indicate that fed funds rate shocks were highly persistent under both regimes and of comparable size. A typical shock is 50 to 70 basis points on impact, dying out very slowly to around 45 to 55 basis points fifteen quarters later. In both regimes a fed funds rate shock has a negligible effect on unemployment that is insignificantly different from zero over the entire fifteen year horizon.

The inflation effects of a monetary policy shock are significant and positive under Volcker-Greenspan and point estimates are similar in both regimes. A one standard deviation shock to the fed funds rate of approximately 70 basis points is associated with an immediate 40 basis point increase in inflation. The effect on inflation is persistent and roughly constant for the entire fifteen year horizon and is consistent with the long-run interest rate effect which returns from an impact of 70 basis points to a fifteen quarter effect of approximately 50 basis points. It is interesting to note that the fifteen quarter effects on inflation and the fed funds rate are insignificantly different from each other and are consistent with the predictions of an economic model in which money is super-neutral, expectations are rational and prices are perfectly flexible.

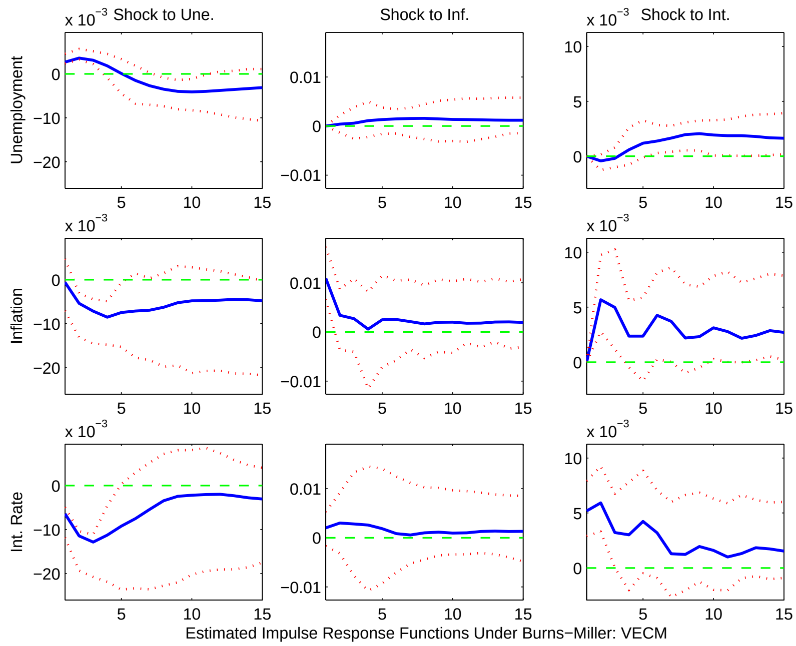

Figure 6

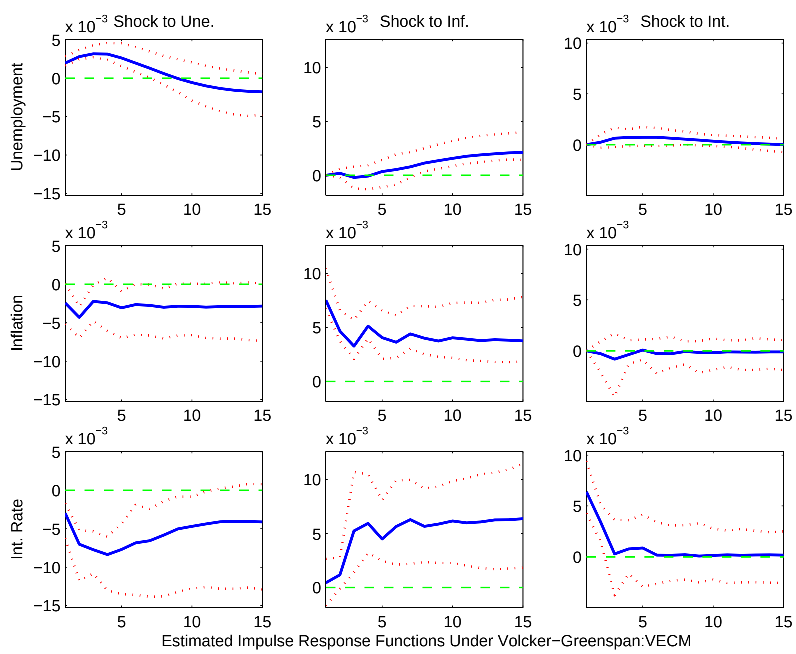

Figure 7

For comparison with the benchmark model, Figures 6 and 7 report impulse response functions for the model in which structural shocks are identified by a conventional causal ordering. This figure imposes the same cointegrating vectors as the benchmark case but, unlike the benchmark model, it allows both cointegrating vectors to enter all three equations and it permits all of the parameters to break across regimes.

8 Conclusion

We have developed a method for identifying the parameters of a block of equations in a linear rational expectations model by exploiting the assumption that the parameters of this block remain invariant to structural change. Although our identification method is related to alternatives that have been proposed in the literature on structural VARs it has, we believe, a distinct advantage over existing methods. By including expectations in the private sector structure our estimated model can be used as a guide to the policy maker. In this sense it is immune to the "Keating critique" (Keating [23]).

We applied our method to the U.S. monetary transmission mechanism and we found evidence that the period from 1970Q1-1999Q4 can be modeled as stable parameter dynamic stochastic general equilibrium model with a single structural break in 1979Q3. This model has a reduced form representation as a VAR. We made the identifying assumption that all parameter change in the parameters of the VAR could be attributed to changes in the structural parameters of the policy rule of the Fed and we found this identifying assumption to be consistent with the overidentifying restrictions implied by the model.

