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Cross-Country Estimates of the Degree of Fiscal Dominance and Central Bank Independence

by Carlos de Resende

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Abstract

This paper studies the interdependence between fiscal and monetary policies, and their joint role in the determination of the price level. The government is characterized by a long-run fiscal policy rule whereby a given fraction of the outstanding debt, say δ, is backed by the present discounted value of current and future primary surpluses. The remaining debt is backed by seigniorage revenue. The parameter δ characterizes the interdependence between fiscal and monetary authorities. It is shown that in a standard monetary economy, this policy rule implies that the price level depends not only on the money stock, but also on the proportion of debt that is backed with money. Empirical estimates of δ are obtained for OECD and developing countries using data on nominal consumption, monetary base, and debt. Results indicate differences in the degree of fiscal dominance between developed and developing economies. Estimates of δ correlate positively with some institutional measures of *de facto* central bank independence.

JEL classification: E31, E42, E50, E63
Bank classification: Central bank research; Fiscal policy; Inflation: costs and benefits

Résumé

L'auteur étudie l'interdépendance des politiques budgétaire et monétaire ainsi que leur rôle combiné dans la détermination du niveau des prix. L'État est caractérisé par une règle de politique budgétaire à long terme selon laquelle une fraction du service de la dette – δ – est garantie par la valeur actualisée des excédents primaires présents et futurs, et la fraction restante – 1-δ – par les revenus de seigneuriage. Le paramètre δ mesure l'interdépendance des autorités budgétaire et monétaire. L'auteur montre que, dans une économie monétaire, la règle en question implique que le niveau des prix est fonction non seulement du stock de monnaie mais aussi de la part des emprunts garantie par la création de monnaie. À l'aide des données sur la consommation nominale, la base monétaire et la dette, l'auteur estime empiriquement le paramètre δ pour les membres de l'OCDE et les pays en voie de développement. D'après ses résultats, les deux groupes d'économies se différencient par leur degré de prépondérance budgétaire. Les estimations de δ sont corrélées positivement avec certaines mesures institutionnelles de l'indépendance *effective* des banques centrales.

Classification JEL : E31, E42, E50, E63
Classification de la Banque : Recherches menées par les banques centrales; Politique budgétaire; Inflation : coûts et avantages

1 Introduction

This paper studies the interdependence between fiscal and monetary policies, and their joint role in the determination of the aggregate price level. In general, fiscal and monetary policies are linked through the consolidated government budget constraint. A combination of taxes, new debt issue, and seigniorage revenue must finance government expenditures in every period. In terms of the intertemporal budget constraint, outstanding debt must be backed by a combination of the present discounted value of current and future primary surpluses and seigniorage revenues. More specifically, this paper investigates if the proportion of debt that is backed by each source of revenue, primary surplus or seigniorage, matters for the determination of the price level.

The theoretical analysis is carried out in a standard competitive monetary economy. The government is characterized by a long-run fiscal policy rule whereby a given fraction of the outstanding debt, say δ, is backed by the present discounted value of current and future primary surpluses. The remaining debt is backed by seigniorage revenue. The parameter δ is structural and summarizes the degree of interdependence between fiscal and monetary authorities in a given institutional setup. It is shown that in a standard monetary economy, this policy rule implies that the price level depends not only on the money stock, but also on the proportion of debt that is backed with money.

This paper draws on earlier research by Aiyagari and Gertler (1985), extending their work in at least three directions. First, results are derived using only the long-run fiscal policy rule without having to specify a particular period-by-period rule. This long-run rule is compatible with the time-stationary rule in Aiyagari and Gertler, but also with other (perhaps not time-stationary) period-by-period rules. Second, the determination of the price level is characterized at all times, rather than only at the steady state. Finally, a simple empirical strategy is proposed to construct estimates of the δ parameter for a cross-country sample of developing and industrialized economies.

In order to understand the importance of the empirical analysis, note that in this model there is a continuum of fiscal regimes indexed by δ. There are two polar cases. First, in the case where $\delta = 1$, the fiscal authority backs fully all government debt. Fiscal policy accommodates monetary policy in the following sense: whenever the monetary authority sells government bonds in the open market, the fiscal authority increases current or future taxes, and/or reduces current or future expenditures, to back the principal and interest payments on the newly issued debt. The monetary authority never responds to the increase in the stock of government debt associated with a budget deficit. Sargent (1982) and Aiyagari and Gertler (1985) refer to this case as a Ricardian regime. In this paper, it will be referred to as one of zero fiscal dominance or central bank independence.

Second, in the case where $\delta = 0$, the monetary authority backs fully all government debt. In particular, the monetary authority accommodates the fiscal authority whenever a budget deficit is financed with debt. This accommodation takes the form of an increase in current or future seigniorage revenues to back the principal and interest payments on the newly issued debt. The fiscal authority is insensitive to monetary policy in that neither taxes nor expenditure react (today or in the future) to changes in stock of outstanding government debt. Sargent, and Aiyagari and Gertler refer to this case as a polar Non-Ricardian regime. In this paper, it will be referred to as one of complete fiscal dominance.

Aiyagari and Gertler correctly argue that one cannot distinguish between Ricardian and Non-Ricardian regimes on the basis of long-run correlations between nominal interest rates and money growth. The reason is that there exist monetary policy rules for which the Non-Ricardian regimes ($0 \leq \delta < 1$) generate the same correlation as the Ricardian regime ($\delta = 1$). However, we show that under certain conditions, the dynamics of money, debt, and private consumption allow the direct estimation of δ and standard statistical inference can be used to draw conclusions regarding the regime that better describes policy in a given economy. The estimation strategy is based on now standard results in unit-root econometrics that were not well developed at the time Aiyagari and Gertler wrote their contribution.

Using data from a sample of developed and developing economies, country-specific estimates of δ are constructed . The estimates reveal important cross-country heterogeneity. For instance, the null hypothesis that δ equals 1 cannot be rejected at standard levels for most industrial (OECD) countries in the sample, but is more frequently rejected among developing countries. In addition, only within the subsample of developing countries can we find examples for which the null hypothesis that δ equals 0 cannot be rejected. This findings suggest that fiscal dominance is more common among developing countries, while central bank independence seems to be the case for most OECD countries, implying that, for OECD countries: (i) the fiscal authority backs most, if not all, outstanding debt, and (ii) debt plays only a minor role in the determination of the price level. This conclusion is less straightforward for developing economies.

Additional empirical implications of the model are also examined. First, estimates of δ are compared with measures of central bank independence proposed in the literature. Results indicate a positive and significant correlation between δ and the legal autonomy index proposed by Grilli, Masciandaro and Tabellini (1991) and a negative (also significant) correlation, as expected, between δ and a central bank independence index based on the turnover rate of governors proposed by Cuckierman, Webb and Neyapti (1992).

In Sargent and Wallace (1981), the interaction between fiscal and monetary authorities takes the form of a coordination game. The central bank could move first, determine how much seigniorage revenue can be raised, and force the fiscal authority to follow a policy that satisfies the government's consolidated intertemporal budget constraint. Then, a central bank that is committed to price stability could indeed deliver price stability regardless of fiscal policy. Alternatively, the fiscal authority could move first by defining the path of the primary surplus. Since higher seigniorage revenues would be necessary to avoid explosive debt paths, fiscal policy would have an effect on the price level. Given a predetermined path for the primary surplus, "tight" money today triggers higher interest rates, increases interests rate payments on the government's debt, and requires "loose" money later. Rational agents anticipate the future increase in money creation and bid the price level up today. This is Sargent and Wallace's *unpleasant monetarist arithmetic*. The results in this paper imply that, for most industrialized countries in the sample, the central bank is the first mover, but this result is less clear for developing economies, where fiscal dominance is more common. That is, in OECD countries, it seems to be the monetary authority that sets its policy in advance and imposes discipline on the fiscal authority.

This work is related to, but conceptually different from, the literature on the Fiscal Theory of the Price Level (FTPL) [see, for example, Woodford (1995) and Cochrane (1998, 2001)]. Under the FTPL, the price level is determined by the intertemporal budget constraint as the quotient between the nominal value of the interest bearing debt and the present value of the surplus, that might include seigniorage revenues. The underlying assumption is that the government's actions are not constrained by budgetary issues. Consequently, the intertemporal budget constraint holds as an equilibrium condition, rather than as a constraint, and only for equilibrium prices. Any change in fiscal policy must impact the price level, regardless of how committed the monetary authority is to price stability. Both the model presented in this paper and the FTPL predict a relationship between the price level and fiscal variables. However, in this paper it is assumed that the intertemporal budget constraint is always satisfied for any arbitrary sequence of prices, whereas the FTPL assumes it is an equilibrium condition. This difference means that the econometric results presented here should not be interpreted as a formal test of the FTPL.

The paper is organized as follows. Section 2 presents the theoretical model. Section 3 outlines the estimation strategy and reports the empirical results. Section 4 concludes.

2 The Model

2.1 Private Sector

The economy is populated by identical, infinitely-lived consumers with perfect foresight.[1] The objective of the representative consumer is:

$$\max_{\{c_t, n_t, m_t, b_t, k_t\}} \sum_{t=0}^{\infty} \beta^t u\left(c_t, m_t/p_t, 1-n_t\right), \tag{1}$$

[1] The assumption of perfect foresight is not crucial for the theoretical results, but it is analytically convenient. Aiyagari and Gertler (1985) allow uncertainty but focus on a steady state with constant asset prices. Leeper (1991) permits shocks to the fiscal and monetary policy rules, but output, consumption, and government expenditure are deterministic.

where $\beta \in (0,1)$ is the subjective discount factor and u is strictly increasing in all arguments, strictly concave, twice continuously differentiable, and satisfies the Inada conditions.

In each period, consumers choose consumption (c_t), labor (n_t), and next-period holdings of capital (k_t), money (m_t) and nominal one-period government debt (b_t). The variable p_t is the aggregate price level. The time endowment is normalized to one. The population size is constant and normalized to one. Capital and labor services are rented each period to a representative competitive firm that produces output according to a standard neoclassical production function.

The inclusion of real balances (m_t/p_t) as an argument of the utility function reflects the convenience of using money in carrying out transactions. Feenstra (1986) shows the equivalence between including real balances in the utility function, assuming liquidity costs that appear in the budget constraint, and introducing a cash-in-advance constraint. In this sense, the approach followed here to motivate money demand is not restrictive. Since the model is concerned with the composition of government liabilities, following Woodford (1995), m_t is interpreted as the consumer's holdings of the monetary base.

A logarithmic and separable instantaneous utility function is assumed because it is analytically very tractable and allows us to exploit the linearity of the government's budget constraint:[2]

$$u\left(c_t, m_t/p_t, 1-n_t\right) = \ln(c_t) + \gamma \ln(m_t/p_t) + \theta \ln(1-n_t),$$

where γ and θ are positive constants that measure the relative importance of real money holdings and leisure in utility.

The consumer's optimization problem is subject to a no-Ponzi-game condition and to the sequence of budget constraints (expressed in real terms):

$$c_t + \frac{m_t}{p_t} + \frac{b_t}{p_t} + k_t = w_t n_t + r_t k_{t-1} + \frac{m_{t-1}}{\pi_t p_{t-1}} + i_{t-1}\frac{b_{t-1}}{\pi_t p_{t-1}} - \tau_t, \tag{2}$$

for all t, where τ_t is a lump-sum tax, $\pi_t = p_t/p_{t-1}$ is the gross inflation rate, i_{t-1} is the gross nominal interest rate on government debt which is set in period $t-1$ and paid in period t, w_t is the wage rate, and r_t is the gross return on capital between periods $t-1$ and t. In equilibrium, the absence of arbitrage profits will require r_t to equal the real gross interest rate i_{t-1}/π_t.

