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DISTORTIONARY TAXATION AND LABOR SUPPLY : EVIDENCE FROM CANADA

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RÉSUMÉ

Cet article examine de façon empirique les effets des taxes distorsionnaires sur l'offre de travail dans le cadre d'un modèle d'équilibre général. Les relations de long terme impliquées par le modèle sont dérivées et testées en utilisant des données canadiennes entre 1966 et 1993. Alors que les relations de coïntégration prédites par le modèle sans taxe distorsionnaire sont rejetées par les données, celles avec taxe ne le sont pas. Les accroissements persistants du taux de taxation du revenu de travail semblent jouer un rôle important dans la tendance décroissante observée des heures travaillées.

Mots clés : taxation du travail, offre de travail, coïntégration, Canada

ABSTRACT

This paper examines empirically the effects of distortionary taxation on labor supply using a general equilibrium framework. The long-term relations predicted by the model are derived and tested using Canadian data between 1966 and 1993. While the cointegrating predictions of the model without taxation are rejected, the ones of the model with labor taxation are not. Persistent labor tax rate increases appear to play an important role in the observed downward trend in hours worked.

Key words : labor taxation, labor supply, cointegration, Canada

1 Introduction

This paper examines empirically the effects of distortionary taxation on labor supply. The analysis is carried out using a general equilibrium model where public consumption can act as partial substitute of private consumption and taxes are paid on labor income. At the theoretical level, it is clear that because utility maximization equates the marginal rate of substitution of consumption and leisure to the real wage, labor taxation alters the agents' choice by reducing their take-home real wage. Since the government's intertemporal substitution of debt for (distortionary) taxes can affect the agents' consumption and labor supply decisions, Ricardian equivalence fails [see Barro (1989), Trostel (1993), and Cardia (1997)]. Although some researchers suggest that, for a given level of fiscal spending, distortionary taxation has only second-order implications [see, among others, Barro (1989)], the question of whether the magnitude of the effects just described is economically and statistically important is primarily an empirical question.

That distortionary taxation could affect labor supply is suggested by figure 1 that plots the effective labor tax rate and the number of hours worked per person per week in Canada between 1966 and 1993.[1] Notice that the upward trend in the tax rate is mirrored (almost literally) by a downward trend in hours worked. While other causes, like sustained technological progress and demographic shifts, could also account for the reduction in the number of hours worked, figure 1 is certainly provocative and motivates the inquiry of this paper.

More generally, the analysis of distortionary taxation is important for several reasons. First, alternative methods of public financing are associated with different levels of social welfare. Cooley and Hansen (1992) quantify the costs of various forms of taxation and find that replacing labor-income with lump-sum taxes reduces the welfare loss by 5.2% of GNP compared with a benchmark model that represents current US tax policy. Ohanian (1997) compares war-financing by debt or taxes and concludes that would have World War II been financed solely with distorting capital and labor taxes, the welfare loss would have been approximately 2-percent of steady-state

[1]The tax rates for this graph were taken from Mendoza, Razin, and Tesar (1994) and Ruggeri, Laroche, and Vincent (1997). The series of hours worked was constructed by the authors using data on total hours worked per week and labor force (in persons) supplied by Statistics Canada.

GDP. Second, the effect of government purchases might depend on whether they are financed by means of lump-sum or distortionary taxes. Baxter and King (1993) find that the government spending multiplier is positive when expenditure is financed by lump-sum taxes but negative when financed by distortionary taxes. Finally, taxes can be an important source of economic disturbances and/or play an important role in the transmission of shocks. For example, McGrattan (1994) finds that 27% (4%) of the variance of output is explained by innovations to the labor (capital) tax rate.

In related research, Stuart (1981) constructs a two-sector model where labor taxes are paid on income earned in the market sector. The calibration of the model to the Swedish economy indicates that increasing the marginal tax rate from 58% to 65% reduces labor supplied to the market sector by between 1.8% and 2.5% depending on the scenario considered. Braun (1994) shows that introducing distortionary taxes in a real business cycle (RBC) model improves its ability to reproduce features of the US economy like the variability of hours worked and the weak correlation between real wages and employment. Braun suggests that substantial intertemporal substitution effects in the labor supply decision may be the result of changes in taxes.

The papers above employ as analytical tool different versions of the neoclassical growth model and provide evidence primarily in the form of simulations of the model economy. Rather than using calibration to asses the relevance of taxation on labor supply, this paper derives the empirical predictions of the model and, based on unit-roots tests, frames these predictions in terms of cointegration. It is shown that the behavioral rules and resource constraints of the model imply that a precise combination of the variables should be stationary. Because our model includes as special cases specifications without distortionary taxation and no substitutability of public spending and private consumption, cointegration tests provide a simple and transparent strategy to evaluate the competing theories. The estimate of the cointegrating vector delivers estimates of the preference parameters, and impulse-response analysis allows us to examine the dynamic response of the leisure-labor ratio to innovations in the effective tax rate. Although our methodology and data set differ from Braun's, we validate his conclusion that distortionary taxation can be an important determinant of the agents' labor supply decision.

The econometric strategy builds on previous work by Ahmed and Yoo (1995) who study the effect of fiscal trends on the time-series predictions of

RBC models. Ahmed and Yoo show that although the US consumption-to-output and leisure-labor ratios appear nonstationary, including government expenditure yields a relation among the variables that is stationary. This paper generalizes their specification by introducing distortionary labor taxes and confirms, using Canadian data, the rejection of a basic neoclassical growth model that ignores the effect of fiscal variables.

Most of the empirical macroeconomics literature on the effects of taxation examines only indirectly the relevance of distortionary taxes. For example, reduced-form estimation designed to capture the effects of government debt and taxation on private consumption yields conflicting results [on this see Cardia (1997)]. Even in the cases when Ricardian equivalence is rejected, it is not possible to distinguish between the possible sources of the failure (whether finite horizons, liquidity constraints, or distortionary taxation). Empirical analysis is also complicated by the fact that labor income tax rates depend on the household's income bracket and generating aggregate tax series is nontrivial. However, in a recent paper, Mendoza, Razin and Tesar (1994) compute effective tax rates consistent with the tax distortions faced by a representative agent in a general equilibrium model and show that their time series properties are similar to other tax measures that employ data on income distribution, statutory taxes, and other institutional characteristics.

Our results show that changes in the labor tax rate affect the leisure/labor supply decision in a manner consistent with the theoretical model. While the cointegrating predictions of the specification with distortionary taxation are not rejected by the data , they are rejected for the simple neoclassical growth model without taxes. The parameter estimate for the leisure preference parameter in the utility function is statistically different from zero, and its magnitude is comparable to previous estimates based on US data and the values generally used in calibration by the RBC literature. When the predicted values for the leisure-to-employment ratio are plotted against the actual series, the fitted values and the actual series are remarkably close.

Impulse-response analysis shows that an increase in labor taxation decreases labor supply. Roughly speaking, an increase of a 1 percentage point in the labor tax rate decreases weekly hours worked (or equivalently, increases leisure) by 0.3 hours. Given the upward trend in the labor tax rate between 1966 and 1993 (when this rate rose from 15.1% to 31%), distortionary taxation appears to explain a substantial part of the decrease in hours worked per person in Canada documented in figure 1. While these results do no constitute direct evidence against Ricardian Equivalence, they show that the weight of leisure in the utility function is at least as large as the one of consumption[2] and that distortionary taxation can affect labor supply. Hence the intertemporal substitution of debt for taxes might produce real effects.

The paper is organized as follows: section 2 presents and solves the general equilibrium model; section 3 analyses the univariate properties of the series, defines the model predictions in terms of cointegrating relations, tests the implications of the competing specifications, and obtains estimates of the structural parameters; section 4 employs impulse-response analysis to examine the short- and long-term effect of changes in taxes on the leisure-labor decision; and section 5 concludes.

