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Can we Rely upon Fiscal Policy Estimates in Countries with Unreported Production of 15 Per Cent (or more) of GDP?

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Abstract

This paper analyzes the effects of fiscal policy in Italy by employing a database containing two statistical novelties: quarterly fiscal variables on accrual basis and a time series estimate of tax evasion for the period 1981:1-2006:4. Following Revenue Agency suggestions, we use in a VECM the time series of the concealed VAT base as a proxy for the size of "unreported production", and define a *regular GDP* measure constructed as GDP net of government expenditure and evaded VAT base. The results reveal that we cannot rely upon the estimates of fiscal policy multipliers in countries with a sizeable unreported production unless the dynamics of the hidden and regular components of the GDP are disentangled. Changes in public spending and the tax rate generate a reallocation from underground to the regular economy which contributes to obscure the spending and tax effect on total GDP. In this setup the spending multiplier shows large long-run effects, considerably stronger than those registered in a model with no attention paid to unreported production. The drop in regular output, after an increase in the effective tax rate, tends to be considerable after one year, producing long-lasting effects and a significant increase in unreported production and tax evasion.

JEL-Code: C320, E620, H260, H620.

Keywords: fiscal policy, VECM, fiscal multipliers, unreported GDP, tax ratio, effective tax rate.

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The views expressed in the working paper are those of the authors and do not necessarily reflect those of the MEF and the DT.

1. Introduction[1]

In recent years policymakers have implemented an array of discretionary fiscal measures to stimulate economic activity and soften the economic downturn. At the same time, almost all OECD economies and many emerging countries have announced or implemented some sort of fiscal stimulus package. However, there persists a lack of consensus on the effects of fiscal policy and, in the economic literature, the impact of such fiscal measures remains uncertain. This is certainly the case for the euro area, and in particular for Italy, given the scarcity of relevant studies.[2]

The empirical literature often provides numerical estimates of the impact of an increase in government spending on GDP and employment in both the United States and in European countries. Such estimates contribute to determine the appropriate size and timing of countercyclical fiscal policy packages. The uncertainty about the quantitative effects of fiscal policy derives not only from the usual errors in empirical estimation but also from different views on the proper theoretical framework and econometric methodology. Here we emphasize that a crucial source of uncertainty is tied to the size and dynamic of tax evasion. A question that naturally arises in evaluating the effects of fiscal policy is whether we can rely upon fiscal policy multipliers estimated in countries with unreported production of 15 per cent (or more) of GDP. This question is relevant to many economies such as the Italian economy and most other economies in Mediterranean Europe. However, it can be generalized to other European countries and even to the US economy, where unreported production and tax evasion statistics are rare and the methods employed to calculate the phenomenon are not entirely clear. Thus, apart from the novelty of the results for Italy itself, we may generalize our findings to those obtained for all the countries with a sizeable unreported production.[3]

This paper analyzes the effects of fiscal policy in Italy by employing a database containing quarterly fiscal variables for the period 1981:1-2006:4. In addition, by exploiting the new yearly time series estimate of the unreported Value Added Tax (VAT) base provided by the Italian Revenue Agency we also provide a quarterly time series estimate of unreported production.[4] This estimate is extremely important, not only because it provides a long enough time series of tax evasion, but also because it allows the size of the underground production to be estimated. After all, evading VAT means under-reporting production, labor activities and revenues. Hence, the time series of the concealed VAT base, covering the period 1981-2006, can be used as a proxy for the size of "unreported production". This allows us to construct and use in the models two important GDP measures: the GDP net of government

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[2] For the Euro area see Perotti (2004), Biav and Girard (2005), De Castro and Hernandez De Cos (2006), Burriel et al. (2009) and Alfonso and Sousa (2009), whereas for Italy one of the few analyses is found in Giordano et al. (2007).

[3] There are several analyses on the size and dynamics of the underground economy for most of the industrialized and developing countries. See, for instance, Schneider and Buehn (2009). However, Statistical Offices and Revenue Agencies of developed countries have not produced time series data for the underground economy and tax evasion.

[4] See Marigliani and Pisani (2007). Appendix A reports further details about the construction of quarterly time series.

expenditure referred to as private GDP and regular GDP defined as GDP net of government expenditure and evaded VAT base. We perform a wide range of simulations with government expenditure and the tax rate to assess how these policies impact on regular GDP. The results are interpreted and compared with the private GDP responses, providing remarkable insights into the role played by underground production in the Italian economy and its relationships with the fiscal variables.

Finally, we do not follow the literature that seeks to remedy the shortage of fiscal data of national accounts by using OECD quarterly general government data or national quarterly cash-basis data, but we employ quarterly government national account estimates. Quarterly government estimates are computed by making use of a dynamic extension of the disaggregation method currently applied by the Italian Statistics Institute (ISTAT), avoiding the shortcomings associated with the use of the sources mentioned above.[5]

Our empirical analysis of fiscal policy, through Vector Autoregressive (VAR) models, is therefore, enriched by the decomposition of the GDP and the use of accrual data for public finance variables.
Since the works of Fatás and Mihov (2001) and Blanchard and Perotti (2002), there has been a large body of literature which used the structural VAR approach to estimate the effect of fiscal policy on macroeconomic aggregate variables, developing different methods to identify fiscal shocks. By contrast, we use a Vector Error Correction model, and study the short- and long-run effects of aggregated government spending and net tax shocks, controlling for the effects of the unreported production on the fiscal variables and GDP.