First-order necessary conditions for the representative consumer's problem include:

$$\begin{aligned} 1/c_t &= \beta(i_t/\pi_{t+1})(1/c_{t+1}), \qquad (3) \\ m_t/p_t &= \gamma c_t i_t/(i_t - 1), \qquad (4) \end{aligned}$$

Equation (3) is an Euler equation for consumption and equation (4) defines money demand as a function of consumption and the return on money. We will see below that only these two conditions

[2] All results of the paper follow through if agents derive utility from government expenditures, as long as they enter separably in the utility function.

are necessary to derive the model's implications for the aggregate price level, without reference to the remaining first-order conditions.

2.2 Government

In every period, the government spends an exogenous amount of resources G_t. Government expenditures may be financed by levying lump-sum taxes (τ_t), by issuing money (M_t), and by increasing public debt (B_t). The government is subject to a no-Ponzi-game condition and to a dynamic budget constraint (expressed in real terms):

$$G_t + (i_{t-1} - 1)\frac{B_{t-1}}{p_t} = \tau_t + \frac{(M_t - M_{t-1})}{p_t} + \frac{(B_t - B_{t-1})}{p_t}. \tag{5}$$

Forward iteration on (5) and the government's no-Ponzi condition imply an intertemporal budget constraint:

$$\begin{aligned} i_{t-1}\frac{B_{t-1}}{p_t} &= \sum_{j=0}^{\infty}\frac{\tau_{t+j}}{R_t^{(j)}} + \sum_{j=0}^{\infty}\frac{M_{t+j} - M_{t+j-1}}{p_{t+j}R_t^{(j)}} - \sum_{j=0}^{\infty}\frac{G_{t+j}}{R_t^{(j)}}, \\ &= \mathcal{T}_t + \mathcal{S}_t - \mathcal{G}_t, \end{aligned}$$

where $R_t^{(j)} = \prod_{h=1}^{j} r_{t+h}$ is the j-periods-ahead market discount factor, and $\mathcal{T}_t$, $\mathcal{S}_t$ and $\mathcal{G}_t$ are the present value of tax receipts, seigniorage revenue, and government expenditure, respectively. Without loss of generality, we assume that the government's present value budget constraint holds with equality.[3]

The government is assumed to follow a "long-run" fiscal policy rule whereby it commits itself to raise large enough primary surpluses (in present value terms) to back a constant fraction of the currently outstanding debt. More formally:

Definition (The δ-backing Fiscal Policy): *Given a sequence of prices $\{i_{t+j-1}, p_{t+j}\}_{j=0}^{\infty}$ and an initial stock of nominal debt B_{t-1}, a δ-backing fiscal policy is a sequence $\{G_{t+j}, \tau_{t+j}, B_{t+j}\}_{j=0}^{\infty}$ such that, for all t:*

$$\mathcal{T}_t - \mathcal{G}_t = \delta i_{t-1}\frac{B_{t-1}}{p_t}, \tag{6}$$

where $\delta \in [0, 1]$.

Put simply, this fiscal policy rule means that a constant fraction (δ) of the outstanding government debt, including interest payments, is backed by the present discounted value of current and future primary surpluses. Since the government's intertemporal budget constraint is always satisfied, it follows that:

$$\mathcal{S}_t = (1 - \delta)i_{t-1}\frac{B_{t-1}}{p_t}. \tag{7}$$

[3] Note that we impose a no-Ponzi game condition on total government liabilities. Under the assumption that the government does not waste revenues, this amounts to

$$\lim_{j\to\infty} \left(M_{t+j} + B_{t+j}\right)/p_{t+j}R_t^{(j)} = 0.$$

Hence, the policy (6) also implies that a fraction $(1-\delta)$ of the currently outstanding debt is backed by the present discounted value of current and future seigniorage revenue.

The set of possible fiscal regimes is indexed by the fraction δ of the outstanding debt that is backed by the primary surplus. Because $\delta \in [0,1]$, this set is a continuum limited by the following two polar cases:

(i) In the case where $\delta = 1$, the fiscal authority backs fully all outstanding debt. It commits itself to adjust the stream of future primary surpluses in order to match the current value of the government's bond obligations. There is complete accommodation of the fiscal policy to any open market sale by the monetary authority. Whenever the monetary authority sells government bonds in the open market, the fiscal authority increases current or future taxes (and/or reduces current or future expenditures) to back the principal and interest payments on the newly issued debt. On the other hand, the monetary authority never responds to the increase in the stock of government debt associated with a budget deficit. Sargent (1982) and Aiyagari and Gertler (1985) refer to this case as a Ricardian regime, while Leeper (1991) refers to it as one of active monetary/passive fiscal policy. Here it will be called one of zero fiscal dominance and complete central bank independence.

(ii) In the case where $\delta = 0$, all outstanding debt is backed by the monetary authority in the form of current and future seigniorage revenues. The monetary authority fully accommodates the fiscal authority whenever a budget deficit is financed with debt. This accommodation takes the form of an increase in current or future seigniorage revenues to back the principal and interest payments on the newly issued debt. The fiscal authority is insensitive to monetary policy in the sense that neither taxes nor expenditure react (now or in the future) to changes in the stock of outstanding government debt. Sargent, and Aiyagari and Gertler refer to this case as a polar Non-Ricardian regime. Leeper calls it one of passive monetary/active fiscal policy. Here, this case will be defined as one of complete fiscal dominance.

The long-run rule (6) is consistent with multiple period-by-period fiscal policy rules. As an example, consider the following version of the rule used by Aiyagari and Gertler (1985):

$$p_t(\tau_t - G_t) = \delta \left[(i_{t-1} - 1)\, B_{t-1} - (B_t - B_{t-1}) \right]. \tag{8}$$

Under (8), the nominal primary surplus is adjusted in every period (increasing τ_t or reducing G_t) in the exact amount needed to finance a fixed fraction δ of the interest on the outstanding debt (B_{t-1}) net of an adjustment for debt growth. To see that this stationary policy satisfies (6), simply iterate forward on (8) and use the government's no-Ponzi-game condition. In principle, there might be other period-by-period policy rules (perhaps not time-stationary) that are consistent with the rule (6). An advantage of this approach is that it allows both the determination of the price level and the construction of empirical estimates of δ using the long-run policy rule (6) without having to assume that a particular policy like (8) is satisfied in every period, for every country in the sample.

The parameter δ characterizes the degree of interdependence between fiscal and monetary authorities. In the paper, it will be treated as a "deep parameter," that reflects the revealed preferences of governments regarding the backing of its debt either by the fiscal or the monetary authority. This parameter should not be interpreted narrowly, as capturing a publicly announced policy commitment, or a commitment formally written in a country's budget, constitution, or central bank organic law. Instead, δ is a value that arises from the interaction of the fiscal and monetary authorities given a stable institutional setup. This interpretation is reinforced by the observation that the price level is derived here using a long-run fiscal policy rule without any reference to particular period-by-period fiscal or monetary policy rules.

Our specification of government behavior follows an earlier literature that describes monetary and/or fiscal policies in terms of explicit rules. See, among others, Taylor (1993) and Clarida, Galí, and Gertler (2000) for monetary policy rules; and Sargent and Wallace (1981), Aiyagari and Gertler (1985), Leeper (1991), and Bohn (1998) for fiscal policy rules. Leeper and Bohn point out that fiscal rules relating taxes to debt can be consistent with an optimizing government that minimizes the cost of tax collection by smoothing marginal tax rates over time [see Barro (1979)]. We view the δ-backing rule as a fairly unrestrictive way to parameterize government behavior that is convenient both analytically and empirically. It captures in a reduced-form way the idea that in response to different institutional settings, the monetary authority will face different obligations regarding fiscal policy. Whether this rule satisfies some optimality criterion, or whether it is a realistic description of government behavior beyond that just mentioned is an open question to be addressed in future research.

2.3 Equilibrium

The competitive equilibrium for this economy may be defined in an entirely standard way. Specifically, it corresponds to a price system, allocations for the representative consumer and the representative firm, and a government policy, such that (*i*) the representative consumer and the representative firm optimize given the government policy and the price system, (*ii*) the government policy is budget-feasible given the price system and the choices of consumers and firms, and (*iii*) markets clear.

In this model, the price level is determined by the clearing of the money market

$$M_t = m_t. \tag{9}$$

Money supply is determined by the combination of the fiscal rule and the government's intertemporal budget constraint [eq. (7)], while money demand is given by the consumer's intratemporal condition relating money and consumption [eq. (4)]. From equation (7), money supply can be written after some manipulations as

$$\frac{M_t}{p_t} = \frac{i_t}{i_t - 1}\left[(1-\delta)i_{t-1}\frac{B_{t-1}}{p_t} + \frac{M_{t-1}}{p_t} - \sum_{j=1}^{\infty}\left(\frac{M_{t+j}}{p_{t+j}R_t^{(j)}}\frac{i_{t+j}-1}{i_{t+j}}\right)\right]. \tag{10}$$

Using the equilibrium condition (9) and money demand (4) in (10) yields

$$\gamma c_t = (1-\delta)i_{t-1}\frac{B_{t-1}}{p_t} + \frac{M_{t-1}}{p_t} - \sum_{j=1}^{\infty}\left(\frac{m_{t+j}}{p_{t+j}R_t^{(j)}}\frac{i_{t+j}-1}{i_{t+j}}\right).$$

Exploiting the recursive nature of the Euler equation [eq. (3)] to find an expression for the infinite sum, $\sum_{j=1}^{\infty}(m_{t+j}/p_{t+j}R_t^{(j)})((i_{t+j}-1)/i_{t+j})$, in terms of current consumption, and after some algebra:

$$p_t = \frac{(1-\beta)(M_{t-1} + (1-\delta)i_{t-1}B_{t-1})}{\gamma c_t}. \tag{11}$$

This equation describes the aggregate price level as a function of consumption and of the beginning-of-period stocks of money and debt. Aiyagari and Gertler obtain an expression for the price level similar to the one above, but assuming a specific period-by-period rule and focusing on a stationary solution with constant asset prices.

As an alternative, one can use the fact that $M_{t-1} + (1-\delta)i_{t-1}B_{t-1} = M_t + (1-\delta)B_t$,[4] to write the price level in terms of the end-of-period stocks of money and debt:

$$p_t = \frac{(1-\beta)[M_t + (1-\delta)B_t]}{\gamma c_t}. \tag{12}$$

Note that equations (11) and (12) are equivalent, but the empirical analysis of (12) would not require data on the gross nominal interest rate. Regardless of whether one focuses on (11) or (12), this model implies that the price level depends not only on the money stock, but also on the proportion of the outstanding debt that is backed by money. In this sense, the proportion of the outstanding debt that is backed by money behaves like money itself.