2 The Model

2.1 Preferences and Utility Maximization

The economy is populated by identical, infinitely-lived agents who choose optimal sequences of consumption and leisure to maximize their life-time utility. Formally, the representative agent's problem is,

$$\underset{\{C_{t+i}, L_{t+i}\}_{i=0}^{\infty}}{Max} \quad E_t \sum_{i=0}^{\infty} \beta^i u(C^*_{t+i}, L_{t+i}) \tag{1}$$

where $\beta \in (0, 1)$ is the subjective discount factor, $u(\cdot)$ is the instantaneous utility function assumed concave and strictly increasing in both of its arguments, L_t is leisure, and C^*_t is effective consumption. Effective consumption is a composite of private and public consumption:

$$C^*_t = C_t + \theta G_{c,t}, \tag{2}$$

where C_t is private consumption, $G_{c,t}$ is government consumption, and the coefficient θ measures the contribution of $G_{c,t}$ to the agent's well-being. Notice that under this specification, a unit of government consumption is equivalent

[2] This is important because if leisure is not a quantitatively important component of utility, labor taxation does not distort the agents' labor supply decision and all taxation is effectively lump-sum.

to θ units of private consumption in utility terms [see, among others, Barro (1981), Aschauer (1985) and Ahmed and Yoo (1995)].[3]

In the rest of the analysis we specialize instantaneous utility to the logarithmic form [see also Baxter and King (1993) and Braun (1994)]:

$$u(C_t^*, L_t) = \log(C_t^*) + \gamma \log(L_t), \tag{3}$$

where γ is a positive constant that measures the relative weight of leisure in $u(\cdot)$.[4] The agent's budget constraint is given by

$$A_{t+1} = (1 + r_t)A_t + (1 - \tau_t)w_t N_t - T_t - C_t,$$

where A_t is financial wealth, r_t is the real interest rate, τ_t is the tax rate on labor income, and T_t is a lump-sum tax (net of transfers).[5] In equilibrium, financial wealth is held only in the form of private capital. That is, $A_t = K_t$, where K_t denotes private capital. Finally, the total time endowment is normalized to 1 so that

$$L_t + N_t = 1. \tag{4}$$

In addition to the transversality condition, the first-order conditions that characterize the solution of the dynamic programing problem above are

$$1/C_t^* = \beta(1 + r_t)E_t\left(1/C_{t+1}^*\right) \tag{5}$$

and

$$\gamma C_t^*/L_t = (1 - \tau_t)w_t. \tag{6}$$

[3]Baxter and King (1993) also include an additive function, say $\Gamma(G_{c,t})$, in the instantaneous utility to capture the notion that certain government purchases (for example, military expenditures) increase the agent's welfare without affecting her consumption/leisure decision.

[4]Another specification of the utility function assumes that it is linear on leisure [see, for example, Hansen and Cooley (1992) and Ohanian (1997)]. Hansen's (1985) model where households can work only a fixed number of hours or none at all (*i.e.,* labor is indivisible) also yields a linear dependence of instantaneous utility on leisure. These specifications imply that the marginal disutility of work is constant, the intertemporal elasticity of substitution is infinity, and agents readily substitute labor across periods. Hence, under linear utility, changes in labor taxes are likely to have a larger effect on labor supply than under logarithmic preferences. For some evidence in favor of logarithmically separable utility see McGrattan (1994).

[5]Because our focus is on the effect of labor taxes on labor supply, we abstain from explicitly incorporating capital and consumption taxes.

Equation (5) describes the optimal rate of substitution between current and future consumption while (6) dictates that the marginal rate of substitution between leisure and consumption should equal the after-tax real wage.

Relations (5) and (6) imply that the intertemporal substitution of leisure follows

$$1/[(1-\tau_t)w_t L_t] = \beta(1+r_t)E_t[1/((1-\tau_{t+1})w_{t+1}L_{t+1})].$$

In order to illustrate the implications of the above relation, it is useful to consider the case of perfect-foresight (so that the expectations operator drops), take logs and rearrange the resulting expression to obtain

$$\Delta \log(L_{t+1}) = \log(\beta(1+r_t)) + (\tau_{t+1} - \tau_t) - \Delta \log(w_{t+1}),$$

where $\Delta \log(L_{t+1}) = \log(L_{t+1}) - \log(L_t)$, $\Delta \log(w_{t+1}) = \log(w_{t+1}) - \log(w_t)$, and we have approximated $\log(1-\tau_t)$ by $-\tau_t$. Hence, an increase in the real interest rate and a transitory decrease (increase) in labor tax rate (wage) at time t, constitute incentives for the intertemporal substitution of leisure for labor. In all these instances, agents reduce current leisure *vis a vis* future leisure, or equivalently, increase their labor supply at time t compared with time $t+1$. The prediction that the agents' labor supply reacts not only to movements in the real wages and interest rates, but also to changes in the labor tax rate, is important for two reasons. First, as pointed out by Braun (1994), it can explain the weak correlation observed in aggregate data between hours worked and real wage. Second, as recognized by Barro (1989), it means that distortionary taxation could lead to the empirical failure of Ricardian Equivalence.

Notice, however, that the intertemporal effects of movements in the tax rate and real wage do depend on the time-series properties of the variables. If all tax/wage changes are perceived by agents to be largely permanent, then there is no reason to substitute intertemporally labor supply. An illustrative example is the case when the log of the real wage and the tax rate follow random walks. Then, their first-difference are white noise and the changes in (the log of) leisure are also white noise around the time-varying component $\log[\beta(1+r_t)]$.

On the other hand, regardless of the persistence associated with tax changes, labor taxation has a level, intratemporal effect as seen from the

Euler condition (6). Thus, increases in the labor tax rate, reduce the take-home real wage and motivate agents to substitute consumption for leisure within the period. As we will see below, this could partly explain labor supply changes in Canada during the post-war period when labor tax movements have been persistent [see figure 1].

2.2 Production and Public Sectors

Production of the (single) consumption good is carried out by perfectly competitive firms using a constant-returns-to-scale technology of the form

$$Q_t = a_t K_t^{1-\phi} N_t^{\phi}, \tag{7}$$

where Q_t is output, K_t is private capital stock, N_t is labor (measured in hours worked), ϕ is a constant coefficient that satisfies $0 < \phi < 1$, and a_t is an exogenous productivity shock. As shown below, the equilibrium condition examined empirically in section 3 is robust to allowing government expenditure and/or public capital to increase factor productivity [as in Baxter and King (1993) and Ahmed and Yoo (1995)] and to the precise time-series specification of the technology shock.