We used our estimates to contruct sample impulse response functions to private sector and policy structural shocks and we estimated the variance-covariance matrix of the structural shocks over two different policy regimes. Our estimates were found to be consistent with the work of Orphanides and Williams [31] who found that policy makers in the Volcker-Greenspan regime were less accommodative to shocks to the real economy than those under Burns-Miller. Our estimates of impluse response functions are broadly consistent with a forward looking rational expectations model that exhibits long-run monetary super-neutrality and price flexibility.

Our method does not come without cost. We imposed weak identifying assumptions and, as a consequence, we were not able to pin down parameters with great precision. It seems likely that our instruments could be strengthened, thereby increasing the precision of our estimates, by extending our analysis to the case of multiple breaks. Additional leverage might be gained by expanding the information set. Our empirical example is based on a data set that contains only three variables. Extending the model by adding additional variables or by including summary information contained in factor variables, as in the work of Beyer et. al. [6], might improve the quality of the estimates.

We also find it likely that one could extend our work to the case of structural Markov switching although this poses considerable technical difficulties because of the nature of the non-linearities introduced and their interaction with forward looking behavior. It is tempting to believe that a policy maker might gain from experimentation by random switches across regimes in order to gain information about the private sector structure. This too poses an interesting extension that we leave for future research.

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Appendix A

This Appendix discusses our evidence for structural breaks and motivates the choice of a restricted sample.[14] In the text, we develop a method that uses a single structural break. To implement this method we must find two adjacent sub-periods, each of which can be described a by a stable parameter model. To operationalize this method our first task was to find a suitable pair of adjacent sub-samples.

Results for the whole sample: 1960-1999

The econometric literature on testing for structural breaks is large, see for example the special issue of *Journal of Econometrics* in 1996. We applied a range of recursive tests, drawn from this literature, for a break within a cointegrated system. We began by applying the Ploberger-Krämer-Kontrus test (PKK, [33]). This is a joint test for constancy of the short run adjustment coefficients in each of the equations of a reduced form VECM. The graphs of the PKK test statistics suggest strongly that there might be breaks close to 1970 and 1979; see Figures (A1–A3) below.