There are at least three issues that lead us to use a different framework. First, as mentioned in Perotti (2004) and many others, there is little guidance, theoretical or empirical, on how to identify structural shocks. Second, using the SVAR approach, important information, particularly useful in fiscal policy, is lost. Generally, economic theory has more to offer on the determination of equilibrium than on the nature of dynamic adjustments. When we perform empirical policy analysis, we like to obtain information on the underlying equilibrium tendency among a set of variables, but also to know the short-run dynamics and the adjustment coefficients. We are interested to know how, given a shock, the variables react and adjust on their path to equilibrium. Is adjustment slow or fast? Do some variables react more quickly and in response to different disequilibria? These dynamic interactions can often be very important and insightful in a policy analysis. Other than disentangling the single components of a shock (short-adjustment and long-run reactions) we are interested in the whole effect, because this is what we really observe with an uncritical impulse response analysis. Finally, in this framework a situation of special interest arises since several variables are driven by one or more stochastic trends, that is they have a particularly strong link that may be of interest in economic terms.

The main findings of this paper can be summarized as follows. Government spending shocks have, in general, a positive effect on GDP, which becomes increasingly relevant in the model with regular GDP (after one year the multiplier is 0.8 in the private GDP model and 3.7 in the regular). Tax rate shocks strongly reduce regular GDP and the effective tax rate displays a significant and positive influence on unreported production and tax evasion. Conversely, there is weak evidence of non-Keynesian effects on GDP when the hidden economy is not considered. As to fiscal variable interactions, a rise in the tax

[5] In Appendix B we summarize the characteristics of the estimated quarterly series for fiscal variables.

rate leads, in the short run, to a drop in government spending, whereas in the medium term we observe a complex dynamic interaction. Conversely, there is evidence of a striking and significant negative effect when looking at the reverse dynamic (i.e. a rise in public spending reduces the tax rate), which becomes increasingly powerful in the medium-long term. Moreover, shocks in the GDP (both private and regular) lead to a robust and permanent increase in government expenditure. Overall evidence about the public finance variables shows that the conduct of fiscal policy, during the observed sample, did not follow a rule aiming to comply with the intertemporal budget constraint, since a higher path of public spending was not associated with a higher path of taxes. Finally, the interaction between regular and unreported production demonstrates that the link between the two sectors is very strong, but also very harmful in the long term, since there is strong evidence that shocks to the unreported production have long-lasting negative effects for the regular economy.

All these points raise economic and policy implications which are extensively discussed in the paper, that is organized as follows. Section 2 and Appendices A-B describe the data. Section 3 outlines the specification of the VEC model and the identification method. Section 4 presents the results for government spending, tax rate and unreported production shocks, whereas Appendix C reports the statistical output of the models and the full set of impulse response analysis. Section 5 concludes and provides some policy issues.

2. Description of the data set

The data set used in this paper has two statistical novelties. First, since quarterly general government data, based on the European System of National Accounts, are available only from 1999, we elaborate our own time series estimates for several aggregates of public expenditure and taxation on an accrual basis. Second, we explicitly include in the model a time series for unreported production based on the undeclared VAT base provided by the Revenue Agency.

In Italy, VAR analysis with fiscal variables has been seriously impeded by the absence of long time series for government accrual data. Empirical applications have been "forced" to rely on either OECD quarterly general government data or quarterly cash-basis data. The sources of government budget data in recent works in Italy are the Ministry of Economy and Finance which has published quarterly cash figures since the early 1980s, and the Bank of Italy. To quote a recent paper, Giordano et al. (2007) provide SVAR estimates using cash quarterly data on the public expenditure and net taxes. [6]

So far, policy analysis relying on official quarterly general government data, compiled on a national accounts basis, has been feasible only when considering a few budgetary items, specifically non-market sector final consumption expenditure, non-market sector compensation of employees and VAT and other taxes on imports. The time series for these variables are available from ISTAT and cover the period 1981:1-2008:4. In this paper we elaborate long-run quarterly government accrual estimates by

[6] In Giordano et al. (2007) there is detailed analysis of the construction of the series and a comparison with national accounts data. The controversy on cash-basis and accrual basis data for analyzing the fiscal policy effects is explicitly admitted and discussed by the authors.

making use of a dynamic extension of the disaggregation method currently applied by ISTAT to compute estimates of Quarterly Economic Accounts (the method is described in Appendix B).[7]