The representative firm chooses labor demand and the level of capital to maximize profits:

$$\pi_t = Q_t - r_t K_t - w_t N_t - \delta K_t, \tag{8}$$

where the price of the good has been normalized to 1, r_t is the rental price of capital, w_t is the real wage, and δ is the rate of depreciation. As usual, necessary conditions for the maximization of (8) subject to (7) are

$$a_t \phi \left(K_t/N_t\right)^{1-\phi} = w_t, \tag{9}$$

and

$$a_t(1-\phi)\left(K_t/N_t\right)^{-\phi} - \delta = r_t. \tag{10}$$

The first condition simply states that the firm's demand for labor equates the marginal productivity of labor and the real wage while the second condition determines the optimal level of capital as a function of its rental price. In turn, private capital accumulates according to

$$K_t = (1-\delta)K_{t-1} + I_t.$$

Following the literature [see, for example, Cooley and Hansen (1992), Braun (1994), and McGrattan (1994)], we abstract from assigning a utility function to the government and instead model it as an exogenous sequence of expenditures and taxes. This sequence satisfies in every period the government's budget constraint:

$$G_{c,t} + G_{i,t} = \tau_t w_t N_t + T_t, \tag{11}$$

where $G_{i,t}$ is government investment.[6]

2.3 Equilibrium Conditions

Replacing the factor prices that solve the firm's profit maximization, (9) and (10), into the agent's Euler equations yield the arbitrage relations

$$1/C_t^* = \beta[1 + a_t(1-\phi)\,(K_t/N_t)^{-\phi} - \delta]E_t\left(1/C_{t+1}^*\right) \tag{12}$$

and

$$\gamma C_t^*/L_t = (1-\tau_t)a_t\phi(K_t/N_t)^{1-\phi}. \tag{13}$$

Finally, the goods-market equilibrium condition for the economy is[7]

$$Q_t = C_t + I_t + G_{c,t} + G_{i,t},$$

or, dividing both sides by output, Q_t,

$$1 = c_t + i_t + g_{c,t} + g_{i,t}, \tag{14}$$

where the lower-case letters denote ratios-to-GDP.

Equation (12) describes the optimal consumption smoothing behavior on the part of agents and has been extensively examined, both in partial and general equilibrium, in previous literature. More interesting, from the perspective of examining the effects of taxation, is relation (13) that describes the intratemporal substitution of consumption and leisure. Plugging (2)

[6]Since allowing government debt does not alter the intratemporal empirical predictions examined below, we simplify the formulation of the model by assuming that expenditures equal receipts in every period.

[7]This condition implicitly treats the economy as closed. However, as shown in the following section, econometric results are robust to modeling the economy as open.

and (4) into (13), dividing by Q_t, using (7), and imposing the equilibrium condition (14) yield

$$(1-\tau_t)x_t = (\gamma/\phi)\left[1-(1-\theta)g_{c,t}-g_{i,t}-i_t\right], \tag{15}$$

where the term

$$x_t = (1-N_t)/N_t,$$

is the leisure-labor ratio. Equation (15) encapsulates the optimizing choices of agents and firms and the economy's resource constraint. It also relates at the theoretical level, fiscal policy to the agents' labor supply decision and (as we will see below) generates empirical predictions that can be readily compared with actual data. Its tractable form is the deliberate result of using a logarithmic specification for the utility function and a constant-returns-to-scale technology. Notice that by examining the variables as percentages of GDP, any trend effect associated with population growth is eliminated. Also, multiplicative increases in factor productivity (whether due to technology shocks, public capital, or government consumption) cancel out in the derivation of (15) and, as a result, have no bearing on its empirical predictions. This is important because it means that the econometric results below are robust to specific assumptions regarding the time-series properties of productivity shocks.

The above model includes two interesting specifications as special cases. First, by setting $\tau_t = 0$ for all t one obtains a model with no distortionary labor taxation. Because an important goal of this paper is to examine empirically the effects of taxation on labor supply, this restricted model constitutes a natural alternative. It is easy to see that the equivalent expression of (15) for this restricted specification is

$$x_t = (\gamma/\phi)\left[1-(1-\theta)g_{c,t}-g_{i,t}-i_t\right]. \tag{16}$$

Second, one could consider the case where government consumption is not a substitute of private consumption by setting $\theta = 0$ in (15) to obtain

$$(1-\tau_t)x_t = (\gamma/\phi)\left[1-g_{c,t}-g_{i,t}-i_t\right]. \tag{17}$$

3 Econometric Analysis

3.1 The Data

The data set consists of 112 quarterly, seasonally-adjusted observations of output, consumption, investment, government consumption (that includes both durables and nondurables), government investment and the leisure-labor ratio for Canada. All variables, except for the leisure-labor ratio, are in constant 1986 Canadian dollars. The sample period from 1966:1 to 1993:4 was determined by the availability of data on effective tax rates. These rates were computed by Mendoza, Razin, and Tesar (1994) for 1965-1988, and later extended by Ruggeri, Laroche, and Vincent (1997) until 1993. In addition, because Statistics Canada collects the series of hours worked on a quarterly basis only since 1966, the year 1965 could not be included in the analysis.

For the calculation of the leisure-labor ratio, we followed Ahmed and Yoo (1995). The total endowment of time available to the economy is defined as $EE = 12 * 7 * 16 * LF$, where LF is the labor force (in thousand of persons),[8] 12 is the number of weeks in a quarter, 7 is the number of days in a week, and 16 is the number of (non-sleeping) hours in the day. The leisure-labor ratio is then computed as $x = (EE - TN)/TN$ with TN denoting the total number of hours worked. The tax-adjusted leisure-labor ratio, $(1 - \tau_t)x_t$, is calculated assuming that the annual tax rate applies equally in all four quarters of the year.[9]

3.2 Univariate Properties of the Series

Prior to the econometric analysis of the equilibrium relations derived in section 2, we examine the time-series properties of the data. Graphical analysis appears to indicate that the series could be characterized by nonstationary

[8] The Labor Force Survey is designed to represent all persons in the population 15 years and over with the exception of persons living on Indian reserves, full-time members of the armed forces and people living in institutions for more than six months. It is based on a sample of 58000 representative households across the country (but excluding the Yukon and the Northwest Territories), involving some 103000 respondents.

[9] We also considered the alternative approach of interpolating the available annual tax rates to construct quarterly estimates. However, results were basically the same as the ones reported below.

processes. Consider first figure 2 that contains the plots of consumption, investment, government consumption, and government investment as percent of GDP and the leisure-labor ratio (both adjusted and unadjusted for taxes). These graphs suggest that the variables are very persistent and likely to be integrated of order 1. This intuition is statistically confirmed by the results of KPSS and Augmented Dickey-Fuller (ADF) tests.

For most of the variables under consideration, the natural specification to estimate is that of an autoregression with constant term but no trend [see Hamilton (1994, ch. 17)].[10] The choice of the optimal level of augmentation, (*i.e.,* the number of lagged first differences included in the OLS regression) was based on the recursive application of t-tests as suggested in Campbell and Perron (1991). In order to assess the robustness of the results, we also employed the Modified Information Criterion (MIC) [Ng and Perron (1998)] and included a deterministic time trend in the regression with similar conclusions to the ones reported. The KPPS procedure tests the null hypothesis of stationarity and requires the construction of an estimate of the variance of the disturbance in the regression that takes into account its autocorrelation. Kwiatkowski, Phillips, Schmidt, and Shin (1992) propose the Newey-West estimator and, given the trade-off between size distortion and power found in their simulations, suggest the use of 8 lags for the Barlett kernel.

Note in table 1 that the null hypothesis of stationarity is rejected by the KPSS test for all variables except for the tax-adjusted leisure-labor ratio. However, in this case the non rejection is rather marginal (the statistic is 0.28 and the 10% critical value is 0.347) and is not robust to including a time trend in the regression.[11] On the other hand, ADF test results imply that the null hypothesis of a unit root cannot be rejected for any variable at standard significance levels. Similar results are reported by Ahmed and Yoo (1995) using US data for the somewhat-larger sample period 1955:1 to 1993:1. Ahmed and Yoo find that the leisure-labor ratio and the ratios of

[10] Since most of the variables are ratios to GDP it could be argued that perhaps a nonlinear alternative would be even more plausible. However, since standard unit root tests are derived under the assumption of linearity, this proposition is not pursued further. Campbell and Perron (1991, p. 157) stress that, in finite samples, any covariance-stationary process can be approximated arbitrarily well by a unit root process, in the sense that the autocovariance structures will be arbitrarily close.