Notice that the derivation of the aggregate price level, p_t, does not involve the production side of the economy. In particular, it does not involve the consumer's first-order conditions for their choice of capital and labor, the firm's first-order conditions, or the market clearing in goods and factors markets. Since this model displays the property of money superneutrality, the production side of the economy is solved in a completely independent set of equations that do not include nominal

[4] Write equation (7) as:

$$\begin{aligned}(M_t - M_{t-1})/p_t - (1-\delta)i_{t-1}B_{t-1}/p_t &= -\mathcal{S}_{t+1}/r_{t+1}, \\ &= -(1-\delta)i_tB_t/p_{t+1}r_{t+1}, \\ &= -(1-\delta)B_t/p_t,\end{aligned}$$

where the last line follows from multiplying and dividing the right-hand side by p_t, and using the definitions of gross inflation and gross real interest rate.

variables.[5] The consumption level, c_t, that appears in the denominator of (12) is determined in that subsystem as well. Thus, p_t is the outcome of monetary policy (reflected in the sequence of M_t) and how government debt is backed (summarized by the parameter δ).[6]

In order to develop further the reader's intuition, consider a long run situation where all real variables are constant. By dividing and multiplying the right-hand side of (12) by y, we obtain

$$p_t = \frac{M_t V}{y} + \frac{(1-\delta)B_t V}{y},$$

where $V \equiv (1-\beta)y/(\gamma c)$ can be interpreted as a measure of velocity of the broad monetary aggregate, $M_t + (1-\delta)B_t$, that consists of the sum of money and the monetized debt (*i.e.,* the proportion of debt that is backed by seigniorage). Note that only for the special case where $\delta = 1$, can the constant V be interpreted as money-velocity and the Quantity Theory of Money holds. More generally, for any $\delta \in [0, 1)$, the stock of debt plays a role in the determination of the price level. This point was made before by Aiyagari and Gertler.

Government debt also plays a crucial role in the determination of p_t under the Fiscal Theory of the Price Level (FTPL). The FTPL assumes that the government does not have to satisfy its intertemporal budget constraint for all possible sequences of p_t. Any particular path for the price level that does not satisfy the intertemporal budget constraint could be automatically excluded as an equilibrium by the government because it would not satisfy market clearing nor the consumer's optimality conditions. As a result of this assumption, p_t is determined as the quotient between the nominal value of interest-bearing debt and the present value of the all government revenues (including seigniorage) regardless of whether the government debt is, or will be, monetized. In contrast, in the model used here, the no-Ponzi-game condition on the government's behavior implies an intertemporal budget constraint that is satisfied for all price sequences and the equilibrium sequence is determined by the clearing of the money market.

This conceptual difference between the FTPL and this model has both theoretical and empirical implications. At the theoretical level it implies that, under the FTPL, B_t affects the price level even if it is never monetized, while in this model, only the proportion that is monetized (now or in the future) will affect p_t. The effect of B_t on p_t increases linearly with the proportion of debt that is backed by current or future seigniorage revenues, $(1-\delta)$. When $\delta = 1$, given the path of government expenditures, savings in the form of government debt will be used to pay future taxes. Consequently, debt has no effect on the current demand for goods or money and Ricardian

[5] In general, the Sidrauski model can exhibit nonsuperneutrality outside the steady state. Fischer (1979) shows that for the CRRA utility function, the rate of capital accumulation is positively related to the rate of money growth, except for the case of log-separable utility used here.

[6] Results are also robust to allowing distortionary taxation on capital and labor. The reason is that the Euler equation (3) and the intratemporal condition (4) are unchanged when the model is generalized in this manner. All that is required to make our results go through is to redefine $\mathcal{T}_t$ as the present discounted value of all lump-sum and distortionary taxes on capital and labor income.

equivalence holds. When $\delta \in [0,1)$, a proportion of debt does not require future tax increases but implies an increase in current and/or future seigniorage revenue. Anticipating future inflation, forward-looking agents reduce their current money demand and bid the price level up today.

At the empirical level, the next section will show that under certain conditions, the long-run dynamics of money, debt, and private consumption permit the econometric estimation of δ in our model. Statistical inference can then be used to draw conclusions regarding the policy regime (whether Ricardian or not) in a given economy. However, given the assumption that the intertemporal budget constraint is always satisfied, the econometric results have no direct bearing on the impossibility result in Cochrane (1998), whereby the FTPL cannot be falsified empirically because only equilibrium prices are observable.

3 Empirical Analysis

3.1 Econometric Strategy

This section describes a simple econometric strategy to obtain estimates of the parameter that measures the degree of interdependence between fiscal and monetary policies, δ. Rewrite equation (12) as:

$$M_t = \frac{\gamma}{(1-\beta)} C_t - (1-\delta) B_t, \tag{13}$$

where $C_t \equiv p_t c_t$ denotes nominal private consumption. Consider the empirical counterpart to this relation:

$$M_t = \alpha_0 + \alpha_1 C_t + \alpha_2 B_t + e_t, \tag{14}$$

where α_0 is an intercept, α_j for $j = 1, 2$ are constant coefficients, and e_t is a disturbance term that captures specification error. In terms of the structural parameters of the model, $\alpha_1 = \gamma/(1-\beta)$, and $\alpha_2 = -(1-\delta)$. Although not all structural parameters can be identified from the ordinary least squares (OLS) projection of M_t on C_t and B_t, δ would be identified from the coefficient on the stock of debt.

In principle, because all three variables are endogenous to the model, the OLS regression would yield biased and inconsistent estimates if the variables were covariance-stationary. However, if M_t, C_t, and B_t are nonstationary variables, and equation (13) is a cointegrating relationship, then the same regression would yield superconsistent parameter estimates (Phillips and Durlauf 1986).[7]

This approach is not the only one that could deliver estimates of the parameter δ. There are at least two other strategies. First, one could consider estimating δ directly from the fiscal rule (6).

[7] In principle, the reduced-form (14) may be written with either M_t, C_t, or B_t on the left-hand side. In adopting the formulation above, we are normalizing the coefficient of M_t in the cointegrating vector to unity. Provided M_t belongs to the cointegrating relation, results are robust to this normalization. The reason we choose to write the reduced-form in this manner is that its estimation delivers δ directly without the need to use, for example, the Delta method to compute its standard error.

An advantage of this strategy is that it would deliver a "theory-free" estimate without the need to model the consumer's behavior or make assumptions about functional forms. Unfortunately, this strategy requires the computation of the present discounted values $\mathcal{T}_t$ and $\mathcal{G}_t$ that involve infinite future values for taxes and government expenditure. Since the econometrician only has access to a finite number of observations, the implementation of this approach would necessarily involve truncation and the loss of many degrees of freedom.

Second, one could follow the literature and construct inferences about government behavior on the basis of particular period-by-period rules [see, for example, Bohn (1998)]. This strategy would overcome the problem created by the computation of infinite summations. However, it seems unlikely that the same period-by-period rule describes government behavior in a cross-section of countries with different institutional arrangements. Instead, the approach here makes the hypothesis of similar consumer preferences across countries (at least in terms of functional form if not of preference parameters) but avoids imposing a common period-by-period institutional framework for governments in different countries.

Notice that it is possible to identify δ even if the theoretical model only assumes a long-run fiscal policy rule, allowing any period-by-period rule that satisfies (6). The reason is that current money supply is derived directly from the implication of the long-run fiscal rule and the government's intertemporal budget constraint. Then, the money market equilibrium and the agents' first-order conditions are used to derive the price level. Thus, there is a sense in which the long-run rule is directly estimated, using the restrictions from economic theory to solve out the infinite sum.[8] Hence, by developing a fully-specified model, we can construct econometric inferences about the policy regime, even if we do not know the particular period-by-period rule followed by a given government in a given country.

3.2 Data

The empirical analysis is based on annual, nominal (in local currency), per-capita data on monetary base, government debt, and private consumption from 18 industrialized countries, all members of the Organization for Economic Cooperation and Development (OECD), and 20 developing economies. We included all IMF member countries for which reasonably long time series of the variables were available. In addition to data availability, the sample period for some countries was limited by substantial institutional changes. In particular, the sample for Germany ends before the reunification and the samples for member countries of the European Monetary Union end before the introduction of the Euro, in January 1999. Table 1 shows the cross-country sample used in the

[8] Recall that we used the money market equilibrium to substitute M's (money supply) with m's (money demand) in (10). Then, we used the agents' intratemporal condition (4) to express the infinite sum in terms of future consumption and, finally, we used consumption smoothing to write the infinite consumption sum in terms of current consumption alone.

empirical analysis.[9]

All series come from the International Financial Statistics (IFS) database compiled by the International Monetary Fund, with the exception of government debt for the United States, Canada and Brazil, which come from national sources.[10] For all other countries, government debt corresponds to the IFS series 88 (Total Debt), or the sum of IFS series 88a or 88b (Domestic Debt) with IFS series 89a or 89b (Foreign Debt). Monetary base corresponds to IFS series 14 (Reserve Money) or to the sum of IFS series 14a, 14c, and 14d, which are disaggregated liabilities of the monetary authority. Private consumption corresponds to the series 96F (Household Consumption Expenditures or Private Consumption). Population is IFS series 99Z..ZF (mid-year estimate of the total population by the United Nation's *Monthly Bulletin of Statistics*).

Table 1
Cross-Country Sample

OECD countries	Sample	Developing countries	Sample
Australia	1949 – 2002	Brazil	1964 – 2005
Austria	1970 – 1997	Colombia	1950 – 1987
Belgium	1953 – 1998	Costa Rica	1951 – 2002
Canada	1948 – 2005	El Salvador	1951 – 2000
Finland	1950 – 1997	Guyana	1955 – 1997
France	1951 – 1998	Honduras	1954 – 2004
Germany	1950 – 1990	India	1960 – 2001
Iceland	1950 – 2005	Indonesia	1972 – 2001
Italy	1962 – 1998	Israel	1972 – 2001
Luxembourg	1974 – 1997	South Korea	1953 – 1998
Netherlands	1951 – 1998	Malaysia	1960 – 1999
New Zealand	1970 – 2000	Malta	1960 – 2001
Norway	1971 – 2003	Mexico	1965 – 2005
Spain	1962 – 1998	Nigeria	1968 – 2004
Sweden	1950 – 2005	Morocco	1962 – 2005
Switzerland	1960 – 2004	Pakistan	1960 – 2003
United Kingdom	1970 – 1997	Philippines	1949 – 1994
United States	1948 – 2005	South Africa	1956 – 2005
		Thailand	1950 – 2005
		Tunisia	1971 – 1999

[9] We acknowledge the fact that data from member countries of the OECD may differ, in terms of quality and reliability, from data from developing economies. We also point out that OECD countries are market economies with relatively few prices under direct or indirect government control, which is not always the case for developing economies (for example, Argentina, Brazil, and Israel used widespread price and wage controls during inflation stabilization programs in the 1980s). These factors must be taken into account in the interpretation of results.

[10] For the United States, government debt is the Gross Federal Debt Held by the Public from the U.S. Department of Commerce, available from the Federal Reserve Bank of St. Louis (*www.stls.frb.org*). For Canada, it corresponds to the series D469409 (Net Federal Government Debt) in the CANSIM database of Statistics Canada, and for Brazil, it is represented by the series BM_DPIPP (end-of-period outsanding federal debt not held by the central bank), available from the *Banco Central do Brasil*

3.3 Results

The econometric strategy outlined in the previous section is valid only if M_t, C_t, and B_t are nonstationary variables and the OLS regression (14) is not spurious, but forms a cointegrating relationship. Unit-root and cointegration tests are used to assess both conditions.