The selection of fiscal variables used in the VEC model follows Blanchard and Perotti (2002) and Giordano et al. (2007). In particular, direct government spending is defined as the sum of government consumption and investment (which includes wages, current purchasing of public goods and services and public investment) while net taxes are defined as total government receipts, less transfers to households and firms.[8] More concretely, transfers include all expenditure items except public consumption, public investment and interest payments. Differently from most of the literature on fiscal policy effects, we use the average effective tax rate rather than the collected net revenues to better appreciate the complex interaction between tax rates and unreported production/tax evasion examined elsewhere (see Chiarini et al. 2011).[9] Such fiscal aggregates have become standard in this literature since government spending on goods and services might have different effects, as it directly affects the aggregate demand of the economy, while transfers and taxes exert their effects through real disposable income that could be partially saved.[10]

2.1 Fiscal aggregate time series

The seasonally adjusted figures in real terms for spending on goods and services are plotted in Figure 1. Government current spending on goods, services and public investment shows a steady increase over the sample period, with the significant exception of the period 1992-1996. This drop should be related to the considerable fall in public investments and the striking corrective measures taken by the government after the 1992 budget law which cut all public expenditure items in order to combat strong speculation on the exchange rate for default risk. After 1997, the slope of the positive trend becomes slightly less steep due to the consolidation effort in the period prior to Monetary Union. However, it then rose again in recent years due to the relaxation of the policy.

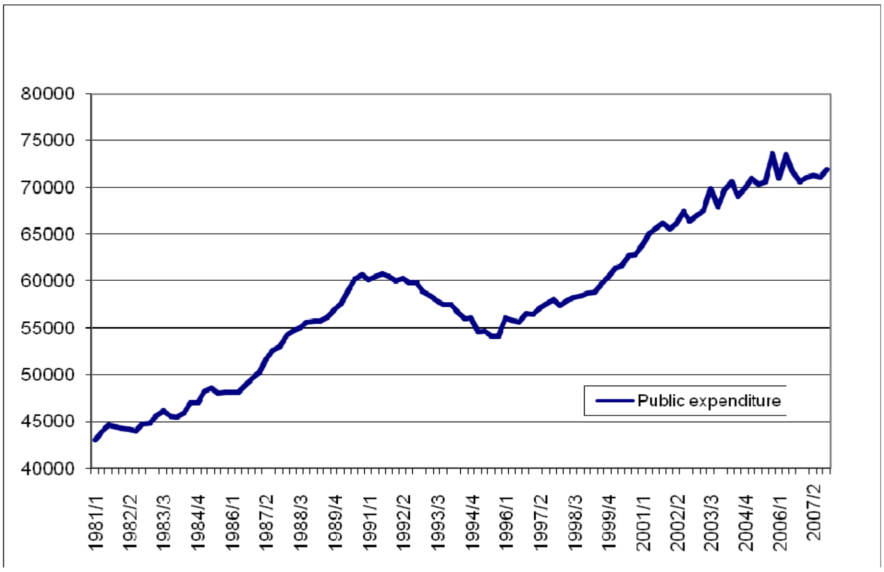
[7] We apply the dynamic extension of the Chow-Lin (1971) temporal disaggregation procedure suggested in Proietti (2005). This method is based on the state-space representation of a first order Autoregressive Distributed Lag model, which transforms the distribution problem into one of unknown observation.

[8] Similarly to other studies we excluded interest payments as they largely depend on the debt stock and therefore are not a discretionary fiscal policy tool.

[9] Although the traditional economic theory on tax evasion derives ambiguous predictions as to the tax rate impact on tax evasion, as emphasized in the literature reviewed in Slemrod and Yitzhaki (2002), and despite the substantial focus on these issues in policy analyses on tax evasion, surprisingly little is known about these problems from empirical studies. However, the evidence provided in a macroeconomic setting is quite homogeneous in claiming a positive effect of the tax rate on shadow activities (Schneider, 2002; Giles and Tedd, 2002; Davis and Henrekson, 2004). The evidence is less clear when considering tax evasion, since data are not available on an international basis and are difficult to collect even for OECD countries.

[10] See Burriel, de Castro et al. (2009).

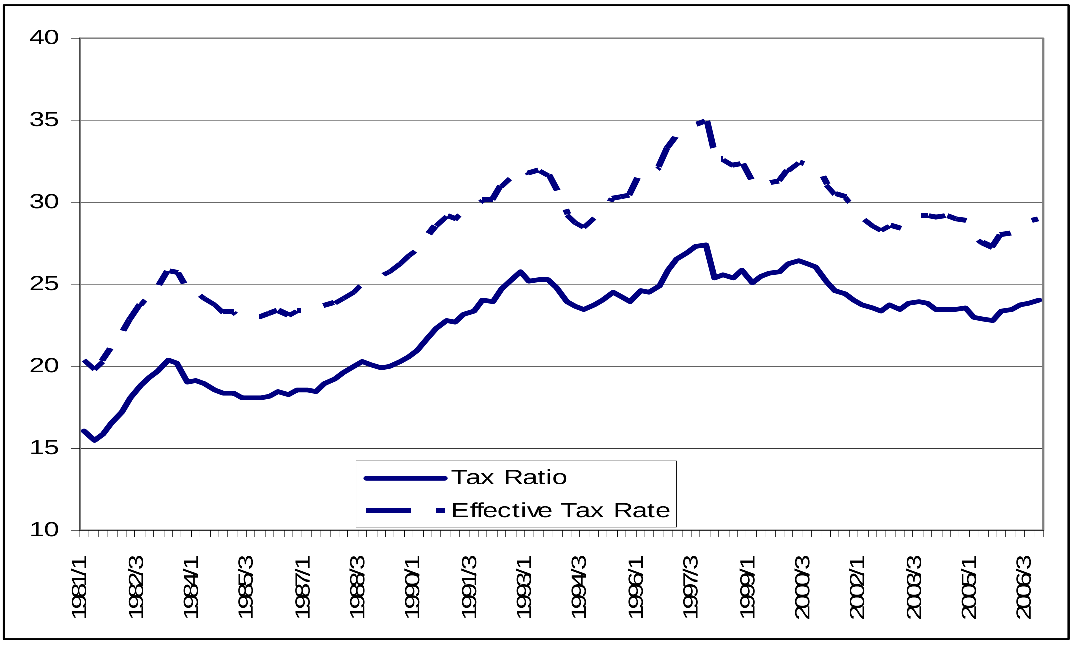
Figure 1 Direct public spending (consumption +investment, *euro millions, at 2000 prices***)**