[11] The test stastic obtained when adding a trend is 0.21, that is larger than the 5% critical value of 0.146.

consumption and various government expenditure components to output are nonstationary, but the investment to output ratio is stationary.[12]

In addition, we also test whether the effective labor-tax rate for Canada is persistent enough to be described by a nonstationary time-series process. To that effect we carry out ADF and KPSS tests using raw annual data from Mendoza, Razin, and Tesar (1994) and Ruggeri, Laroche, and Vincent (1997). The estimated regressions are stationary processes without trend and, as before, the level of augmentation of the ADF test is determined using recursive t-tests. Results are presented in the last row of table 1 and indicate that while the hypothesis of a unit root cannot be rejected at standard significance levels, the hypothesis of stationarity is rejected at the 1% level. (This conclusion is robust to including a time trend in the test regressions.) This finding is in line with Mendoza, Razin, and Tesar's observation that the tax rate on labor income has followed an upward trend for all countries in their sample. The evidence in table 1 simply suggests that this trend appears to be stochastic rather than deterministic.

3.3 Tests of the Empirical Implications

The above results are important because they allow us to frame empirically the equilibrium conditions derived in section 2 in terms of cointegrating relations [Engle and Granger (1987)]. That is, even if individual variables are characterized by nonstationary processes, the behavioral rules and resource constraints that underlie (15), (16), and (17) imply that a precise combination of these variables should be stationary. Because these equations are based on different assumptions about the relevance of distortionary taxation and the substitutability between public and private consumption, cointegration tests provide a simple and transparent strategy to evaluate the competing models.

Two types of cointegration tests are employed. First, the null hypothesis of no cointegration is tested using the residual-based approach proposed by Engle and Granger (1987). Gonzalo and Lee (1998) show that this test is

[12] King *et al.* (1991) present some empirical evidence that, for the United States, the consumption/output and investment/output ratios are stationary. The discrepancy in the results could be partly attributed to the fact that King *et al.* measure output as private output alone (since there is no government in their model) while Ahmed and Yoo (1995) and this paper use GDP.

more robust than Johansen's trace test [Johansen (1991)] to certain departures from unit root behavior like long memory and stochastic unit roots. Still, as an additional check, we use Johansen's Maximum Likelihood (ML) procedure to determine the number of cointegrating relations among the model variables.

Engle and Granger's test evaluates the null hypothesis of no cointegration and 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. Results for relations (15), (16), and (17) are respectively presented in the three top rows in table 2. Recall that (15) includes both taxation and partial substitution of public and private consumption, (16) ignores distortionary taxation by assuming that the labor tax rate is zero in all periods, and (17) corresponds to the case of no substitutability between public and private consumption obtained when $\theta = 0$. The restriction that the coefficients on government consumption, government investment, and investment are the same in (17) is imposed by running the OLS regression of the tax-adjusted leisure-labor ratio on a constant and the sum of the three variables. The test statistics in column 3 indicate that the null hypothesis of no cointegration is rejected at the 5% level for (15) (the p-value is approximately 3%), and at the 10% level for (17) (in this case the p-value is approximately 9.6%). In contrast, the same hypothesis cannot be rejected for the relation without taxes (16) at any standard significance level.

These results constitute evidence in favor of the model with distortionary taxation in that (*i*) its econometric prediction – that the model variables form a stable long-run relation – is supported by the data, and (*ii*) the same prediction by the alternative specification without taxation is rejected. This conclusion is robust to different time-series specifications of the technology shocks, to relaxing the assumption that public expenditure and private consumption are partial substitutes, and to modeling the economy as open.[13] Hence, labor taxation appears to be an empirically important element in models concerned with the agents' labor supply decision. The nonnegligible difference in the p-values associated with the rejection of the hypothesis of

[13] In this case, the trade balance (as percentage of GDP) also appears in the right-hand-side of the cointegrating relations. Test statistics for the null hypothesis of no cointegration in the open-economy model with (without) distortionary taxation are -4.53 (-2.57), respectively.

no cointegration of (17) and (15), suggests that the partial substitutability of public spending and private consumption reinforces the finding that (15) is a cointegrating relation.

In addition to the above test, we also employ the trace test [Johansen (1991)] to verify the number of cointegrating relations among the variables. The number of lags to be included in the Vector Error Correction (VEC) model were chosen using a sequence of Likelihood Ratio tests in a vector autoregression in levels as suggested by Enders (1995). Results for this test are presented in table 3 and indicate the presence of two cointegrating relations at the 1% significance level (though 3 at the 5% level) for the model with distortionary taxation and none (1 at the 5% level) for the model without taxes. Although, Gonzalo and Lee (1998) show that in certain circumstances Johansen's test tends to overestimate the number of cointegrating relations, it seems worthwhile to explore the possibility of a second cointegrating relation among the variables. A natural candidate is a cointegrating relation between the two components of government expenditure. A residual based test [see the fourth row of table 2] suggested that indeed the government consumption and investment are cointegrated.[14]

3.4 Estimates of the Structural Parameters

The estimation of the cointegrating relation (15) is of particular interest because it provides us with estimates of the structural parameters of the model. A number of strategies to estimate cointegrating vectors (some of them asymptotically equivalent) have been proposed in the literature. A nonexhaustive list includes OLS [Engle and Granger (1987)], nonlinear least squares [Stock (1987)], canonical correlations [Bossaerts (1988)], maximum likelihood in a fully specified VEC model [Johansen (1991)], three-step-estimation [Engle and Yoo (1989)] and dynamic generalized least squares

[14]The result reported in table 2 corresponds to the one obtained regressing government investment on a constant term and government consumption. Notice however that for the inverse regression (that is government consumption on a constant and government investment), the hypothesis of no cointegration cannot be rejected at the 5% significance level. This result highlights the well-known fact that normalizations can play a nontrivial role in cointegration testing [see Hamilton (1994, ch. 19)]. Additional evidence on the cointegration of the expenditure components is provided by Johansen's trace statistic that rejects the null hypothesis of no cointegrating vector between g_c and g_i at the 5% level.

(DGLS) [Stock and Watson (1993)]. Gonzalo (1994) uses Monte Carlo simulations to compare some, but not all, the above methods[15] and concludes that in finite samples the maximum likelihood method has the smallest variance among the estimators considered. On the other hand, this approach has the disadvantage that it only delivers the basis of the cointegrating vectors rather than the cointegrating relations themselves. Phillips (1991) stresses that if researchers want to make structural interpretations on the separate cointegrating relations, this logically requires the use of restrictions from economic theory.

With the above considerations in mind, we employ the DGLS method proposed by Stock and Watson (1993) that is asymptotically equivalent to maximum likelihood [see Gonzalo (1994, p. 204)] but makes use of the restrictions of the general-equilibrium model. This approach involves running the OLS regression

$$\begin{aligned}(1-\tau)x_t = \; & \alpha + \rho_1 g_{c,t} + \rho_2 g_{i,t} + \rho_3 i_t + \sum_{s=-p}^{p} \xi_{1,s}\Delta g_{c,t-s} \\ & + \sum_{s=-p}^{p} \xi_{2,\mathrm{s}}\Delta g_{i,t-s} + \sum_{s=-p}^{p} \xi_{3,s}\Delta i_{t-s} + u_t\end{aligned} \tag{18}$$

where α is an intercept, ρ_j and $\xi_{j,s}$ for $j = 1, 2, 3$ denote constant coefficients, and u_t is a disturbance term. The serial correlation of the residuals (if any) is then characterized in a parametric time-series model and the equation is reestimated using GLS. We selected the appropriate number of leads and lags by the sequential application of F-tests starting with the maximum number $p = 4$. Test results indicated that the most parsimonious yet statistically accurate representation involved setting $p = 0$, so that only the lagged first-difference of the variables were included in (18). The residuals of the OLS regression were parameterized as an AR(1) process based on a regression of $\hat{u}_t$ on four of its lags but no constant.[16] Results for the GLS regression are presented in table 4.[17]

[15]The author explains (see p. 204) that some estimators, most notably the dynamic GLS procedure by Stock and Watson, were proposed after his article was submitted.