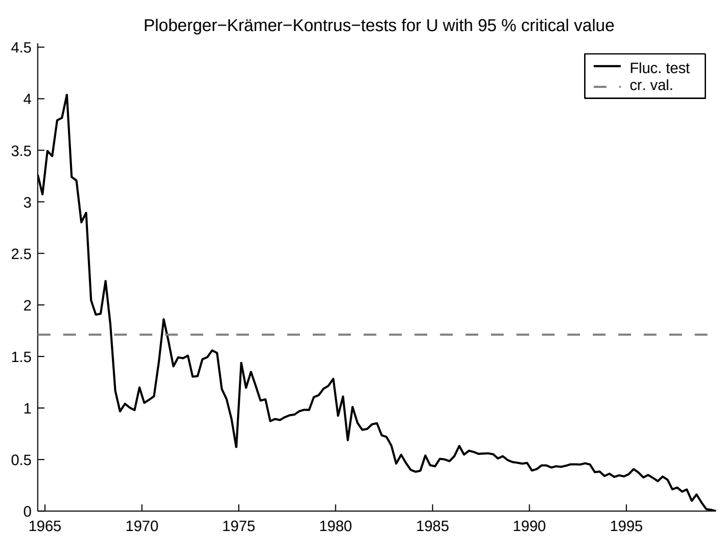

Figure A1: Parameter instability in unemployment equation (1960-2000).

[14] Additional details of test results in Appendix A are available from the authors on request.

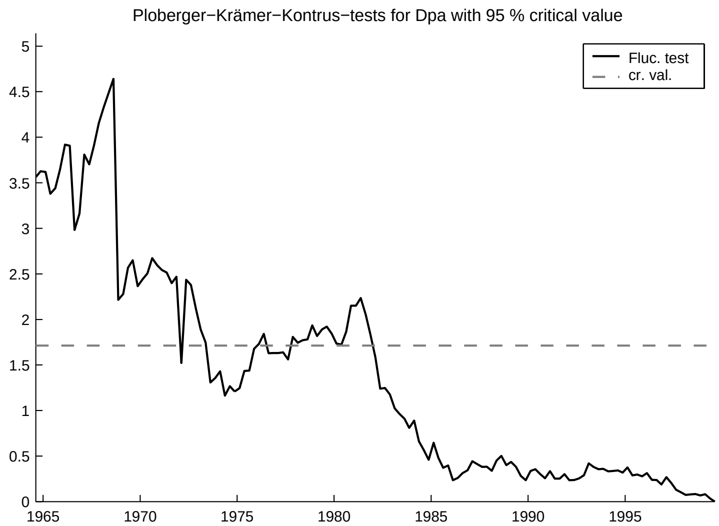

Figure A2: Parameter instability in inflation equation (1960-2000).

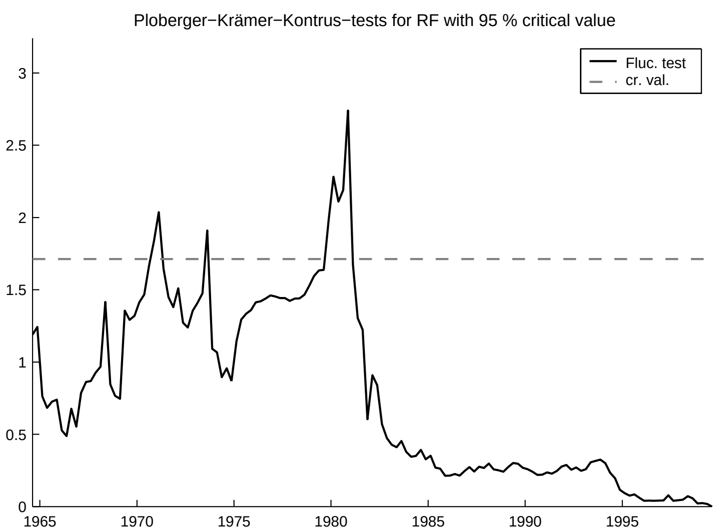

Figure A3: Parameter instability in interest rate equation (1960-2000).

Since PKK is not a test for multiple breaks, the evidence from these graphs is informal. To supplement this evidence we ran recursive Chow tests on the residuals of the VAR and found that these tests also indicated the possibility of breaks around the same dates. This can be seen in Figure A4: panels $a-c$, which graph the residuals and two standard error bounds from reduced form VAR equations for the three variables, u_t, i_t, and π_t. Panels $d-e$ show one-step ahead and panels $g-i$ show recursive n-step ahead forecast errors; horizontal lines are the corresponding critical values, normalized at 5%. Panels k and l

show recursive Chow-tests for parameter constancy in the overall VAR.