In Figure 2 we display the evolution observed for the (net) tax revenues expressed as percentages of GDP (tax ratio hereafter) and the pattern of the (net) tax revenues expressed as percentages of GDP net of unreported production (effective tax rate hereafter).[11] Since GDP, by definition, includes unreported production, the relative average weight of taxes, in an economy with large tax evasion, is poorly represented by the above-defined tax ratio. In order to find an appropriate measure of the average fiscal burden we can consider the so-called effective tax rate, which is the ratio between tax revenues and the corresponding theoretical tax base, i.e. regular GDP. The tax ratio is the one usually considered when examining average fiscal pressure. However, this measure is downward-biased in countries with a considerable underground economy: by definition, the effective tax rate is always higher than the tax ratio, and a gap oscillating around 10-11 percentage points is observed, clearly indicating the huge burden that taxpayers have undergone due to the existence of underground activity and tax evasion.

[11] Following Barro (1981) and Bohn (1990) we use these measures instead of marginal tax rates because it is unclear what the marginal tax rate is on aggregate.

Figure 2: Net fiscal pressure: tax ratio and effective tax rate -percentages



A steep increasing trend may be observed the period 1980-1997, followed, in more recent years, by a slightly declining pattern, more accentuated for the effective tax rate and by a new surge at the end of the sample considered. However, significant reductions occurred over the periods 1984-85, 1994-96, 1998-99 and 2001-2005. These drops in net revenues were caused by many factors, such as fiscal reforms, policy changes and the introduction of new taxes. The fall in the first two periods reflects the drop in tax revenues due to a fall in economic activity and to the expiration of temporary tax increases in the previous years: the peaks in the tax rate curve also show the increases in tax revenues due to fiscal amnesties granted in 1982, 1991, and 1994, whereas the amnesties granted in 2002 are not clearly discernible. The sharp rise in tax rates, especially the effective one, observed in 1997 and, to a lesser extent, in 1993, is mainly due to extraordinary revenues, connected respectively to the so-called tax for Europe (1997) and to taxation on assets and buildings (1993). The introduction of a new tax (IRAP) replacing health contributions and many other taxes may also contribute to explain the reduction in 1998-99. Finally, the fall in interest payments on public debt after 2001 led to an expansionary policy and may well explain the new increase in the fiscal burden at the end of the sample.

2.2 Unreported and regular production time series

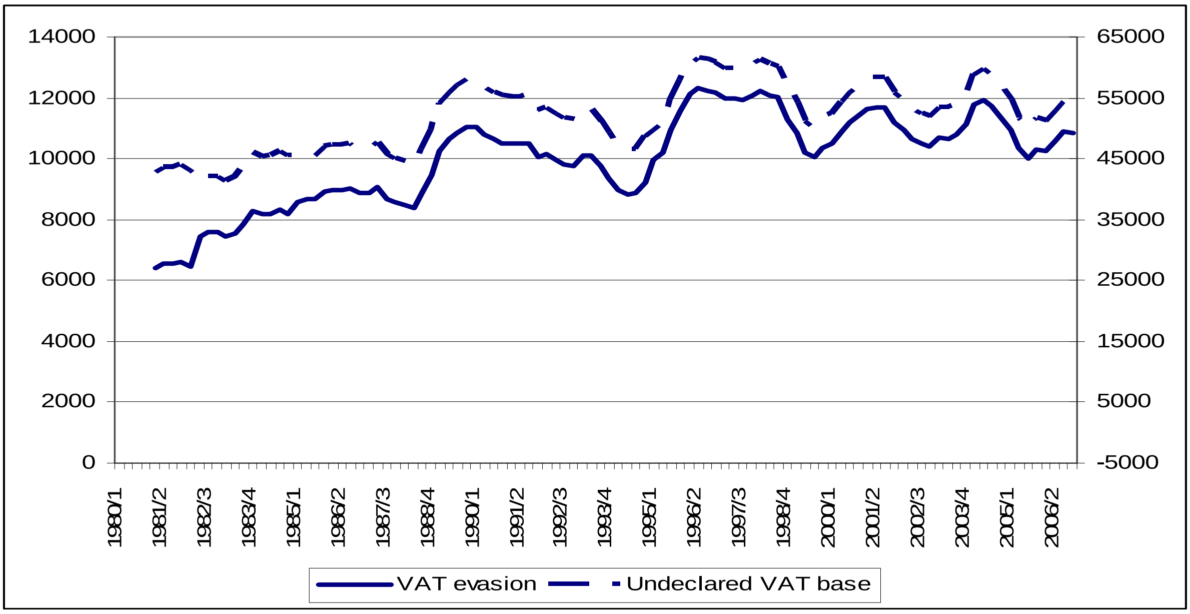
Data on unreported production in Italy are currently provided by the Revenue Agency of the Ministry of Economy and Finance, which has recently estimated a yearly time series of the non-reported VAT base. This fiscal aggregate is relevant to both unreported production and tax evasion. According to the data constructors, Marigliani and Pisani (2007), evading VAT means under-reporting production, labor and revenues. Hence, this time series estimate for the period 1981-2006 can be used as a proxy for the size of underground production which, in turn, can be used to estimate regular production.[12]