[16]The coefficients of the second to fourth lag were not significantly different from zero at standard levels. For the first lag the estimate was only 0.40 (0.10) but, since it is statistically different from zero, efficiency gains are possible by using GLS.

[17]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,

Using the reduced-form parameters, it is possible to construct estimates of the structural parameters of interest. Note, however, that the share of labor to total income (ϕ) and the weight of leisure in the utility function (γ) are not separately identified because the estimated intercept corresponds to the ratio γ/ϕ. This issue can be addressed by constructing an estimate of the share of labor on the basis of national income data. Mankiw and Scarth (1995, p. 78) report that this share has been roughly constant in Canada at $\phi = 0.67$ since 1945. Multiplying the intercept estimate, $\gamma/\phi = 1.95$ (0.31), by this figure yields an estimate of $\gamma = 1.31$ (0.21), where the terms in parenthesis denote standard errors.[18] This estimate is somewhat smaller but still consistent with the ones reported by Braun (1994), who finds values ranging from 4.21 to 5.59 for different tax specifications, and McGrattan (1994) whose preference estimates imply that $\gamma = 2.95$.[19] In calibrated models, Hansen (1985) and Ohanian (1997) employ values of γ of 2 and 1.5, respectively. Thus, Canadian data appears to confirm earlier estimates of γ obtained using postwar US data and suggests that the weight of leisure in the utility function is comparable to the one assigned to consumption.

The fact that the DGLS procedure involves lags and leads of the variables complicates the structural interpretation of the remaining coefficients. However from the estimates of the level and lagged difference of government consumption and the intercept (see table 4) it is possible to recover a point estimate of θ. For example, an estimate of θ can be computed as $\theta = (\rho_1 + \xi_{1,0})/(\gamma/\phi) + 1 = (3.06 - 1.33)/1.95 + 1 = 1.89$ (0.70) where the standard error is obtained using the delta method. This estimate is positive as predicted by theory and significantly different from zero. Although this point estimate appears numerically large, it is not possible to reject the hypothesis that its true value is smaller than 1, as would be expected if government consumption were an imperfect substitute for private consumption. To see this, construct the 95% confidence interval for θ to obtain (0.52, 3.26) and note that any null hypothesis $\theta = \tilde{\theta}$ for $\tilde{\theta} \in (0.52, 0.99)$ would not be

hypothesis tests regarding the coefficients that do not have an exact unit root can be subject to size distortions.

[18]Note that the calculation of the standard error of γ implicitly assumes that the labor share is measured without error.

[19]McGrattan employs a different specification of the utility function than the one used in this paper. However a simple log transformation implies that $(1-\gamma)/\gamma$ (in her notation) corresponds to our parameter γ.

rejected at the 5% significance level.

Other estimates of θ obtained by earlier researchers using US data include Kormendi (1983) who finds $\theta = 0.28$ (0.15),[20] and Ahmed and Yoo (1995) who obtain 0.59 for durable government consumption and 0.94 for nondurable consumption. Using total government expenditure, Aschauer (1985) reports values between 0.23 and 0.42 depending on the number of lags employed in the estimation procedure, McGrattan (1994) finds -0.026 (0.126), and Katsaitis (1987) (using Canadian data) estimates values ranging from 0.35 to 0.42. Our findings are in agreement with the above estimates (except McGrattan's) and provide independent support for the idea that public spending can act as partial substitute of private consumption.

The Euler equation derived from the model determines the variables that enter (18) and, taken literally, predicts that the coefficients of their leads and lags should not be significantly different from zero. Results reported above provide some support for this implication of the model in that F-tests indicate that the coefficients of leads and lags for $p > 0$ in (18) are not statistically different from zero. For the estimated equation (with $p = 0$) only lagged government investment has a significant coefficient.

Finally, consider figure 3 that contains plots of realized and fitted values of the tax-adjusted leisure-labor ratio and notice that the model successfully tracks the behavior of the leisure-labor ratio along the business cycle.

4 The Effect of Taxation on the Leisure/Labor Ratio

While the above results suggest the empirical importance of distortionary taxation, more precise statements about the effect of distortionary taxation on labor supply are obtained in this section. To that end we linearize (15) around the steady-state, estimate the associated VEC model and perform impulse-response analysis. Consider first the linear Taylor-series expansion of (15):

$$x_t = \alpha + \lambda_1 g_{c,t} + \lambda_2 g_{i,t} + \lambda_3 i_t + \lambda_4 \tau_t + v_t, \tag{19}$$

[20] Because Kormendi's actually estimates $-\theta$ from an OLS regression of consumption on government consumption, government investment and other variables, he reports -0.28 in his article (see table 6 in p. 1006).

where α is an intercept, λ_1 through λ_4 are constant coefficients, and v_t is a random term that includes approximation error. The advantage of this linearization is apparent from (19) as taxes now affect the leisure-labor ratio additively, rather than multiplicatively, and allow the separate inclusion of τ_t in the VEC model below.

As an intermediate step, it is important to verify that the linearization has not fundamentally altered the long-run relation among the variables. The residual-based test reported in the last row of table 2 indicates that the linear version (19) still constitutes a cointegrating relation, and the Johansen's trace test reported in table 3 confirms that (as before) two cointegrating vectors are present among the variables. Thus, these results are consistent with the ones found in the previous section for the exact, nonlinear version of the model.

The VEC model (with two cointegrating vectors) is estimated by the method of FIML. Then, in order to examine the effect of an innovation in the labor tax rate on the leisure-labor ratio, we construct the response of the latter to a one-standard-deviation shock to the fiscal variable(s). In general, impulse-response analysis requires the orthogonalization of the model disturbances. This can be achieved by either imposing structural identification restrictions [for example, as in King *et al.* (1991)] or using the Choleski decomposition of the variance-covariance matrix [for example, as in McGrattan (1994)]. A possible drawback of the latter approach is that results are not usually independent of the ordering of the variables in the system. On the other hand, in the absence of a fully-specified model of government behavior that imposes constraints on the interaction of fiscal variables, the Choleski decomposition provides a general platform to analyze the data. We will see below that regardless of whether the innovations to the tax rate are exogenous or the result of changes in government expenditure (that is, regardless of the ordering), their effect on the leisure labor ratio is basically the same.

Consider first the direct effect of taxation alone on the leisure-labor ratio, obtained by considering the ordering (τ, x, g_c, g_i, i). Note that the last three variables could be interchanged without affecting the final result and only shocks to the tax rate are interpreted as exogenous. The associated impulse response is presented in figure 4 (ordering 1). Because the tax rate is persistent (persistent enough that the hypothesis of a unit root could not be rejected), the effect of a positive tax disturbance is pretty much permanent and as predicted by theory entails an increase in leisure relative to labor.

More precisely, a one-standard-deviation shock to the labor tax rate yields a long-run increase of 0.33 percentage points in the labor tax rate and an increase of 0.0106 in the leisure-labor ratio. Equivalently, an increase of 1 percentage point in the tax rate (say from 15% to 16%) raises the leisure-labor ratio by 0.032.