Table 2
ADF Unit Root Test on C_t, M_t, and B_t

	C_t			M_t			B_t		
Country	lags	*t*-stat	*p*-value	lags	*t*-stat	*p*-value	lags	*t*-stat	*p*-value
OECD countries:									
Australia	1	-0.14	0.99	6	0.07	1.00	9	-1.64	0.76
Austria	0	-2.60	0.28	2	-1.69	0.73	2	-1.63	0.75
Belgium	1	-1.99	0.59	8	-1.06	0.92	1	-2.40	0.37
Canada	1	-0.89	0.95	1	-1.81	0.69	1	-2.68	0.25
Finland	7	0.61	1.00	9	-0.22	0.99	2	-0.69	0.97
France	1	-2.18	0.49	0	-1.74	0.09	1	0.94	1.00
Germany	5	0.32	1.00	1	0.18	1.00	3	0.81	1.00
Iceland	2	1.28	1.00	9	-1.19	0.90	5	-2.36	0.40
Italy	0	-1.48	0.82	1	-0.50	0.98	3	-0.44	0.98
Luxembourg	0	-2.49	0.33	1	-2.01	0.56	0	-0.63	0.97
Netherlands	1	-1.17	0.91	8	-0.24	0.99	1	-1.72	0.72
New Zealand	0	-2.37	0.39	0	-4.07	0.02	0	0.20	1.00
Norway	2	-1.06	0.92	2	0.05	1.00	0	-0.08	0.99
Spain	7	-0.91	0.94	1	-2.04	0.56	9	0.36	1.00
Sweden	3	-1.84	0.67	2	-2.19	0.48	8	-1.39	0.85
Switzerland	4	-2.50	0.33	1	-1.45	0.83	7	-0.09	0.99
United Kingdom	4	-1.64	0.74	0	-1.47	0.82	0	0.58	1.00
United States	9	-0.08	0.99	2	1.27	1.00	7	-1.19	0.90
Developing countries:									
Brazil	2	-0.25	0.99	1	0.40	1.00	2	1.14	1.00
Colombia	1	1.18	1.00	1	2.16	1.00	0	23.89	1.00
Costa Rica	3	0.49	1.00	1	-1.78	0.70	8	0.21	1.00
El Salvador	2	-0.85	0.95	0	-2.56	0.30	7	0.05	1.00
Guyana	1	0.91	1.00	0	1.12	1.00	5	0.38	1.00
Honduras	4	0.02	1.00	10	0.05	1.00	0	1.79	1.00
India	0	3.55	1.00	5	0.05	1.00	0	10.84	1.00
Indonesia	0	2.14	1.00	1	-0.71	0.96	2	-0.28	0.99
Israel	2	-2.44	0.35	0	-1.41	0.84	0	-2.49	0.33
Korea	0	0.65	1.00	8	0.28	1.00	1	1.96	1.00
Malaysia	9	-0.26	0.99	3	-0.29	0.99	1	-2.56	0.30
Malta	0	0.34	1.00	0	-1.82	0.68	1	0.66	1.00
Mexico	1	0.05	1.00	0	4.40	1.00	6	0.64	1.00
Morocco	1	-2.64	0.26	0	1.71	1.00	0	-1.76	0.71
Nigeria	4	1.97	1.00	2	0.97	1.00	9	0.22	1.00
Pakistan	5	0.23	1.00	4	1.46	1.00	4	0.08	1.00
Philippines	1	2.52	1.00	2	0.57	1.00	0	1.30	1.00
South Africa	5	1.64	1.00	10	-0.03	0.99	2	-0.85	0.95
Thailand	2	1.00	1.00	6	0.05	1.00	10	-0.15	0.99
Tunisia	0	0.94	1.00	1	-0.70	0.96	2	0.10	1.00

Notes :
(1) ADF Test equations include a constant and a linear trend.
(2) Number of lags selected according to the Modified Akaike Information Criteria.

Table 2 report results of augmented Dickey-Fuller (ADF) unit-root tests. For all ADF tests, the estimated alternative is a covariance-stationary autoregression with both a constant and a deterministic trend. The level of augmentation in the tests (i.e., the number of lagged first differences included in the OLS regression) is based on the Modified Information Criterion (MIC) proposed by Ng and Perron (2001).[11] Note that, for all countries, the null hypothesis of a unit root with drift cannot be rejected against the alternative of a deterministic trend at the five per cent significance level. The only exception is the per-capita nominal money stock of New Zealand.[12]

The null hypothesis of no cointegration is tested using the residual-based method proposed by Engle and Granger (1987) and Phillips and Ouliaris (1990). Gonzalo and Lee (1998) show that this test is more robust than Johansen's trace test to certain departures from unit root behavior like long memory and stochastic unit roots. The residual-based test requires running OLS on the relation of interest and then testing the hypothesis that the regression residuals have a unit root. Nonstationarity of the residuals constitutes evidence against cointegration. For some countries, the test results, reported in Table 3, depend on the method used to select the level of augmentation. Four different criteria are considered: sequential t-tests, Modified Akaike (MAIC), Modified Schwarz (MSIC), and a standard Schwarz information criteria.

Note that, for the OECD countries, rejection of no cointegration at the 15 per cent significance level or less is the common outcome from tests based on sequential t-tests and Schwarz lag-selection methods. For Finland, Germany, Iceland, Luxembourg, and Norway, tests based on MAIC and MSIC lag-selection methods suggest no cointegration. The null of no cointegration is also not rejected for Canada, when considering sequential t-tests and MAIC.

Among developing countries, on the other hand, there is strong evidence of no cointegration for Nigeria, Pakistan, and Thailand with the null hypothesis not being reject even at the 65 per cent level, regardless of the lag-selection method. For all other countries, the null hypothesis of no cointegration is not rejected at the 15 per cent level or less, for at least two lag-selection methods. The exception is South Korea, for which only the test based on MSIC lag-selection method shows evidence of cointegration.

Based on these results, it is reasonable to conclude that there is cointegration between nonstationary variables M_t, B_t, and C_t in all OECD countries, except New Zealand (since M_t is found to be stationary), and in all developing countries, except South Korea (weak evidence of no cointegration), and Nigeria, Pakistan, and Thailand (strong evidence against cointegration). For all other countries, the tests show evidence that 1) M_t, B_t, and C_t are nonstationary and, 2) for at least two different lag-selection methods, those variables form a cointegration relationship.

[11] For robustness to the lag-selection method, we also applied recursive t-tests with similar conclusions.

[12] Results (available from the author upon request) are robust to alternative unit root tests such as the KPSS test (Kwiatkowski, Phillips, Schmidt, and Shin 1992), the ADF test with GLS detrending, the ERS point-optimal test (Elliott, Rothenberg, and Stock 1996), the Phillips-Perron test (Phillips and Perron 1988), and the Ng-Perron modified unit root test (Ng and Perron 2001).

Table 3
Engle-Granger Cointegration Tests

	Lag Length Selection Criteria											
	seq. t-tests			MAIC			MSIC			Schwarz		
Country	*lags*	*t-stat*	*p-value*	*lags*	*t-stat*	*p-value*	*lags*	*t-stat*	*p-value*	*lags*	*t-stat*	*p-value*
OECD countries:												
Australia	7	-1.49	0.13	7	-1.49	0.13	7	-1.49	0.13	0	-6.28	0.00
Austria	0	-4.77	0.00	0	-4.77	0.00	0	-4.77	0.00	0	-4.77	0.00
Belgium	8	-2.28	0.02	0	-3.50	0.00	0	-3.50	0.00	0	-3.50	0.00
Canada	10	-1.15	0.23	10	-1.15	0.23	1	-2.12	0.03	0	-2.93	0.00
Finland	7	-3.41	0.00	4	-1.28	0.18	4	-1.28	0.18	10	-4.40	0.00
France	2	-3.37	0.00	0	-1.59	0.10	0	-1.59	0.10	2	-3.37	0.00
Germany	9	-4.49	0.00	0	-1.10	0.24	0	-1.10	0.24	1	-2.43	0.02
Iceland	7	-2.25	0.03	9	-1.33	0.17	6	-1.32	0.17	7	-2.25	0.03
Italy	3	-3.51	0.00	1	-0.41	0.53	1	-0.41	0.53	3	-3.51	0.00
Luxembourg	1	-1.87	0.06	0	-0.90	0.31	0	-0.90	0.31	1	-1.87	0.06
Netherlands	8	-2.19	0.03	0	-4.41	0.00	0	-4.41	0.00	8	-2.19	0.03
New Zealand	0	-4.73	0.00	0	-4.73	0.00	0	-4.73	0.00	3	-1.83	0.06
Norway	9	-2.88	0.01	2	-0.77	0.37	2	-0.77	0.37	0	-4.73	0.00
Spain	0	-2.84	0.01	0	-2.84	0.01	0	-2.84	0.01	0	-2.84	0.01
Sweden	7	-3.24	0.00	2	-2.13	0.03	2	-2.13	0.03	0	-2.13	0.03
Switzerland	0	-1.87	0.06	0	-1.87	0.06	0	-1.87	0.06	0	-1.87	0.06
United Kingdom	3	-3.07	0.00	0	-2.56	0.01	0	-2.56	0.01	0	-2.56	0.01
United States	1	-2.16	0.03	3	-1.39	0.15	3	-1.39	0.15	1	-2.16	0.03
Developing countries:												
Brazil	8	-1.66	0.09	0	-4.38	0.00	0	-4.38	0.00	8	-1.66	0.09
Colombia	9	-1.47	0.13	6	-0.19	0.61	6	-0.19	0.61	9	-1.47	0.13
Costa Rica	10	-6.64	0.00	0	-5.29	0.00	0	-5.29	0.00	10	-6.64	0.00
El Salvador	7	-1.25	0.19	7	-1.25	0.19	1	-1.98	0.05	0	-2.53	0.01
Guyana	7	-3.82	0.00	9	-0.66	0.42	4	-1.46	0.13	7	-3.82	0.00
Honduras	10	-2.09	0.04	6	-1.22	0.20	6	-1.22	0.20	10	-2.09	0.04
India	9	-2.15	0.03	7	-0.89	0.32	7	-0.89	0.32	0	-3.56	0.00
Indonesia	3	-3.14	0.00	0	-1.85	0.06	0	-1.85	0.06	3	-3.14	0.00
Israel	1	-2.98	0.00	0	-2.14	0.03	0	-2.14	0.03	1	-2.98	0.00
Korea	8	-0.97	0.29	9	-0.67	0.42	0	-2.87	0.01	8	-0.97	0.29
Malaysia	5	-2.13	0.03	4	-1.22	0.20	4	-1.22	0.20	5	-2.13	0.03
Malta	0	-3.34	0.00	0	-3.34	0.00	0	-3.34	0.00	0	-3.34	0.00
Mexico	9	-4.04	0.00	0	-2.38	0.02	0	-2.38	0.02	9	-4.04	0.00
Morocco	2	-1.41	0.15	1	-0.59	0.46	1	-0.59	0.46	0	-1.51	0.12
Nigeria	9	0.02	0.68	9	0.02	0.68	9	0.02	0.68	9	0.02	0.68
Pakistan	9	0.66	0.85	7	-0.33	0.56	7	-0.33	0.56	9	0.66	0.85
Philippines	0	-3.15	0.00	0	-3.15	0.00	0	-3.15	0.00	0	-3.15	0.00
South Africa	0	-5.15	0.00	9	-1.13	0.23	8	-1.26	0.19	0	-5.15	0.00
Thailand	7	0.65	0.85	6	0.05	0.69	6	0.05	0.69	7	0.65	0.85
Tunisia	7	-2.26	0.03	2	-1.39	0.15	2	-1.39	0.15	0	-4.40	0.00

Note : ADF Test equations do not include either constant or trend.