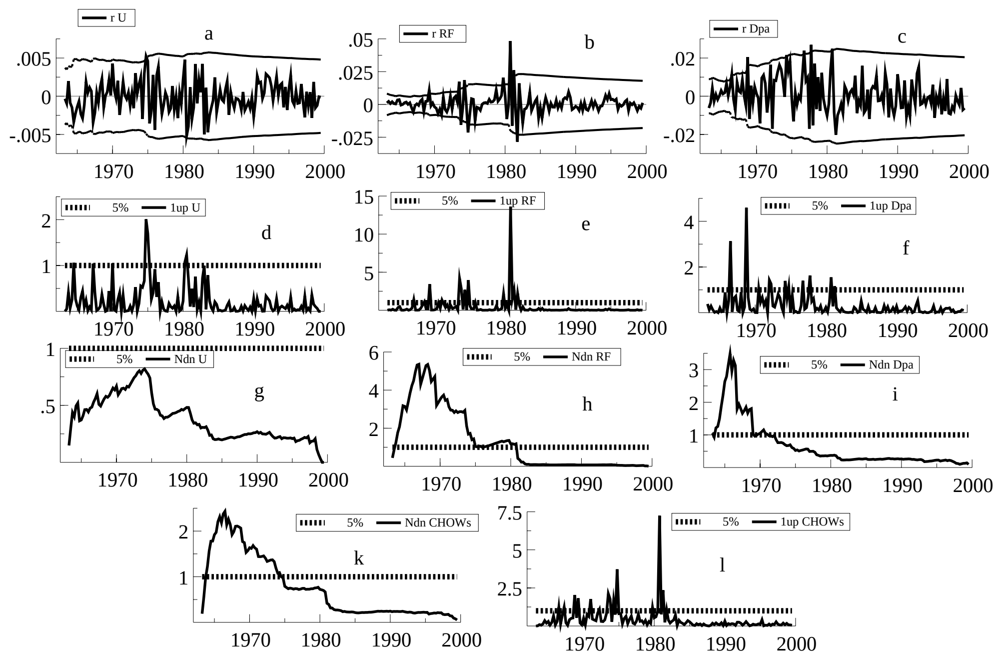

Figure A4: Structural Stability Tests

Next, we ran the Hansen-Johansen [21] fluctuation test which checks for empirical stability of the two non-zero eigenvalues of the cointegration matrix (corresponding to Π in the error correction representation of the model). This test rejects the null hypothesis of parameter stability with $p-values$ below 5% for one and far below 1% for the other eigenvalue. Hansen-Johansen's version of Nyblom's [29] *sup Q* and *Mean Q* test is a test for a break in the estimated cointegrating vectors. In Table A1 we report the critical values for this test, computed with the program *Structural VAR*, by Anders Warne [44].

	60-99		70-99		70-79		79-99	
Test	Q	95% *c.v.*	Q	95% *c.v.*	Q	95% *c.v.*	Q	95% *c.v.*
sup Q	83.6	2.41	4.80	2.45	1.93	2.42	1.77	2.45
Mean Q	3.08	1.10	1.71	1.11	0.37	1.09	0.80	1.10

Table A1: Critical Values for the Hansen and Johansen Nyblom Test

The critical values and test statistics reported in Table A1 show that both tests reject clearly the null of no break over the $60-99$ sample. The test statistic for the *sup Q* test is 83.6 which greatly exceeds the 95% critical value of 2.41. Similarly, the test statistic for the *Mean Q* test is 3.08 which exceeds the 95% critical value of 1.10. Similarly, tests statistics for the $70-99$ period reject the null of no break. In contrast, the test statistics for the two sub-periods $70-79$ and $79-99$ fail to reject the null of a constant parameter model at the 95% level.[15]

To summarize, the informal statistical evidence suggests that the period 1970-1999 can be modeled as two separate stable parameter models with a break around 1979. We gained confidence that the break around 1979 might be attributable to changes in the monetary policy rule from the fact that the stable sample periods coincide closely with dates at which there was a change in the Chairmanship of the Fed. Supporting evidence for 1979 as a good candidate for the second break point is provided by Athanasios Orphanides [30] who analyzes Fed minutes and argues that Fed operating policy changed dramatically when Paul Volcker came into office.