To give meaning to this figures, it is useful to compute the elasticity of the leisure-labor ratio with respect to the tax. However, computing elasticities requires the level of the variables in addition to their relative change. Since the variables are not stationary, no well-defined benchmark (like the mean) is available. Still, with this consideration in mind, we perform some illustrative calculations using the sample average of the variables. These values are approximately 2.3 for the leisure-labor ratio and 23% for the tax rate. For these figures, an estimate of the (long-run) elasticity of the leisure-labor ratio with respect to the labor tax rate is 0.32.[21] This implies that an increase of the tax rate from 23% to 24% would prompt agents to increase their leisure-labor ratio to $2.3+0.032 = 2.332$, or equivalently, reduce the number of hours worked per week by 0.3, from 33.9 to 33.6.[22] Using simulations of a general equilibrium model with taxation, Greenwood and Huffman (1991) show that reducing the labor tax rate from 35% to 25% increases output and hours worked by 10%. Our estimates predict that, starting from the benchmark 23%, the same reduction of 10% points in the tax rate would increase weekly hours worked by 8.8%.

In order to asses the robustness of these results, we also examine an alternative ordering of the variables. Since changes in tax rates can partly reflect changes in the exogenous path of government spending [see Burnside, Eichenbaum, and Fisher (1999)], a natural alternative is the ordering (g_c, g_i, τ, x, i). The associated impulse-responses for the leisure-labor ratio and the tax rate are presented in figure 5 (ordering 2) and show short-term paths and the long-term effects similar to the ones presented in figure 4.[23]

[21] While not directly comparable, our results are consistent with Ziliak and Kniesner (1999) that use PSID data to estimate a labor tax elasticity of -0.06.

[22] Using the definition of the leisure-labor ratio, it is possible to calculate the number of hours worked per week as $(7 * 16)/(1 + x)$, where 7 is the number of days of the week and 16 is the number of (non-sleeping) hours in the day.

[23] For example, a one-standard-deviation innovation in the tax rate would yield a permanent increase in the leisure-labor ratio of 0.0124 under ordering 2 compared with 0.0106 under ordering 1.

Hence, assuming that tax innovations are endogenous with respect to government expenditure is unlikely to drastically affect the calculations reported above.

5 Conclusions

This paper has examined the empirical relevance of fiscal policy for the agents' leisure/labor supply decision in Canada during the period 1966 to 1993. A dynamic general equilibrium model predicts that though the leisure/labor ratio might be nonstationary, a precise combination of this series and other model variables should be stationary. It is shown that this implication is rejected for the model without taxation but cannot be rejected for the model with distortionary taxation. The latter result is robust to allowing substitutability between private consumption and government spending and to the time-series properties of the productivity shock. The estimated leisure-to-labor ratio replicates well changes of the actual series during the sample period, including the large reduction of hours worked during the 1980s. Impulse-response analysis indicate that the effects of changes in the labor tax rate on the leisure/labor ratio are quantitatively important. Some back-of-the envelope calculations suggest that, starting from the benchmark 23%, a reduction of 10% points in the tax rate would increase weekly hours worked by 8.8%. Our empirical results are consistent with earlier general equilibrium models that include labor income taxation and fiscal spending, and suggest that most of the important changes in hours worked during the period considered could be explained by the explicit inclusion of fiscal policy variables and, in particular, by changes in the labor income tax rate.

Table 1. ADF and KPSS Test Results

Variable	ADF	KPSS
c	-2.41	0.57^{*}
i	-1.77	0.92^{**}
g	-1.61	0.91^{**}
g_c	-2.06	0.71^{*}
g_i	-2.22	0.94^{**}
$(1-\tau)x$	-2.06	0.28
x	-2.28	0.85^{**}
τ	-1.46	1.26^{**}

Notes: All data is quarterly except for the effective tax rates that are on an annual basis. The superscripts ** and * denote the rejection of the null hypothesis of a unit root (ADF test) or stationarity (KPSS test) at the 1% and 5% significance levels, respectively.

Table 2. Residual-Based Cointegration Test Results

Variables	k	t-Statistic
$(1-\tau)x, g_c, g_i, i$	1	-4.49^{*}
x, g_c, g_i, i	7	-2.79
$(1-\tau)x, g_c + g_i + i$	9	$-3.09^{\dagger}$
g_c, g_i	1	-3.38^{*}
$x, (1-\tau), g_c, g_i, i$	1	$-4.19^{\dagger}$

Notes: k denotes the level of augmentation of the test and was chosen using recursive t tests [see Campbell and Perron (1991)]. In all cases a constant term was included in the regression. The superscripts **, *, and † denote the rejection of the null hypothesis of no cointegration at the 1%, 5%, and 10% significance levels, respectively.

Table 3. Johansen Cointegration Test Results

Variables	k	Eigenvalue	LR Statistic	5% (1%) Critical Value	Null Hypothesis
$(1-\tau)x, g_c, g_i, i$	4	0.20	65.81	53.12 (60.16)	None**
		0.18	42.12	34.91 (41.07)	At most 1**
		0.11	21.41	19.96 (24.60)	At most 2*
		0.08	8.83	9.24 (12.97)	At most 3
x, g_c, g_i, i	7	0.20	57.07	53.12 (60.16)	None*
		0.17	33.45	34.91 (41.07)	At most 1
		0.08	14.02	19.96 (24.60)	At most 2
		0.05	5.39	9.24 (12.97)	At most 3
$(1-\tau), x, g_c, g_i, i$	4	0.23	82.18	76.07 (84.45)	None*
		0.20	53.24	53.12 (60.16)	At most 1*
		0.11	28.79	34.91 (41.07)	At most 2
		0.09	15.90	19.96 (24.60)	At most 3
		0.05	5.71	9.24 (12.97)	At most 4

Notes: k denotes the level of augmentation of the test and was chosen using the procedure suggested by Enders (1995). In all cases a constant term was included in the regression. The superscripts ** and * denote the rejection of the null hypothesis at the 1% and 5% significance levels, respectively.

Table 4. Estimates of Cointegrating Relation

Variable	Estimate	Standard Error	t-Statistic
intercept	1.95	0.31	6.22^{**}
$g_{c,t}$	3.06	1.07	2.85^{**}
$g_{i,t}$	-11.22	1.63	-6.89^{**}
i_t	-2.95	0.61	-4.87^{**}
$\Delta g_{c,t}$	-1.33	1.23	-1.08
$\Delta g_{i,t}$	17.54	4.13	4.25^{**}
Δi_t	0.97	0.60	1.61

Notes: The superscripts **, *, and † denote the rejection of the null hypothesis that the coefficient is zero at the 1%, 5%, and 10% significance levels, respectively.

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Figure 1. Hours Worked per Person per Week and Labor Tax Rate in Canada