A common dilemma related to the use of the unit-root and cointegration tests has been their low power when applied to time series only available for the postwar period, since it is the span of the data, rather than the frequency, that maters for the power of these tests (Perron,1989, 1991; Pierse and Snell 1995). In the hope that inference about the existence of unit roots and cointegration can be made more straightforward and precise by combining information on the time series dimension with that from the cross-sectional dimension, a number of unit root tests using panel data techniques have been suggested (Banerjee 1999; Baltagi and Kao 2000).

Table 4
Panel Unit Root Tests

	C_t		M_t		B_t	
Method	*stat*	*p-value*	*stat*	*p-value*	*stat*	*p-value*
All countries:						
Null: Unit root (assumes common unit root process)						
Levin, Lin & Chu *t**	6.02	1.00	13.67	1.00	16.49	1.00
Breitung *t*-stat	14.28	1.00	10.90	1.00	10.62	1.00
Null: Unit root (assumes individual unit root process)						
Im, Pesaran and Shin W-stat	13.24	1.00	12.41	1.00	17.71	1.00
ADF - Fisher Chi-square	18.52	1.00	12.56	1.00	28.90	1.00
PP - Fisher Chi-square	19.71	1.00	35.82	1.00	27.68	1.00
Null: No unit root (assumes common unit root process)						
Hadri Z-stat	12.99	0.00	7.13	0.00	12.04	0.00
OECD countries:						
Null: Unit root (assumes common unit root process)						
Levin, Lin & Chu *t**	-0.18	0.43	4.48	1.00	0.20	0.58
Breitung *t*-stat	11.89	1.00	8.49	1.00	5.24	1.00
Null: Unit root (assumes individual unit root process)						
Im, Pesaran and Shin W-stat	4.44	1.00	4.84	1.00	6.66	1.00
ADF - Fisher Chi-square	13.77	1.00	14.36	1.00	8.82	1.00
PP - Fisher Chi-square	14.21	1.00	37.01	0.42	6.50	1.00
Null: No unit root (assumes common unit root process)						
Hadri Z-stat	13.10	0.00	5.55	0.00	12.71	0.00
Developing countries:						
Null: Unit root (assumes common unit root process)						
Levin, Lin & Chu *t**	10.58	1.00	13.42	1.00	24.10	1.00
Breitung *t*-stat	13.04	1.00	12.00	1.00	13.61	1.00
Null: Unit root (assumes individual unit root process)						
Im, Pesaran and Shin W-stat	14.17	1.00	14.22	1.00	20.29	1.00
ADF - Fisher Chi-square	4.92	1.00	4.39	1.00	5.42	1.00
PP - Fisher Chi-square	3.40	1.00	8.55	1.00	8.28	1.00
Null: No unit root (assumes common unit root process)						
Hadri Z-stat	13.96	0.00	11.84	0.00	11.18	0.00

Note: Tests include individual fixed effects and individual linear trends.

Tables 4, shows the results from six different panel unit root tests[13] applied to the whole cross-country sample and to subsamples of OECD and developing countries. All tests include a linear trend and individual fixed effects. The general conclusions for nonstationarity in M_t, B_t, and C_t are strongly confirmed in all tests.

To test for cointegration between M_t, B_t, and C_t in a heterogeneous panel framework, Pedroni's panel cointegration test (Pedroni 1999, 2004) is used. Two sets of statistics are considered: 1) four statistics pooled along the "within-dimension" (the panel cointegration statistics), constructed by summing both the denominator and the numerator terms over the cross-section dimension separately, and 2) three statistics based on pooling along the "between-dimension" (the group mean cointegration statistics), constructed by first dividing the numerator by the denominator and then summing over the cross-sectional dimension. That is, the former are based on estimators that pool the autoregressive coefficient, say ρ_i, across different cross-section members $i = 1, ..., N$, for the unit root tests on the estimated residuals, while the latter are based on simple averages of the individually estimated ρ_i's. As a consequence, the null hypothesis $H_0 : \rho_i = 1$ for all i, is tested against the alternative $H_1 : \rho_i = \rho < 1$ for all i (common value for the autoregreesive coefficient on the residuals), in the case of the panel statistics, and against $H_1 : \rho_i < 1$ for all i, in the case of group mean statistics. Results, displayed in Table 5, strongly suggest the rejection of the null, in favor of cointegration.

Table 5
Pedroni's Panel Cointegration Test

	All countries		**OECD countries**		**Developing countries**	
Statistic	stat	*p*- value	stat	*p*- value	stat	*p*- value
panel ν -stat	7.9625	0.0000	7.9625	0.0000	7.1886	0.0000
panel ρ -stat	-8.4855	0.0000	-6.1984	0.0000	-5.6833	0.0000
panel pp-stat	-7.5782	0.0000	-5.6172	0.0000	-5.0295	0.0000
panel adf-stat	-5.4165	0.0000	-3.6561	0.0001	-4.1159	0.0000
group ρ -stat	-7.2915	0.0000	-5.3058	0.0000	-5.0171	0.0000
group pp-stat	-7.8870	0.0000	-5.4015	0.0000	-5.7471	0.0000
group adf-stat	-6.0431	0.0000	-3.1040	0.0010	-5.3851	0.0000

These results, both from time series and panel frameworks, are important because they allow an empirical description of the money market equilibrium as a cointegrating relationship for most countries in the sample. This means that even if the individual series can be represented as nonstationary processes, the behavioral rules and constraints of the model economy imply that a precise combination of these variables should be stationary. Hence, a simple Least Squares regression yields a superconsistent estimate of the parameter that characterizes the interdependence between

[13] For the ADF and PP Fisher-type tests, see Maddala and Wu (1999) and Choi (2001). For the other tests, see Levin, Lin and Chu (2002), Breitung (2000), Im, Pesaran and Shin (2003), and Hadri (1999).

fiscal and monetary policies.[14]

For the estimation of the cointegrating vector, we employ the dynamic ordinary least squares (DOLS) method proposed by Stock and Watson (1993). This method is asymptotically equivalent to maximum likelihood but exploits the functional relationship predicted by the model. This approach involves running the OLS regression:

$$M_t = \alpha_0 + \alpha_1 C_t + \alpha_2 B_t + \sum_{s=-p}^{q} \xi_{1,s} \Delta C_{t-s} + \sum_{s=-p}^{q} \xi_{2,s} \Delta B_{t-s} + e_t, \tag{15}$$

where $\xi_{j,s}$ for $j = 1, 2$ and $s = -p, -p+1, \ldots, q-1, q$ are constant coefficients. The appropriate number of leads and lags was selected using the Modified Akaike Information Criteria.

Table 6 presents estimates of the structural parameters. Nigeria, Pakistan, and Thailand are excluded from the sample, since no evidence of cointegration was found for those countries. However, South Korea is included even though evidence of cointegration is weak, and New Zealand stays on the sample on the basis of an Ng$-$Perron modified unit root test (Ng and Perron 2001) that does not allow the rejection of the null hypothesis of a unit root on M_t at the ten per cent significance level, or less. Needless to say, estimates for these two countries should be regarded with more caution.

In Table 6, the p-values for $\widehat{\alpha}_1$and $\widehat{\alpha}_2$, and the confidence interval for $\widehat{\delta}$ are based on rescaled standard errors. Standard errors are rescaled to take into account the serial correlation of the residuals that remains after adding the p leads and q lags (see, Hayashi 2000, pp. 654$-$657). Notice that, although the weight of real balances in the utility function (γ) and the subjective discount rate (β) are not separately identified, the coefficient on nominal consumption, $\alpha_1 = \gamma/(1-\beta)$ should be positive. Among the developed economies, except for Iceland and Luxembourg, $\widehat{\alpha}_1$ is positive and statistically different from zero. In the developing countries subsample, the exceptions are Israel, Philippines, and Tunisia, for which $\widehat{\alpha}_1$ is not statistically significant at the ten per cent level.[15]

Estimates of δ are identified from the reduced-form parameter $\alpha_2 = -(1-\delta)$. These estimates, $\widehat{\delta}$, are reported in Column 8 of Table 6. In all cases, this parameter is positive, and with the exception of Costa Rica and Malta, statistically different from zero. At a first glance, the two groups of countries do not seem to be much difference regarding the degree of fiscal dominance: given the 95 per cent confidence intervals, in 6 out of 18 OECD countries (Austria, Belgium, France, Italy, New Zealand, and the United States) and in 7 out of 17 developing countries (Brazil, Costa Rica, South Korea, Malta, Mexico, Morocco, and South Africa), we cannot reject the null hypothesis

[14] Elliot (1998) shows that even if the model variables have roots near but not exactly equal to one, the point estimates of the cointegrating vector are consistent. However, hypothesis tests regarding the coefficients that do not have an exact unit root can be subject to size distortions.

[15] All regressions include the intercept term (not reported), α_0. The theoretical model predicts that the intercept should be zero [see eq. (13)]. However, for some countries in the sample, the intercept was found to be statistically different from zero. Strictly speaking, this constitutes a rejection of the theory. A more constructive interpretation of this result is that the theoretical relation holds *up to* a constant term.

that $\widehat{\delta} < 1$.

Table 6
DOLS Estimation of Structural Parameters

	leads	lags	α_1		α_2		δ		Valid Sample		
	p	q	estimate	p-value	estimate	p-value	estimate	95% conf. interval	start	end	obs
OECD countries:											
Australia	0	0	0.060	0.0000	0.163	0.0002	1.163	[1.081 , 1.245]	1950	2002	53
Austria	0	1	0.178	0.0000	-0.035	0.0228	0.965	[0.936 , 0.995]	1972	1997	26
Belgium	1	3	0.180	0.0000	-0.044	0.0000	0.956	[0.949 , 0.963]	1957	1997	41
Canada	1	1	0.042	0.0000	0.013	0.0032	1.013	[1.005 , 1.022]	1950	2004	55
Finland	1	1	0.182	0.0000	-0.006	0.4967	0.994	[0.974 , 1.013]	1952	1997	46
France	2	1	0.152	0.0000	-0.070	0.0003	0.930	[0.895 , 0.965]	1953	1996	44
Germany	1	0	0.162	0.0000	0.009	0.8147	1.009	[0.935 , 1.082]	1951	1989	39
Iceland	3	2	-0.002	0.9453	0.092	0.0704	1.092	[0.992 , 1.193]	1953	2002	50
Italy	1	1	0.489	0.0000	-0.140	0.0000	0.860	[0.801 , 0.919]	1964	1997	34
Luxembourg	0	1	-0.028	0.0163	-0.142	0.4464	0.858	[0.471 , 1.245]	1976	1997	22
Netherlands	2	1	0.067	0.0572	0.077	0.0003	1.077	[1.039 , 1.115]	1953	1996	44
New Zealand	0	1	0.055	0.0000	-0.030	0.0165	0.970	[0.945 , 0.994]	1972	2000	29
Norway	0	0	0.115	0.0000	-0.021	0.3374	0.979	[0.935 , 1.023]	1972	2003	32
Spain	0	1	0.220	0.0157	-0.072	0.3084	0.928	[0.787 , 1.070]	1964	1998	35
Sweden	1	1	0.070	0.0037	0.012	0.4995	1.012	[0.976 , 1.048]	1952	2004	53
Switzerland	2	1	0.129	0.0004	-0.040	0.5324	0.960	[0.831 , 1.089]	1962	2002	41
United Kingdom	0	1	0.052	0.0000	-0.019	0.1372	0.981	[0.956 , 1.007]	1972	1997	26
United States	2	1	0.122	0.0000	-0.032	0.0003	0.968	[0.952 , 0.984]	1950	2003	54
Developing countries:											
Brazil	1	3	0.399	0.0000	-0.053	0.0000	0.947	[0.935 , 0.960]	1968	2004	37
Colombia	1	1	0.233	0.0000	-0.422	0.0514	0.578	[0.153 , 1.003]	1952	1986	35
Costa Rica	1	1	0.611	0.0000	-0.762	0.0000	0.238	[-0.066 , 0.543]	1953	2001	49
El Salvador	1	1	0.018	0.0000	0.004	0.7558	1.004	[0.978 , 1.030]	1953	1999	47
Guyana	1	1	0.334	0.0118	-0.015	0.0965	0.985	[0.968 , 1.003]	1957	1996	40
Honduras	5	1	0.156	0.0010	-0.023	0.6407	0.977	[0.876 , 1.077]	1956	1999	44
India	0	1	0.137	0.0026	0.075	0.0975	1.075	[0.986 , 1.164]	1962	2001	40
Indonesia	0	1	0.160	0.0000	0.084	0.0334	1.084	[1.007 , 1.160]	1974	2001	28
Israel	0	0	0.107	0.5206	0.104	0.3106	1.104	[0.897 , 1.310]	1973	2001	29
South Korea	1	1	0.197	0.0000	-0.358	0.0174	0.642	[0.351 , 0.933]	1955	1997	43
Malaysia	1	1	0.695	0.0000	-0.145	0.0962	0.855	[0.681 , 1.028]	1962	1998	37
Malta	1	0	1.386	0.0000	-0.908	0.0000	0.092	[-0.037 , 0.222]	1961	2000	40
Mexico	1	0	0.168	0.0000	-0.200	0.0304	0.800	[0.619 , 0.980]	1966	2004	39
Morocco	1	1	0.599	0.0000	-0.192	0.0000	0.808	[0.760 , 0.856]	1964	2004	41
Philippines	1	1	0.082	0.1501	-0.083	0.3929	0.917	[0.715 , 1.120]	1953	1988	36
South Africa	1	1	0.120	0.0000	-0.056	0.0000	0.944	[0.922 , 0.965]	1958	2004	47
Tunisia	1	0	0.060	0.2549	0.070	0.1242	1.070	[0.979 , 1.162]	1972	1998	27