Table A2 reports the dates of changes in the Chairmanship of the Board of Governors of the Federal Reserve System beginning in 1951. There were four such changes during our sample period, beginning in 1960 when data became available on the Federal Funds Rate.

Federal Reserve Chairmen

Wm. McC. Martin, Jr	Apr. 2, 1951-Jan. 31, 1970
Arthur F. Burns	Feb. 1, 1970-Jan. 31, 1978
G. William Miller	Mar. 8, 1978-Aug. 6, 1979
Paul A. Volcker	Aug. 6, 1979-Aug. 11, 1987
Alan Greenspan	Aug. 11, 1987–

Source: www.federalreserve.gov/bios/boardmembership.htm

Table A2: Chairmanship of the Fed

Given our informal evidence for breaks in 1970 and 1979, we built subsamples from 1960Q1-1978Q4 and from 1970Q1 through 1999Q4 and we tested formally for a single break within each subsample. Our prior was that there is a break in each sub-sample, one around 1970 and one around 1980.

[15] Additional tests which we did not apply in this paper are those by Quintos and Phillips [34] and Seo [38].

Results for the sub-sample 1970-1999

For this sub-sample our main results come from the Bai-Lumsdaine Stock test reported in Section 7 of the paper. In addition, we found that the Hansen-Johansen tests again strongly rejects eigenvalue stability far below 1% and Hansen and Johansen's Nyblom test is well above its 99% critical value, (see Table A1). The PKK test strongly rejects the null of parameter stability with $p-values$ again below 1% for each equation and finds a break between 1980Q4 (interest rate equation) and 1981Q2 (inflation equation).

As a further check, we carried out a modified version of the test for structural breaks suggested by Allan Gregory and Bruce Hansen (GH, [19]). We used the version of their test which considers possible breaks in the cointegrating parameters. The null hypothesis is no cointegration against cointegration under a structural break and the null is rejected when the corresponding test statistics (their Z_t^*.and Z_α^* reported in Table 1, p.109) exceed the critical values; our model corresponds to the case where $m = 1$ with specification C/S.

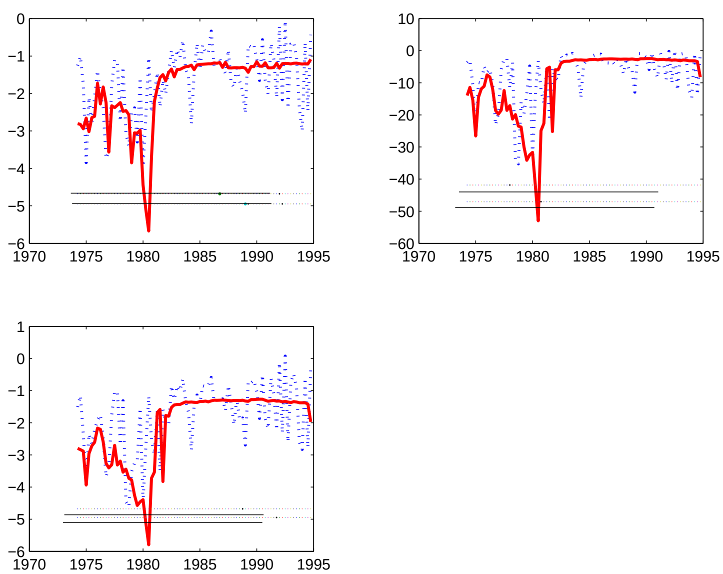

Figure A5: Modified Gregory-Hansen tests

The GH test is based on the residuals of just one cointegrating relationship. To implement it we used a two step strategy. First we simultaneously estimated the two cointegrating vectors using Johansen's [22] method. Second, we applied GH separately to the two series of the corresponding residuals. The results support our earlier finding of a break in one of the two cointegrating vectors.