The size of the unreported VAT base ranges between 170 to 280 billion euros (real value) per year. Figure 3 shows the quarterly series for the unreported VAT base and the estimated VAT evasion. The latter is calculated by multiplying the unreported VAT base by the statutory tax rates. We emphasize that it is only part of the tax evasion phenomenon: it accounts for the uncollected VAT revenues. However, as outlined above, VAT evasion is a "prerequisite" and contains other forms of non-compliance. Therefore, according to the Revenue Agency, the dynamic of uncollected VAT revenues could well approximate the whole evaded tax revenues.

Both the series display considerable volatility, particularly during the second decade of the observed sample. The first half of the 1990s experienced considerable political instability and a fragmented approach to the fiscal policy, whereas during the period 1996-2000 a more stable political framework allowed the start of a process structurally reforming tax collection (Giannini, Guerra 2000). The two troughs observed in 1994 and 1999 are affected by a process of institutional reform. In particular, during the period 1991-93, some minor reforms were introduced, namely the minimum tax and "congruity" coefficients. It is also remarkable that the peak in VAT evasion in 1996 occurred after the tax amnesty (concordato fiscale) granted in 1994, whose receipts were mainly collected in 1995. The sharp reduction in VAT evasion observed during 1996-99 can be explained by structural innovations, such as the tax on line system (fisco telematico) and the new tax returns filing system (Unico form) introduced in 1998, together with Sector Studies (Studi di Settore), procedures midway between audit selection mechanisms and methods of presumptive (normal) taxation (see Santoro 2007b). These two interventions, together with a reorganization process of the fiscal authority starting in 1997, contributed to improve the efficiency of tax administration, indirectly increasing the effectiveness of auditing. The new upward trend registered in the last years of the sample can be explained by a learning process, with tax evaders being "more skilled" with respect to the new tax collection procedures, and perhaps also by an indirect effect due to the fiscal amnesties granted in 2002.
Given the dynamics of tax evasion and the manner in which it mimics that of the unreported VAT base, in this paper we use the two terms indistinguishably, keeping in mind, of course, the diversity regarding their intensity, as depicted in Figure 3.

[12] The approach for assessing declared and undeclared VAT taxable amounts, as well as the corresponding income, is based on a comparison of *actual* values, derived mainly from VAT returns, and *theoretical* ones, derived from National Accounts macroeconomic data. The latter aggregates are estimated selecting the national account expenditure categories that comprehensively cover VAT liabilities: i) household spending and non-profit institutions serving household final consumption expenditure; ii) central government current and capital expenditure; iii) exempt sector intermediate consumption; iv) other expenditure which incurs non-refundable VAT. For each of the listed items the most appropriate data source is chosen in order to respect VAT rules.

Figure 3: Undeclared VAT base (right scale) and VAT evasion (left scale), values in real terms (euro millions at 2000 prices)