Notes:
(1) Individual DOLS equations include a constant.
(2) Number of leads and lags selected according to the Modified Akaike Information Criteria.

However, note that the point estimate of δ is frequently closer to 1 among OECD countries than within the developing countries subsample. There are only two OECD countries (Italy, and Luxembourg) for which $\widehat{\delta}$ is lower than 0.9, as opposed to 7 developing countries (Colombia, Costa Rica, South Korea, Malaysia, Malta, Mexico, and Morocco). As previously mentioned, we cannot even reject the hypothesis that $\widehat{\delta} = 0$ for Costa Rica and Malta.[16] For the two groups of countries, Table 7 shows averages of both the estimated $\widehat{\delta}$ and the 95 per cent confidence lower bound, δ_L. Notice that the average $\widehat{\delta}$ and $\widehat{\delta}_L$ are higher in the OECD countries subsample, and the differences with respect to the developing countries are highly significant.

Table 7
Tests of Equality of Means

	OECD	Developing	*t*-stat	*p*-value
mean δ	0.984	0.831	10.2918	0.0000
mean δ_L	0.914	0.690	8.9805	0.0000
No. obs	18	17		

Regarding the structural parameters of the model, another strategy to assess the differences between OECD and developing countries is the use of heterogeneous panel estimation of long-run relationships (Pesaram and Smith 1995). Table 8 shows the results of three different fixed-effects panel data estimations of equation (15). In the estimation of model 1, coefficients α_1 and α_2 are both assumed to be common along the cross-section of countries. In model 2, they are both assumed to be country-specific. In model 3, $\alpha_1$is common along the cross-sectional dimension, but α_2 is country-specific. Notice that country-specific estimates of α_2, both from models 2 and 3, imply higher values of δ for OECD countries in comparison with developing economies. In addition, in both models 2 and 3, at the five per cent level, we cannot reject the hypothesis that $\delta\,(OECD) = 1$, but we can reject the hypothesis that $\delta\,(\textit{developing}) \geq 1$. As shown in Table 9, in both models 2 and 3, we can strongly reject the null hypothesis $\delta\,(OECD) = \delta\,(\textit{developing})$ using a Wald test.

Table 8
Panel DOLS Estimation of Structural Parameters

		α_1		α_2		δ	
	country	estimate	*p*-value	estimate	*p*-value	estimate	95% conf. interval
model 1:	all	0.090	0.0000	-0.002	0.4955	0.998	[0.990 , 1.005]
model 2:	OECD	0.085	0.0000	-0.004	0.3425	0.996	[0.989 , 1.004]
	developing	0.126	0.0000	-0.079	0.0000	0.921	[0.904 , 0.938]
model 3:	all	0.096	0.0000				
	OECD			-0.008	0.0365	0.992	[0.982 , 1.003]
	developing			-0.025	0.0000	0.975	[0.963 , 0.987]

[16] The theoretical model implies that δ is bounded between zero and one. Rather than incorporating a nonlinear restriction in a linear estimation framework, we follow the simpler approach of first estimating the cointegrating vector and then verifying whether $\hat{\delta}$ falls in the $[0, 1]$ range. This is not the case for Australia, Canada, Germany, Iceland, Netherlands, and Sweden, among OECD countries, and El salvador, India, Indonesia, and Tunisia, in the developing countries' subsample. However, for those countries, except for Australia, Canada (marginaly), Netherlands, and Indonesia, the hypothesis that its true value is 1 cannot be rejected at the 5 per cent level.

Table 9

Wald Test Results

H_0: δ (OECD) = δ (developing)

model 2:

Statistic	Value	df	p-value
F-stat	24.5	(1, 1294)	0.0000
χ^2	24.5	1.000	0.0000

model 3:

Statistic	Value	df	p-value
F-stat	15.2	(1, 1309)	0.0001
χ^2	15.2	1.000	0.0001

Recall that δ is the proportion of current government debt that is backed by the present discounted value of current and future primary surpluses. Hence, finding that δ is more likely to be closer to 1 in the OECD subsample means that outstanding debt in developed economies is essentially backed by the fiscal authority. Backing takes the form of a commitment to adjust the stream of future primary surpluses to match the current value of its bond obligations. In the long-run, there is complete accommodation of fiscal policy to the open market operations by the monetary authority. For example, when the monetary authority sells government bonds, the fiscal authority increases current or future taxes, and/or reduces current or future expenditures, to back the principal and interest payments on the newly issued debt.

This finding also suggests that the interdependence between fiscal and monetary authorities in developed economies is well described by what Sargent (1982) and Aiyagari and Gertler (1985) refer to as a Ricardian regime or, in the language of Leeper (1991), an active monetary/passive fiscal policy regime. In this regime, the fiscal authority backs all outstanding debt, debt plays only a minor role in the determination of the price level, and the Quantity Theory of Money holds as a long-run proposition. Regarding their fiscal/monetary regimes, most industrial countries do not seem to display signs of fiscal dominance.

In terms of Sargent and Wallace's (1981) coordination game between monetary and fiscal authorities, the results imply that, for most OECD countries in the sample, the central bank is the first mover. That is, the monetary authority sets its policy in advance and imposes discipline on the fiscal authority, meaning that the fiscal authority must select a sequence of primary surpluses (and debt) that is consistent with the sequence of M_t supplied by the monetary authority such that the intertemporal budget constraint is always satisfied. In turn, this implies that the unpleasant monetarist arithmetic might not be empirically relevant for developed economies and that "tough" central banks can fight inflation with tight money.

Table 10(a)
DOLS Estimation of Structural Parameters (with breaks)

	leads	lags	α_1		α_2		δ		Breaks	
	p	q	estimate	p-value	estimate	p-value	estimate	95% conf. interval	period	95% conf. interval
Australia	0	0	0.060	0.0000	0.163	0.0002	1.163	[1.081 , 1.245]	1950-2002	No break
Austria	0	1	0.154	0.0000	0.009	0.6700	1.009	[0.965 , 1.053]	< 1986	[1984 , 1990]
					-0.011	0.4890	0.989	[0.957 , 1.021]	> 1986	
Belgium	1	3	3.000	0.1797	0.000	-0.0442	0.000	[0.956 , 0.949]	1957-1997	No break
Canada	1	1	0.035	0.0000	0.059	0.0000	1.059	[1.041 , 1.077]	< 1982	[1980 , 1983]
					0.026	0.0000	1.026	[1.014 , 1.039]	1982-1994	[1992 , 2003]
					0.023	0.0000	1.023	[1.013 , 1.034]	> 1994	
Finland	1	1	0.148	0.0000	-0.238	0.0020	0.762	[0.616 , 0.907]	< 1984	[1982 , 1988]
					-0.003	0.7000	0.997	[0.979 , 1.014]	> 1984	
France	2	1	0.151	0.0000	-0.209	0.0000	0.791	[0.699 , 0.882]	< 1984	[1982 , 1989]
					-0.097	0.0040	0.903	[0.838 , 0.967]	> 1984	
Germany	1	0	0.174	0.0000	-0.264	0.0070	0.736	[0.548 , 0.924]	< 1972	[1970 , 1985]
					-0.013	0.7210	0.987	[0.913 , 1.061]	> 1972	
Iceland	3	2	0.006	0.8140	0.158	0.0010	1.158	[1.074 , 1.243]	< 1989	[1987 , 1990]
					0.088	0.0220	1.088	[1.014 , 1.162]	> 1989	
Italy	1	1	0.395	0.0000	-0.164	0.0000	0.836	[0.784 , 0.887]	< 1986	[1984 , 1992]
					-0.109	0.0000	0.891	[0.839 , 0.943]	> 1986	
Luxembourg	0	1	-0.028	0.0163	-0.142	0.4464	0.858	[0.471 , 1.245]	1976-1997	No break
Netherlands	2	1	0.067	0.0572	0.077	0.0003	1.077	[1.039 , 1.115]	1953-1996	No break
New Zealand	0	1	0.055	0.0000	-0.030	0.0165	0.970	[0.945 , 0.994]	1972-2000	No break
Norway	0	0	0.103	0.0000	-0.046	0.2510	0.954	[0.873 , 1.035]	< 1996	[1994 , 2003]
					-0.001	0.9830	0.999	[0.919 , 1.079]	> 1996	
Spain	0	1	0.272	0.0000	0.129	0.0030	1.129	[1.048 , 1.210]	< 1991	[1989 , 1993]
					-0.060	0.1090	0.940	[0.865 , 1.014]	> 1991	
Sweden	1	1	0.094	0.0120	-0.148	0.3620	0.852	[0.529 , 1.175]	< 1986	[1981 , 1990]
					0.061	0.0540	1.061	[0.999 , 1.123]	1986-1996	[1990 , 2003]
					0.009	0.7410	1.009	[0.953 , 1.065]	> 1996	
Switzerland	2	1	0.048	0.1640	0.517	0.0000	1.517	[1.333 , 1.701]	< 1989	[1987 , 1992]
					0.127	0.0630	1.127	[0.992 , 1.261]	> 1989	
United Kingdom	0	1	0.048	0.0000	0.020	0.2010	1.020	[0.988 , 1.052]	< 1981	[1979 , 1982]
					-0.008	0.5130	0.992	[0.967 , 1.017]	> 1981	
United States	2	1	0.116	0.0000	-0.066	0.2280	0.934	[0.826 , 1.043]	< 1961	[1960 , 1962]
					0.004	0.8120	1.004	[0.969 , 1.039]	1961-1981	[1979 , 1982]
					-0.021	0.1370	0.979	[0.951 , 1.007]	> 1981	

Note : Breaks selected according to the Bayesian Information Criteria (BIC) and/or the sequential method at the 10% significance level.