Figure A5 shows a break in one cointegrating vector in 1980Q4 with a Z_t^* test statistic of -5.7968 and a Z_α^* statistic of -52.99.[16]

Results for the sub-sample 1960-1979

Again for this period we report the BLS results in Section 7 of the paper. In addition, we found that the Hansen-Johansen tests strongly rejects eigenvalue stability far below 1%. Hansen and Johansen 's Nyblom test is again above its 99% critical values (Table A1). The Ploberger-Krämer-Kontrus test strongly rejects with $p - values$ below 1% for each equation and finds a break between 1968Q3 (interest rate equation) and 1969Q1 (unemployment equation). Figure A6 presents the inflation rate results for this sub-sample; additional figures are availa

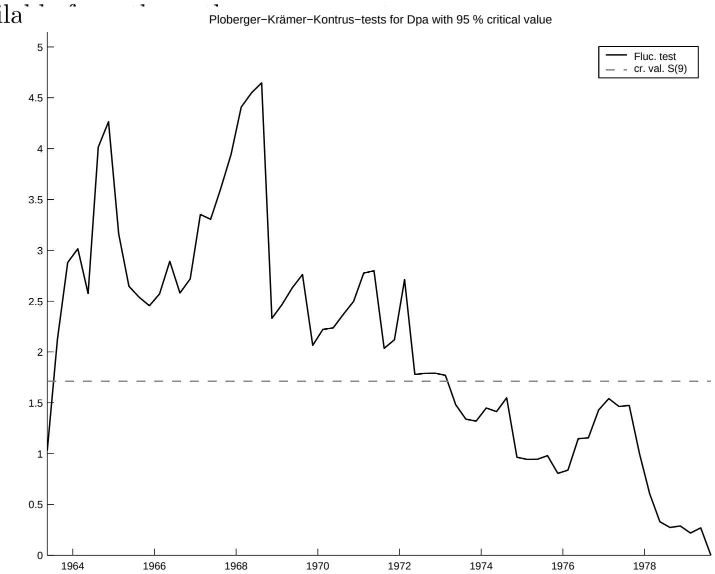

Figure A6: Parameter Instability in Inflation Equation (1960–1978).

[16]The test statistic for the second cointegrating vector which we found to be stable over the whole modelling sample should ideally be outside the critical value bound to be an $I(0)$ variable. This was not the case for our sample; however, the test formaly applies to the case of a single cointegratimg vector. Beyer-Haug [7] have begun to explore the properties of systems in which there are breaks in more than one cointegrating vector but their results are preliminary at this time.

Appendix B

In this Appendix we discuss briefly our estimation and identification of the cointegrating vectors (for a detailed description and statistical analysis, see Beyer and Haug [7]). Since the cointegrating vectors of the reduced form are functions of those of the structural model the assumption that the two equations of the private-sector policy block are invariant across regimes implies that at least one of the cointegrating vectors must also remain unchanged. In our empirical work we used this assumption to identify one of the two cointegrating vectors.

Our estimates of the cointegrating vectors in Definition (9) are presented below

$$\underset{2\times 6}{ci_{t-1}} = \begin{pmatrix} \underset{2\times 3}{\beta} & \underset{2\times 3}{\beta^{\Delta}} \end{pmatrix} \begin{pmatrix} \underset{3\times 1}{Y_{t-1}} \\ \underset{3\times 1}{Y^S_{t-1}} \end{pmatrix} = \begin{bmatrix} 1 & \underset{(0.05)}{-0.643} & \underset{(0.01)}{0.029} & 0 & 0 & 0 \\ 0 & \underset{(0.1)}{-0.659} & 1 & 0 & \underset{(0.2)}{-1.289} & 0 \end{bmatrix}.$$

To account for the shifts in the cointegrating vectors we also estimated ci^S_{t-1} :

$$\underset{2\times 6}{ci^S_{t-1}} = \begin{pmatrix} \underset{2\times 3}{\beta} & \underset{2\times 3}{\beta^{\Delta}} \end{pmatrix} \begin{pmatrix} \underset{3\times 1}{Y_{t-1}} \\ \underset{3\times 1}{Y^S_{t-1}} \end{pmatrix} = \begin{bmatrix} 0 & 0 & 0 & 1 & -0.643 & 0.029 \\ 0 & 0 & 0 & 0 & \underset{(0.2)}{-1.948} & 1 \end{bmatrix}$$

where βY_{t-1} entries are trivially equal to zero. Note that the first cointegrating vector is constant by assumption, and the second one has changed its inflation coefficient from -0.65 to -1.948. Beyer and Farmer [4] attribute this change in the inflation coefficient to a shift in monetary policy.