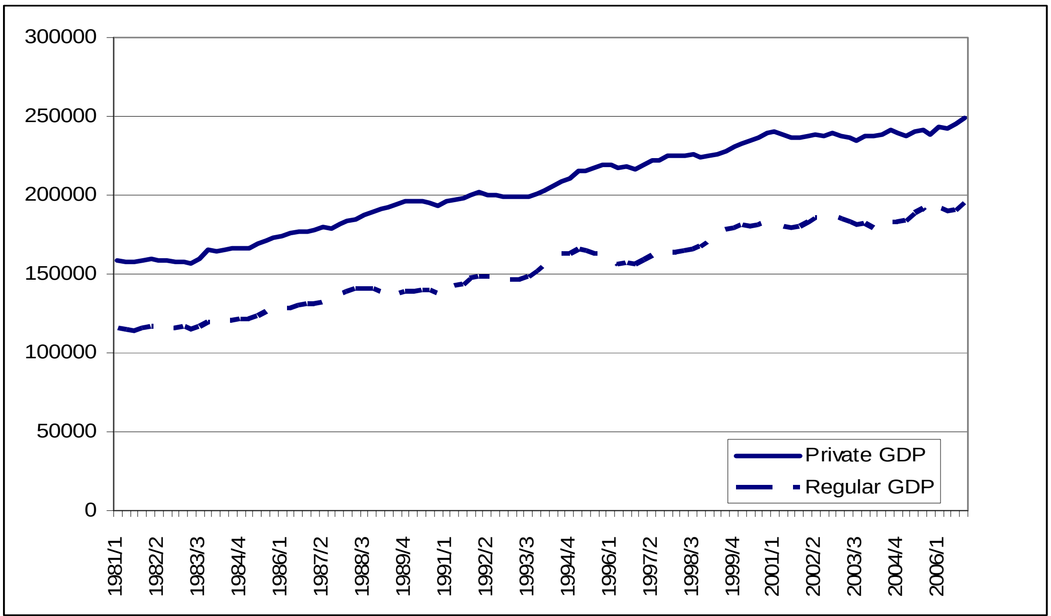


Since the concealed VAT base can be considered a measure of underground production, we are able to provide an estimate of regular production (Figure 4). Actually, Italian national accounts accomplish the requirement of exhaustiveness, as stated by OECD and Eurostat, including the value added generated in the underground economy. Therefore, on subtracting the estimated undeclared VAT base from the official GDP (national accounts), we get the regular production. This way of dealing with aggregate regular GDP may be considered to be rather crude. However, the literature provides empirical measures of the hidden economy that vary enormously in terms of methodology employed, reliability of the data and magnitudes estimated. Here we try to overcome many of these weaknesses using the official data available and, without "heroic assumptions" [13], we explicitly measure the share of national production or income deliberately concealed from observation, against the VAT revenues not reported to the tax authorities (i.e. produced in underground activities). These two measures may have much in common, since the unreported VAT base may be considered an important device that helps to conceal the tax base of other taxes, hiding shadow activities.[14]

[13] See, for instance, the Economic Journal symposium on the Hidden Economy and Schneider and Enste (2002).

[14] Indeed, Italian national accounts provide an exhaustive estimate of GDP, but only since 1992 have they also distinguished the share to attribute to missing economic activity. Marigliani and Pisani (2007) compare their estimate of tax evasion (here exploited) with the ISTAT estimate of the underground economy for the available years in common, i.e. 1992-2004, finding no large differences.

Figure 4. GDP net of public spending: private and regular (euro millions, at 2000 prices)



3. A VEC model for fiscal policy

The primary objective of our analysis is to quantify the fiscal policy effects on regular GDP, Y_R, disentangling the role of the unreported production, Y_U.[15] The benchmark statistical model is in four stochastic variables, namely regular GDP (Y_R), unreported production (Y_U), public expenditure (G) and effective tax rate (τ). In order to fully appreciate the importance of including the unreported production in the analysis, we compare the fiscal multipliers of our original model with those estimated in a model where the two components of the aggregate economic activity are not discernible. Therefore, our **benchmark** model will be compared to a **control** model where the measure of the aggregate economic activity is private GDP (total GDP net of public spending), and fiscal pressure is measured by the tax ratio, i.e. no attention is paid to the existence of tax evasion and unreported production. Under this hypothesis the control model would be in three stochastic variables, namely private GDP (Y_P), public expenditure (G) and the tax ratio ($\bar{\tau}$). Both the models take into account the dynamic of the public debt, whose growth rate is modeled as an exogenous variable. All the variables are in log (except the tax rate and the tax ratio) and nonstationary time series.[16]

3.1. Deterministic variables

Careful observation of the four previous figures gives a good idea of the difficulties in modeling and determining the statistical properties of the variables involved. The time series behave as random walk with drift and exhibit long swings away from their mean value (undeclared tax base, regular GDP and net fiscal pressure variables), accelerations and sudden "jumps" (effective tax rate, tax ratio and undeclared tax base), and broken trends (public expenditure).

[15] As described in Section 2, the variable Y_U is the unreported tax base calculated by the Italian Revenue Agency.
[16] In Appendix C we report the unit root.

Moreover, the VEC model contains variables both in levels and in differences. To cope with these dynamics a careful analysis of the dummy variables is a linchpin of a correct modeling strategy (see Juselius 2006). Our strategy identifies the outliers which, like swings, accelerations, jumps and reverse trends, cause sudden changes in the variables. To this end we based our investigation on a graphical analysis, aiming to obtain tentative recognition of possible outlier observations in the differenced processes. This analysis suggests augmenting the VEC by five dummy variables. Two dummies were included for describing two transitory shocks in quarters 2003:3-4 and 2005:4-2006:1. These dummy variables are used to model situations characterized by a shock at a time immediately followed by a similar but opposite shock in the aftermath. The two dummies imply a positive outlier in the levels of the public spending, observed in 2003:3 and in 2005:4. In 2003 government consumption grew at a rate of 2.2%, compared to 1.9% registered in 2002, whereas public investment recorded, in the same year, a sharp negative change, -2.1%. These two opposite patterns of the public expenditure variables could explain the peak in third quarter of the year. The outlier for 2005:4, is related to the "*annus horribilis*" for the Italian public finances (Pisauro, 2006), with a maximum peak for primary spending, a minimum registered for fiscal receipts, a deep (upward) revision in the figures of the government deficit and the inevitable official opening of an infraction procedure against Italy (Eurostat, 2005).