Table 10(b)
DOLS Estimation of Structural Parameters (with breaks)

	leads	lags	α_1		α_2		δ		Breaks	
	p	q	estimate	p-value	estimate	p-value	estimate	95% conf. interval	period	95% conf. interval
Brazil	1	3	0.386	0.0000	0.044	0.0000	1.044	[1.034 , 1.054]	< 2000	[1998 , 2001]
					-0.004	0.3270	0.996	[0.989 , 1.004]	> 2000	
Colombia	1	1	0.239	0.0000	-0.463	0.0448	0.537	[0.085 , 0.988]	1954-1986	No break
Costa Rica	1	1	0.611	0.0000	-0.762	0.0000	0.238	[-0.066 , 0.543]	1953-2001	No break
El Salvador (1)	1	1	0.026	0.0000	-0.270	0.0060	0.730	[0.544 , 0.916]	< 1970	[1968 , 1998]
					0.054	0.0000	1.054	[1.028 , 1.081]	1970-1987	[1985 , 1994]
					-0.025	0.0480	0.975	[0.949 , 1.000]	> 1987	
El Salvador (2)	1	1	0.027	0.0000	0.060	0.0000	1.060	[1.030 , 1.090]	< 1987	[1986 , 1988]
					-0.029	0.0490	0.971	[0.942 , 1.000]	> 1987	
Guyana	1	1	0.194	0.0070	0.125	0.0000	1.125	[1.106 , 1.144]	< 1990	[1989 , 1991]
					-0.007	0.1490	0.993	[0.984 , 1.003]	> 1990	
Honduras	3	1	0.252	0.0000	-0.132	0.0000	0.868	[0.807 , 0.929]	< 1988	[1986 , 1995]
					0.023	0.4010	1.023	[0.968 , 1.077]	> 1988	
India	0	1	0.137	0.0026	0.075	0.0975	1.075	[0.986 , 1.164]	1962-2001	No break
Indonesia	0	1	0.144	0.0000	-0.006	0.7090	0.994	[0.959 , 1.029]	< 1998	[1997 , 1999]
					0.104	0.0000	1.104	[1.070 , 1.138]	> 1998	
Israel	0	0	-0.194	0.1360	0.173	0.0290	1.173	[1.019 , 1.326]	< 1997	[1992 , 1999]
					0.266	0.0020	1.266	[1.109 , 1.423]	> 1997	
South Korea	2	2	0.197	0.0000	-0.358	0.0174	0.642	[0.351 , 0.933]	1955-1997	No break
Malaysia	1	1	0.732	0.0000	-0.380	0.0000	0.620	[0.462 , 0.778]	< 1987	[1985 , 1988]
					0.205	0.0000	1.205	[1.107 , 1.302]	> 1987	
Malta	1	0	1.393	0.0000	-0.957	0.0000	0.043	[-0.102 , 0.189]	1961-1999	No break
Mexico (1)	1	0	0.067	0.1610	0.039	0.7640	1.039	[0.944 , 1.135]	< 1998	[1997 , 1999]
					0.116	0.4530	1.116	[0.804 , 1.428]	> 1998	
Mexico (2)	1	0	0.191	0.0000	-0.131	0.1690	0.869	[0.680 , 1.058]	< 1990	[1989 , 1990]
					-0.267	0.0070	0.733	[0.544 , 0.922]	> 1990	
Morocco	1	1	0.324	0.0010	-0.054	0.3610	0.946	[0.826 , 1.065]	< 1998	[1997 , 1998]
					0.038	0.5970	1.038	[0.893 , 1.183]	> 1998	
Philippines	1	1	0.315	0.0000	-0.623	0.0000	0.377	[0.231 , 0.523]	< 1984	[1983 , 1985]
					-0.140	0.0000	0.860	[0.804 , 0.916]	> 1984	
South Africa	1	1	0.117	0.0000	-0.082	0.0000	0.918	[0.891 , 0.945]	< 1989	[1987 , 2003]
					-0.056	0.0000	0.944	[0.923 , 0.964]	> 1989	
Tunisia	1	0	0.112	0.0290	0.010	0.8250	1.010	[0.921 , 1.098]	< 1994	[1991 , 1997]
					0.027	0.5190	1.027	[0.942 , 1.112]	> 1994	

Notes:
(1) Breaks selected according to Bayesian Information Criteria (BIC)
(2) Breaks selected according to the sequential method at the 10% significance level.

We also explore the possibility of regime shifts in the DOLS estimation of (15) shown in Table 6. Since α_1 is a "policy-free" parameter that depends only on preferences, we assume that it is not allowed to change. Structural breaks are only allowed for α_2, which means changes in the $\delta-$ Backing Fiscal Policy Rule. Tables 10(a) and 10(b) show results based on the Bai-Perron procedure for the estimation of linear models with multiple structural changes (Bai and Perron 1998, 2003). In the estimations, a maximum number of two breaks are allowed to be endogenously determined by the data.

Notice that [see Table 10(a)] among OECD economies, even though breaks are found in all but five countries in the subsample, results tend to confirm those of Table 6. In general, the identified structural breaks in $\widehat{\delta}$ do not imply big qualitative changes in terms of the degree of fiscal dominance in OECD countries. Countries for which high values of $\widehat{\delta}$ were reported in Table 6 also display point estimates of $\widehat{\delta}$ that are close to 1 both before and after the breaks. The exceptions are Finland, France, Germany, and Sweden[17], all of which seem to have moved from a higher degree of fiscal dominance $(0.76, 0.79, 0.74$, and 0.85, respectively) to a higher degree of central bank independence as the $\widehat{\delta}$ estimates obtained for the post-break periods are closer to 1, in line with the results of Table 6. For two countries, Sweden and the United States, two structural breaks are found, but in both cases $\widehat{\delta}$ is not statistically different from 1 both pre- and post-breaks. Estimates of $\widehat{\delta}$ that are statistically higher than 1 are found in Canada, Iceland, Netherlands, Spain, and Switzerland.

Similarly, among developing countries [see Table 10(b)], no breaks are found in five countries, and results are consistent with the no-break DOLS estimation shown in Table 6. Important structural changes in the form of increases in the degree of fiscal dominance seem to have taken place in Honduras, Malaysia, and Philippines. A significant reduction in $\widehat{\delta}$ is found for Mexico. [18]

To summarize, the results of this section are as follows:

1. Most of the industrial countries and some developing countries can be reasonably described as economies with low degrees of fiscal dominance and/or higher levels of central bank independence. Fiscal dominance is more common in developing countries.

2. The degree of fiscal dominance is lower on average among OECD countries

The empirical results discussed above are consistent with findings in Fischer, Sahay, and Vegh (2002). These authors use annual panel data from 133 market economies and report that the expected negative relationship between fiscal balance and inflation is not verified for low-inflation, mostly developed, countries. A possible explanation of their finding is that in a fiscal regime of zero

[17] Although in the case of Sweden, the coefficient is not statistically different from 1 in all subperiods, both before and after the breaks.

[18] Two different criteria for the selection of break dates are used, the Schwarz or Bayesian Information Criteria, and a sequential method described in Bai and Perron (1998, 2003). Usually, they produce the same results, but El Salvador and Mexico are exceptions.

fiscal dominance, government debt plays no role in the determination of the price level. This point is related to Sargent's (1982) observation that "one cannot necessarily prove that current deficits are not inflationary by running time-series regressions and finding a negligible effect." The reason is that the question of whether budget deficits are inflationary is intimately related to the policy regime and institutional arrangements.

Results for the U.S. economy are also in line with previous work by Bohn (1998) and Canzoneri, Cumby, and Diba (2001), which suggest that fiscal authorities respond to the level of debt by raising primary surpluses. Bohn finds that, in the United States, an increase in government debt by $100 leads to an increase in the primary surplus by $5.40 in the following year. Canzoneri, Cumby, and Diba (2001) use impulse-response analysis to examine the response of U.S. government debt to a positive innovation in the primary surplus (including seigniorage revenue) and report a negative, persistent, and statistically significant debt response that is explained as the government paying off some of its previously accumulated debt.

3.4 Additional Implications

This subsection examines some additional empirical implications of the model. First, it may be helpful to compare the measure of fiscal dominance obtained here with indices of central bank independence (CBI) available in the literature (for a survey, see Arnone, Laurens, and Segalotto 2006). The comparison with indices of central bank independence is motivated by the idea that δ summarizes the interaction between fiscal and monetary authorities in a given institutional setup, meaning not only the legal characteristics of the central bank's organic law, but also to the informal policy decision-making in practice. Hence, estimates of δ obtained from actual data may capture both formal and informal behavioral elements.

Some CBI indices are constructed on the basis of scores, or points, attached to different legal aspects of central bank operation (Cuckierman, Webb and Neyapti 1992; Grilli, Masciandaro and Tabellini 1991; Eijffinger and Schaling 1993; Alesina and Summers 1993).[19] They measure central bank independence by focusing primarily on legal characteristics like the terms of office of the central bank director(s), restrictions on public sector borrowing from the central bank, conflict resolution between the central bank and the executive branch, etc.

However, since *de jure* central bank independence may be very different from *de facto* autonomy from the fiscal authority, Cuckierman (1992) and Cuckierman, Webb and Neyapti (1992) propose the use of the average turnover rate of central bank governors. Sturn and Haan (2001) update those studies to include more countries in the sample. The idea is that above a certain threshold this indicator may be a proxy for actual central bank independence, which makes it less relevant for developed economies. Rather than autonomy, low turnover rates may reflect subordination

[19] See also Bade and Parkin (1982).

of governors who want to keep their jobs, but high enough turnover rates may imply a higher likelihood that the term of office of the governor is shorter than the average term of a government, which dissuades the central bank from taking a long term view of monetary policy.

Table 11 displays the correlations between a $\delta-$based CBI index and other indices. Correlations with the value of the point estimate, $\widehat{\delta}$, are also presented. The $\delta-$based CBI index is computed according to the average of scores using the following mapping from the country-specific point estimates and (95% confidence interval) lower bounds, $\widehat{\delta}_L$, to a scale from 1 to 5:

Estimated $\widehat{\delta}$, $\widehat{\delta}_L$	Score
≥ 0.99	5.0
$\in [0.95\ ,\ 0.99)$	4.5
$\in [0.90\ ,\ 0.95)$	4.0
$\in [0.85\ ,\ 0.90)$	3.5
$\in [0.80\ ,\ 0.85)$	3.0
$\in [0.75\ ,\ 0.80)$	2.5
$\in [0.50\ ,\ 0.75)$	2.0
< 0.50	1.0

Note that the expected positive correlation between the $\delta-$based CBI index and *de jure* CBI indices is only statistically significant when considering the GMT autonomy index by Grilli, Masciandaro and Tabellini (1991), which is only available for 11 OECD countries in our sample. Figure 1 shows the positive relationship between the $\delta-$based CBI index and the GMT index.