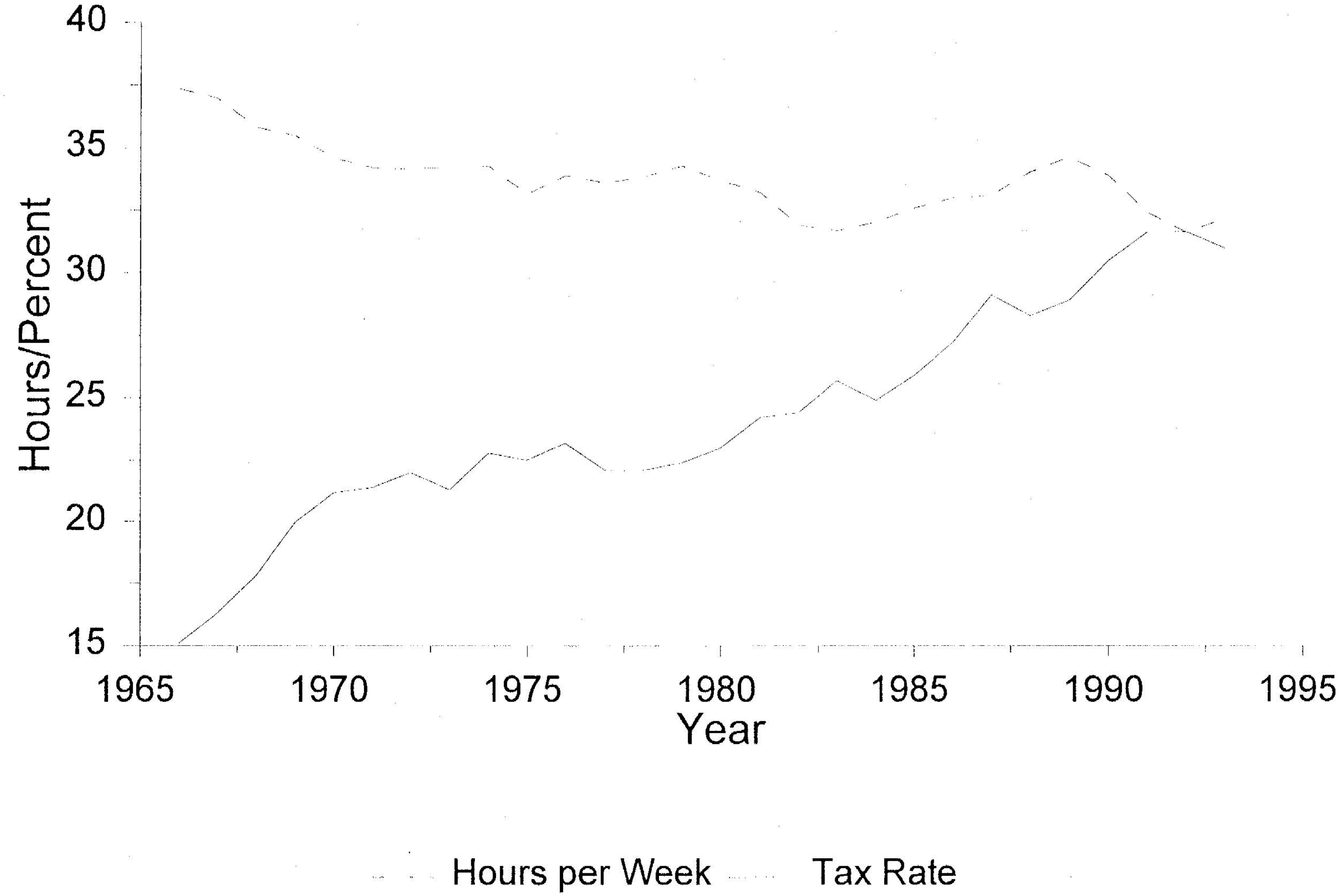


Figure 2. Aggregate Data for Canada

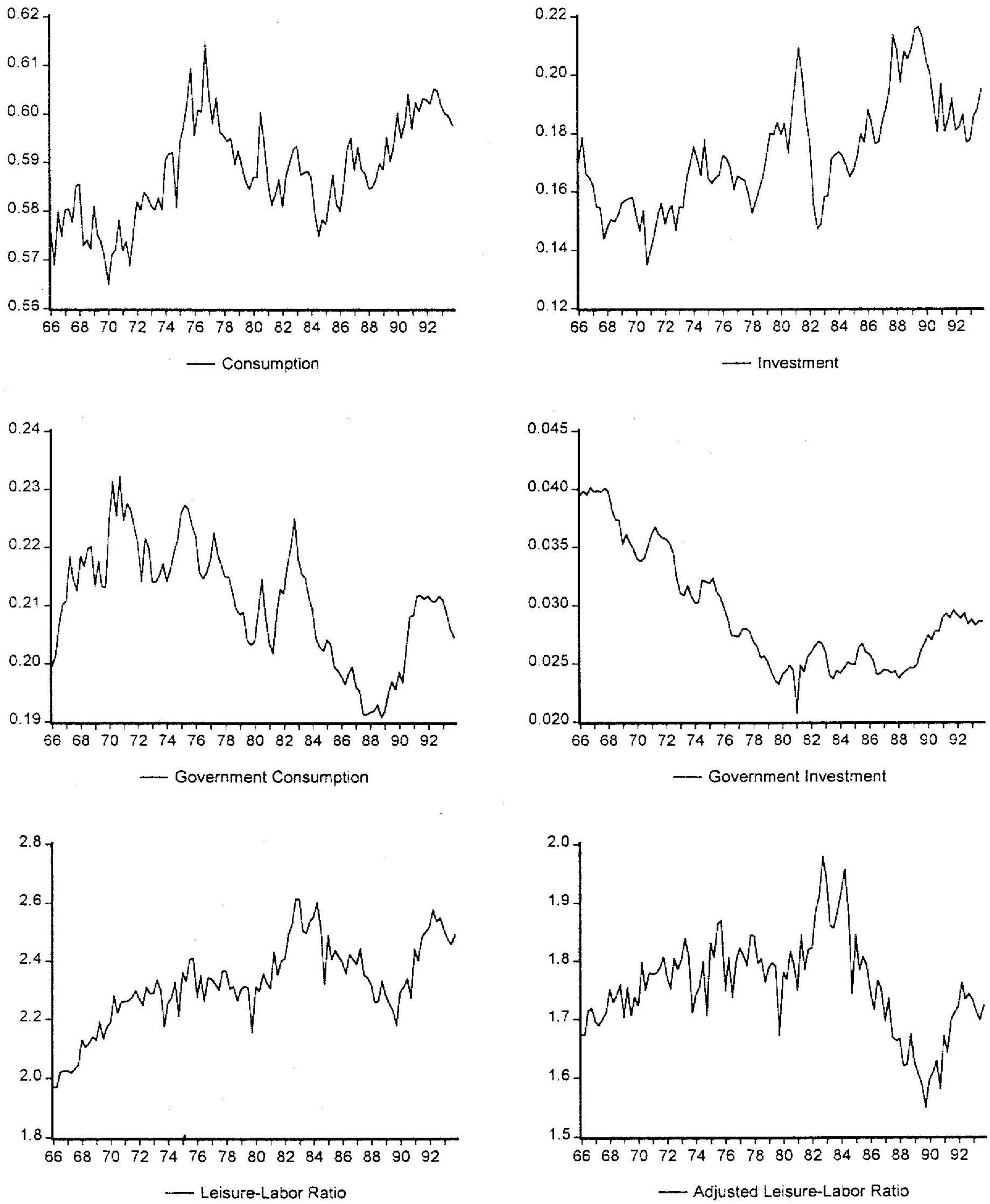


Figure 3. Realized and Fitted Values
Using DGLS

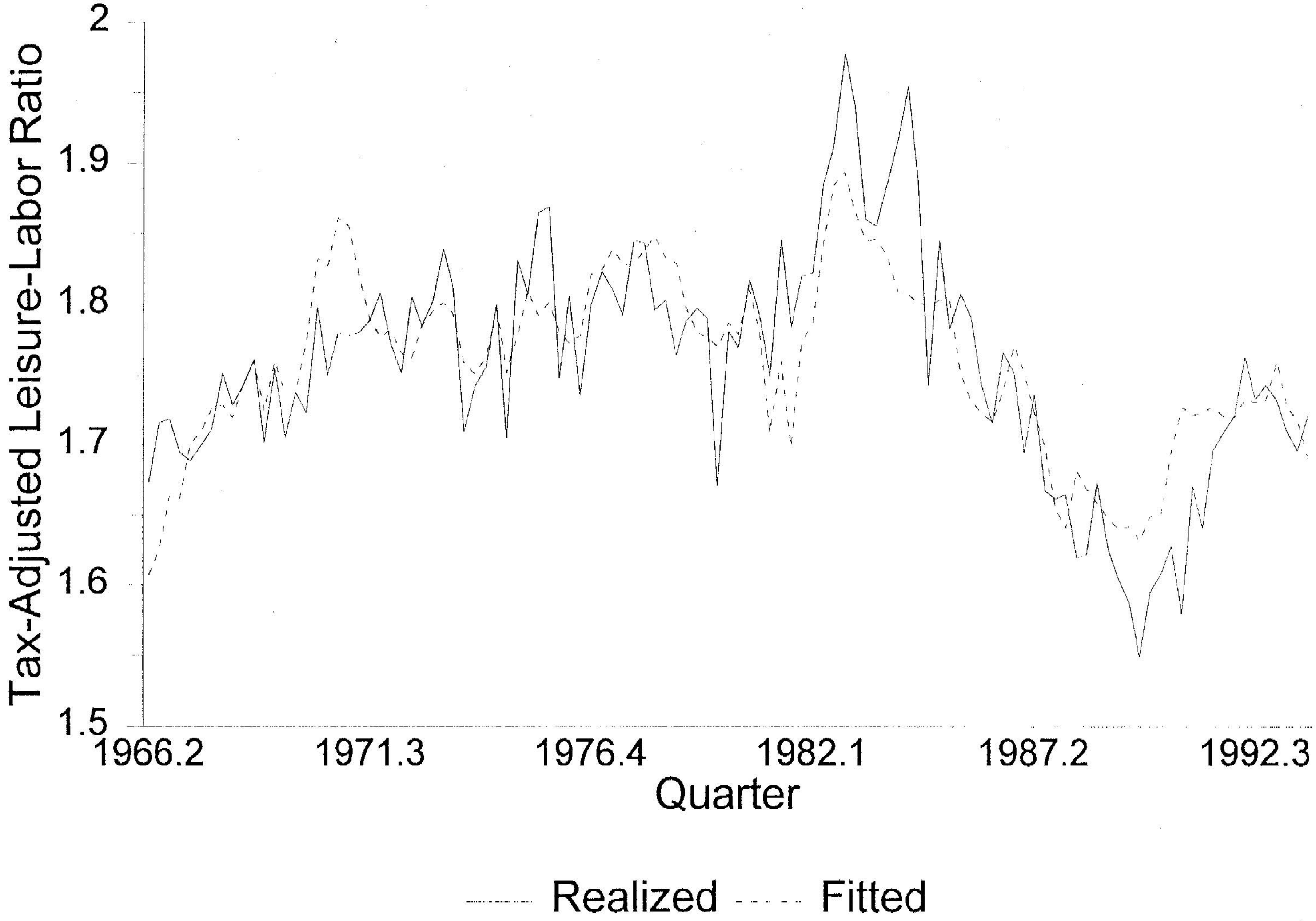
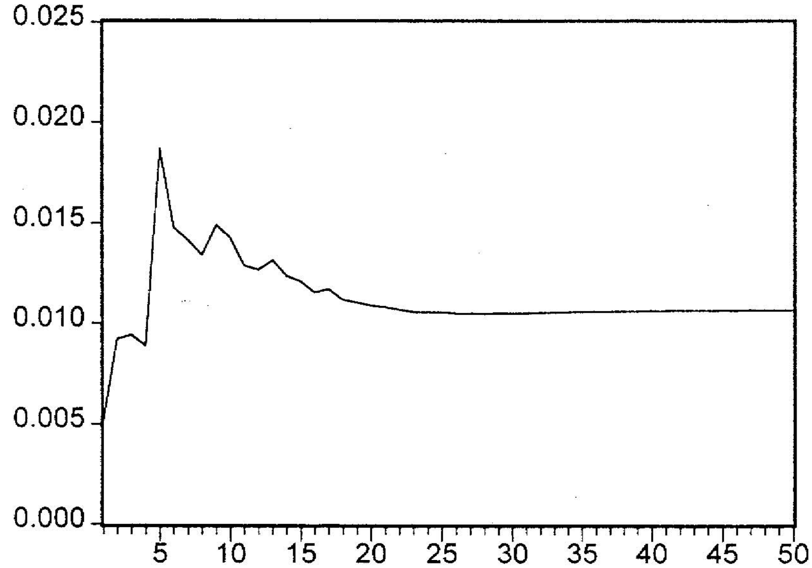