A further three impulse dummies (describing a permanent intervention/shock) augment the VAR for the observations:1989:1; 1983:4; 1998:1. While the first dummy accounts for an outlier in the unreported production series, an upward shift in unreported production in 1989:1, the remaining impulses catch anomalous observations registered for the tax rate: two downward shifts for the net tax rate in 1983:4 and 1998:1 (possibly subsequent to the collection of the revenue from the amnesty granted in 1982 and after the tax for Europe collected in 1997, and the introduction of the local business tax -IRAP- in 1998). Finally, a further shift dummy variable was included in the system, with shift date 1992:1:

$$S92 = \begin{cases} 0, & t < 1992:1 \\ 1, & t \geq 1992:1 \end{cases} \tag{1}$$

This shift function was included (and restricted in the cointegration space) to account for the "regime" change in fiscal policy after 1992, to cope with the exchange rate and debt crisis and the new pattern required from that date of the members of the future European monetary union. [17]

In order to test for cointegration, we conduct our analysis using a VAR with five lags on all stochastic variables.[18] The VAR model can be represented in a vector error correction form:

$$\Delta y_t = \Pi y_{t-1} + \sum_{i=1}^{p-1} \Gamma \Delta y_{t-i} + B z_t + \varepsilon_t \tag{2}$$

[17] Of course, since the shift dummy is restricted to lying in the cointegration space, its difference (current and lagged) is also included as unrestricted in the VEC equations.

[18] The appropriateness of the lag order was tested using the Akaike Information Criterion (AIC) and Final Prediction Error (FPE).

where $$\Pi = \sum_{i=1}^{p} A_i - I; \quad \Gamma_i = -\sum_{j=i+1}^{p} A_j$$

In our case, y_t is a vector containing four nonstationary variables (I(1)), z_t is a vector of conditioning variables (non-stochastic variables such as dummies and others that are weakly exogenous) and ε_t is a vector of innovation. A_i and B are matrices of coefficients to be estimated. It is well known that if the coefficient of the (*4x4*) matrix Π has a reduced rank *(r<4=2* is the number of cointegrating relations in our case), there exist matrices α and β (both *4x2*) such that $\Pi = \alpha\beta'$ where β is a cointegrating vector and α are the adjustment parameters.[19]

3.2. The stationary space

Estimation is carried out over the period 1981:1-2006:4 using a two-stage procedure (S2S).[20] Testing for cointegration (Johansen Trace statistic) with a constant and a structural break, provides evidence of two cointegrating vectors in our data set.

3.2.1 Just-identifying restrictions

It is well known that the existence of one or more cointegrating vectors implies the existence of common stochastic trends among otherwise non-stationary variables. In particular, given *k* variables and a rank of cointegration *r*, there are *k-r* common trends, and hence there can be at most *r* shocks with transitory effects (zero long-run impact) and there must be at least *k-r* shocks with permanent effects. In our benchmark model, we have *2* common trends, and hence two permanent shocks and an equal number of temporary shocks. In order to identify the source of the stochastic trends in our system, following Lutkepohl (2007), we adopt the normalization of the cointegration matrix given as follows:

$$\beta = \begin{bmatrix} I_r \\ \beta_{(k-r)} \end{bmatrix} \tag{3}$$

where $\beta_{(k-r))}$ is a sub-matrix of dimension *((k-r)* x *r)*, so that we get a convenient exact identification of matrix β, with normalization to one and zero restrictions. This representation is known as the triangular representation of a cointegrated system, and it has the convenient property that the variables not normalized to one (public spending and tax rate, in our case) represent stochastic trends driving the system.

As we will see in Section 4.2, impulse response analysis shows long-lasting effects for public finance variables in our benchmark model. This is consistent with identification of the variables as stochastic trends, and their long-run effects are consistent with the ordering chosen in this model.

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[19] See Johansen (1995) and Juselius (2006) amongst others.

[20] See Lutkepohl and Kratzig (2004). Diagnostic tests show a good descriptive power of the system. Diagnostic tests, parsimonious versions of the models and plots of cointegrating vectors, are presented in Appendix C. Further results may be provided by the authors upon request.