Table 11
Correlations between δ and other CBI indices

	CWN			GMT				
	legal	turnover	AS	political	economic	overall	ES	SH
δ-cbi:								
corr.	0.0185	**-0.4310**	0.1737	0.2283	**0.5425**	**0.4710**	-0.1599	0.1511
df	30	29	12	11	11	11	7	12
t-stat	0.102	-2.572	0.611	0.778	2.142	1.771	-0.429	0.530
p-value	0.9198	0.0155	0.5526	0.4531	0.0554	0.1043	0.6810	0.6061
δ:								
corr.	-0.0491	**-0.4742**	0.1216	0.2267	0.4154	0.3930	-0.1093	0.1319
df	30	29	12	11	11	11	7	12
t-stat	-0.2694	-2.9003	0.4242	0.7719	1.5148	1.4175	-0.2910	0.4609
p-value	0.7895	0.0070	0.6789	0.4565	0.1580	0.1840	0.7795	0.6531

Notes:
CWN = Cukierman, Webb, and Neyapti (1992)
GMT = Grilli, Masciandaro, and Tabelini (1991)
ES = Eijffinger and Schaling (1993)
SH = Sturn and Haan (2001)
AS: Alesina and Summers (1993)

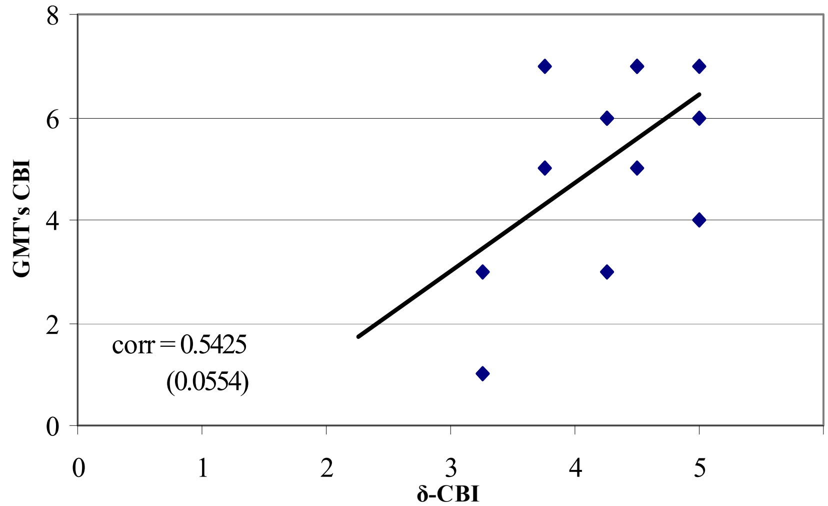

Figure 1: Relationship between δ-CBI and GMT's Economic Autonomy Index

However, the legal CBI index by Cuckierman, Webb and Neyapti (1992), which includes 30 countries from our sample, both industrialized and developing economies, is not correlated with the $\delta-$based measure of CBI (see Figure 2). This suggests that δ may capture legal aspects of CBI that are relevant for OECD countries, but not for developing countries.

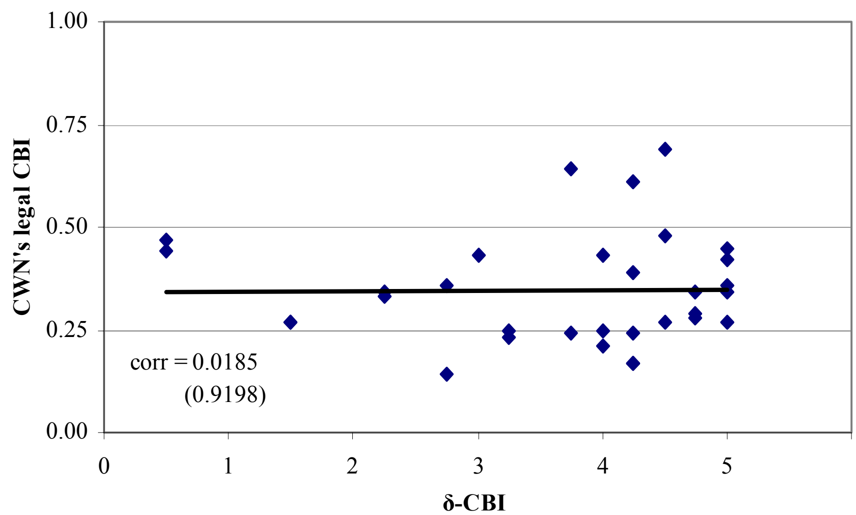

Figure 2: Relationship between δ-CBI and CWN's Legal CBI Index

In addition, considering the CBI indices based on the turnover rate of central bank governors by Cuckierman, Webb and Neyapti (1992) and Sturn and Haan (2001), respectively CWN and SH, Table 11 shows that only the former has the expected negative correlation with the $\delta-$based CBI index. This may be explained by the fact that the SH index, unlike the CWN index, does not cover the same time sample used in our estimations of δ. Figure 3 shows the negative relationship between the $\delta-$based CBI index and CWN's CBI index based on the turnover rate. The fact that the negative correlation is highly significant suggests that the turnover rate may better capture *de facto* CBI, since it correlates well with a measure that is data-dependent, such as the $\delta-$based CBI index.

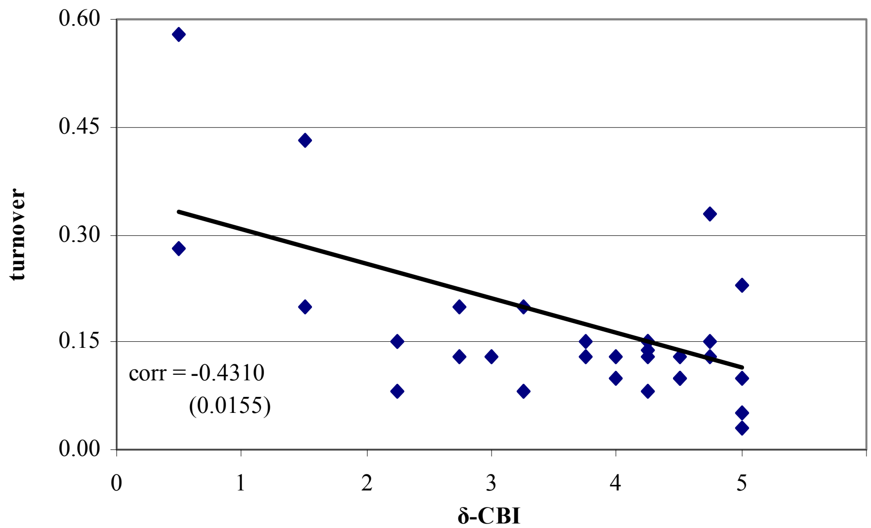

Figure 3: Relationship between δ-CBI and CWN's Turnover Rate

Finally, using the actual data on M, B and c (real consumption), and the country-specific parameters estimated from the model, predictions for the average rate of inflation can be constructed. Figures 4(a) and 4(b) show that the model can approximate reasonably well the inflation rates observed in the data.

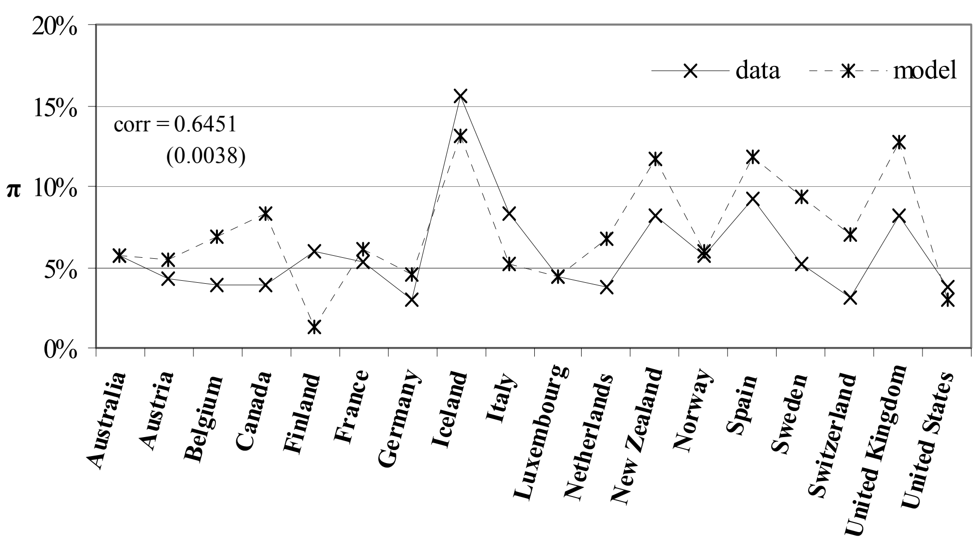

Figure 4(a): Inflation in OECD Countries: Model vs. Data

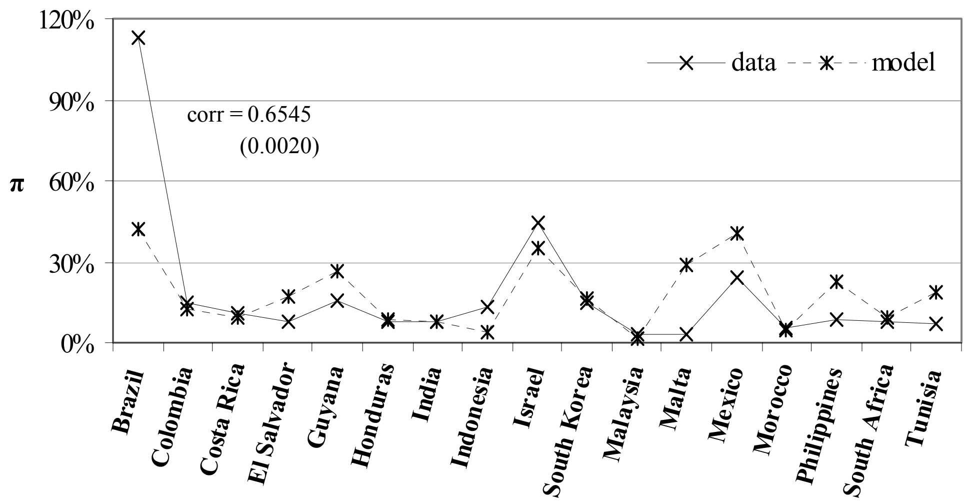

Figure 4(b): Inflation in Developing Countries: Model vs. Data

4 Conclusions

This paper uses a simple infinite-horizon monetary economy to study how fiscal and monetary policy interact to determine the aggregate price level. The government behavior is summarized by a long-run fiscal policy rule, where a fraction of the outstanding debt is backed by the present discounted value of current and future primary surpluses. The remaining debt is backed by the present discounted value of current and future seigniorage revenue. Economies may thus be indexed by the fraction of the debt backed by the fiscal authority. Only when the degree of fiscal dominance is zero, and the debt is fully backed by fiscal policy, is the price level determined by the stock of money alone. More generally, the proportion of debt backed by money behaves like money itself for the purpose of determining the price level.

Simple unit root econometrics techniques are employed to identify the parameter that indexes the policy regimes from the long-run dynamics of nominal money stock, consumption, and government debt. Results suggest that (*i*) a fiscal/monetary regime with a low degree fiscal dominance is a reasonable approximation for most OECD economies and for some developing countries, (*ii*) on average, developing countries have a higher degree of fiscal dominance than OECD countries, and (*iii*) fiscal dominance is more frequent among developing countries than in developed economies.

In addition, it is also shown that the estimates of the parameter that determines the degree of fiscal dominance/central bank independence correlate positively with some institutional measures of central bank independence, especially those based on *de facto*, rather than *de jure*, or legal, autonomy of central banks from the fiscal authority.

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