The loading factors have the following representation:

$$\begin{pmatrix} \underset{\mathbf{3\times 2}}{\boldsymbol{\alpha}} & \underset{\mathbf{3\times 2}}{\boldsymbol{\alpha}^{\boldsymbol{\Delta}}} \end{pmatrix} \begin{bmatrix} \underset{2\times 6}{ci_{t-1}} \\ \underset{2\times 6}{ci^S_{t-1}} \end{bmatrix}.$$

It is straightforward to show that in a structural error correction model as in Equation (8) an unstable cointegrating relationship cannot enter the quasi-reduced form private sector block. This means that the matrices of loading factors $\boldsymbol{\alpha}$ and $\boldsymbol{\alpha}^{\Delta}$ (that appear in the matrix $\mathbf{B}$ in Definition (10)) contain four zero restrictions and have the representation

$$\underset{3\times 2}{\boldsymbol{\alpha}} = \begin{bmatrix} \underset{2\times 1}{\alpha_{11}} & \underset{2\times 1}{0} \\ \underset{1\times 1}{\alpha_{21}} & \underset{1\times 1}{\alpha_{22}} \end{bmatrix} \quad , \quad \underset{\mathbf{3}\times\mathbf{2}}{\boldsymbol{\alpha}^{\boldsymbol{\Delta}}} = \begin{bmatrix} \underset{2\times 1}{\alpha_{11}^{\Delta}} & \underset{2\times 1}{0} \\ \underset{1\times 1}{\alpha_{21}^{\Delta}} & \underset{1\times 1}{\alpha_{22}^{\Delta}} \end{bmatrix}$$

Appendix C

In this appendix we explain how we formed an estimate $\hat{\mathbf{\Omega}}_{\bar{V}\bar{V}}$ of $\mathbf{\Omega}_{\bar{V}\bar{V}}$. Our point estimates of θ were in the determinate regime in both sub-samples (for a discussion of how to handle the indeterminate case the reader is referred to Beyer-Farmer [5]). Given $\hat{\theta}$ for each regime we used a version of Sims [40] code `Gensys` to construct the residuals of the system written as a VAR in levels

$$\hat{\epsilon}_t = Y_t - \bar{C} - \sum_{j=1}^{3} \Gamma_j \left(\hat{\theta}\right) Y_{t-j}. \tag{C1}$$

We estimated a different version of Equation (C1) for each regime. From Equation (C1) we constructed regime dependent estimates of $\mathbf{\Sigma}_{\bar{V}\bar{V}}$

$$\hat{\mathbf{\Sigma}}_{\bar{V}\bar{V}} = \sum_{t=1}^{T} \frac{\left(\hat{\epsilon}_t \hat{\epsilon}_t^T\right)}{T}. \tag{C2}$$

The reduced form errors ϵ_t are related to the structural errors $\bar{V}_t$ by the equation

$$\epsilon_t = \Psi_{\bar{V}}\left(\theta\right) \bar{V}_t. \tag{C3}$$

Taking expectations of the outer product of (C3) leads a system of n^2 equations in n^2 unknowns

$$\mathbf{\Sigma}_{\bar{V}\bar{V}} = \Psi_{\bar{V}}\left(\theta\right) \mathbf{\Omega}_{\bar{V}\bar{V}} \Psi_{\bar{V}}^T\left(\theta\right), \tag{C4}$$

or, exploiting symmetry, this gives $n \times (n+1)/2$ linear equations

$$vech\left(\mathbf{\Omega}_{\bar{V}\bar{V}}\right) = \mathbf{B}\left(\theta\right) vech\left(\mathbf{\Sigma}_{\bar{V}\bar{V}}\right), \tag{C5}$$

where *vech* is the operator that stacks the lower triangular elements of a symmetric matrix into a column vector and the elements of $\mathbf{B}(\theta)$ are functions of $\Psi_{\bar{V}}(\theta)$. To form our estimate $\hat{\mathbf{\Omega}}_{\bar{V}\bar{V}}$, we replaced $\mathbf{B}(\theta)$ and $\mathbf{\Sigma}_{\bar{V}\bar{V}}$ by their estimates from GMM and solved Equation (C5).