Figure 4. Responses to Tax Rate Innovation (Ordering 1)

a) Leisure-Labor Ratio



b) Labor Tax Rate

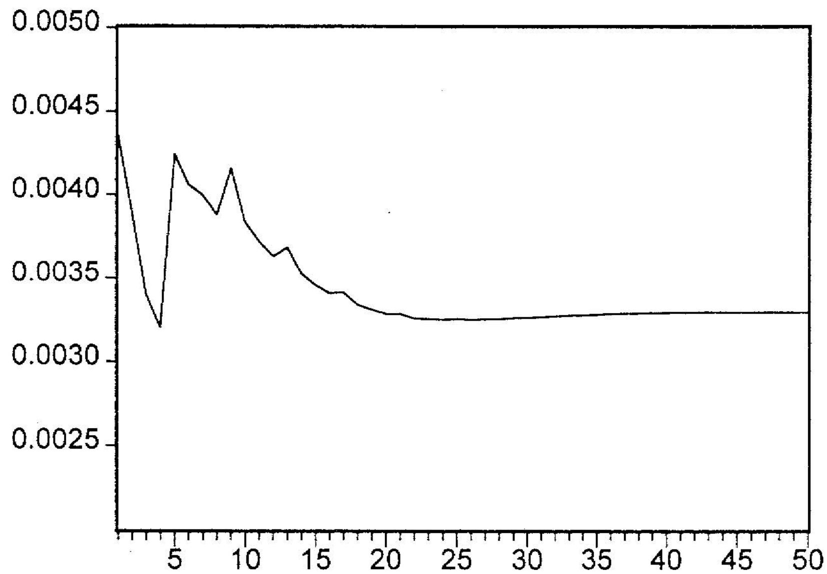
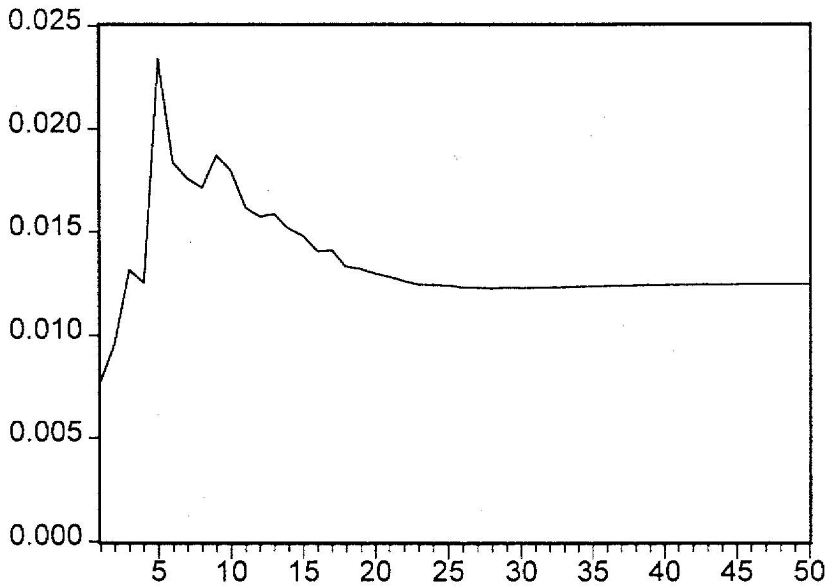


Figure 5. Responses to Tax Rate Innovation (Ordering 2)

a) Leisure-Labor Ratio



b) Labor Tax Rate