3.2.2 Over-identifying restrictions

After having imposed the *r-1* exactly identifying restrictions, and a normalizing coefficient on each cointegration vector, a number of theoretically motivated testable questions may be raised (does government spending cointegrate with the tax rate?, etc) on the cointegration space $Sp(\beta)$:

$$\left(Y_R, Y_U, G, \tau, S92, \mu\right) \in Sp(\beta) \tag{4}$$

where Y_R, Y_U, G, τ, μ represents, respectively, regular GDP, unreported production, public expenditure, the effective tax rate and a constant term. With references to the model with regular GDP, this process ends up with the following description of the cointegration space (standard errors in brackets):

$$\beta' y_{t-1} = \begin{bmatrix} 1 & 0 & \underset{(0.06)}{-0.7} & 0 & \underset{(0.017)}{-0.17} & \underset{(0.63)}{-4.1} \\ 0 & 1 & \underset{(0.06)}{-0.26} & \underset{(0.002)}{-0.02} & 0 & \underset{(0.6)}{-7.4} \end{bmatrix} \begin{pmatrix} Y_{R,t-1} \\ Y_{U,t-1} \\ G_{t-1} \\ \tau_{t-1} \\ S92 \\ \mu \end{pmatrix} \tag{5}$$

The cointegrating vectors are overidentified as two restrictions are imposed on each of them. The Wald test for the beta-restrictions (using Johansen ML estimation) is distributed as a $\chi^2(2)$ and under the null gives a *p-value* of 0.08. In our case we have two linear combinations for which the variance is bounded. One of these seems to support the existence of a long-run positive relationship between G and regular GDP: $ECM_1 = Y_R - 0.7 \cdot G - 0.17 \cdot S92 - 4.1 \cdot \mu$ whereas the other stationary relation clearly emphasizes a positive long-run effect of fiscal variables on tax evasion: $ECM_2 = Y_U - 0.26 \cdot G - 0.02 \cdot \tau - 7.4 \cdot \mu$.

However, it has been widely pointed out that cointegration vectors cannot be interpreted as representing structural equations because they are obtained from the reduced form of a system where all the variables are jointly endogenous.

Caution should be used to interpret the estimated coefficients. They cannot be considered as elasticities, even if the variables are in log, because all the other dynamic relations between the variables which are specified in the VAR model are ignored. Impulse response analysis, taking into account the full system, may provide a more reliable conclusion. This means that the coefficients provided by the cointegration analysis are only indicative.[21]

[21] Lutkepohl (1991, 1994) shows that the *ceteris paribus* assumption may not have a meaning. See also Dickey, Jansen and Thornton (1991). For an updated work, see Juselius (2006) and Lutkepohl (2007).

Nevertheless, the two cointegrating vectors might be thought of as arising from a constraint that an economic structure imposes on the long-run relationships among the jointly endogenous fiscal variables. We can interpret the first cointegrating vector, the stable relationship between regular GDP and public spending (we will see below that all the loading factors are statistically significant), as *the long-run spending multiplier*. Importantly, the linear relationship between these two variables is correctly not forced through the origin. Including a constant in the log-linear relation involves a "scale factor" that has the interpretation of the average propensity to growth. Conversely, the second cointegrating relationship envisages the long-run determinants of tax evasion (*the tax evasion equilibrium*).

In the second ECM, a possible explanation of the positive sign between the fiscal variables and the unreported production concerns Public Choice results on the taxpayer- government exchange relationship. The desired level of tax evasion involves complex evaluation of the fairness of the tax system (see, for instance, Bordignon 1993, Torgler 2007), which combines both the perception of the goods and services offered by the state and the level of tax to be paid. The taxpayer shall assess whether the tax paid differs from his/her fair terms of trade with the government, which depends crucially on the valuation of quantity and quality of public goods supplied. In the examined sample, it seems that the publicly provided goods are perceived as over-provided since larger spending would spur the concealment of tax liabilities. This being the case, we observe a positive long-run equilibrium relationship between public spending and tax evasion. Thus, high tax rates and high public expenditure levels are accompanied by a higher level of tax evasion: in the long term the intertemporal budget constraint requires that high expenditure corresponds to high tax revenues.

3.3. Adjustment coefficients

Using a two-stage procedure in which the beta vectors are estimated first and then fixed in the second stage, we may treat alphas in the same way as the short-run parameters. The strategy chosen is sequential elimination of the short run parameters and loading factors, based on model selection criteria (AIC). [22] The search for zero-restrictions on loading factors provides the following matrix:

$$\begin{pmatrix} \alpha_{Y_R 1} = \underset{(0.02)}{-0.04} & \alpha_{Y_R 2} = 0 \\ \alpha_{Y_U 1} = \underset{(0.04)}{0.21} & \alpha_{Y_U 2} = \underset{(0.03)}{-0.18} \\ \alpha_{G1} = \underset{(0.02)}{0.1} & \alpha_{G2} = 0 \\ \alpha_{\tau 1} = \underset{(0.72)}{1.28} & \alpha_{\tau 2} = \underset{(0.72)}{3.99} \end{pmatrix} \tag{6}$$

[22] The zero restrictions imposed on the short-term parameters of the VECM, including the loading factors, are based on model selection criteria. Precisely, the System Sequential Elimination of Regressors procedure implemented in *JMulTi*, controls, at each step, for the parameter with the smallest t-ratio (a threshold value of 1.6 is specified). The decision regarding the elimination is based on the Akaike selection criterion.

The α coefficients relate the error correction terms ECM_1, ECM_2 (already commented upon), with the first differences (the short-run) of the endogenous variables Y_R, Y_U, G and τ. Thus, $\alpha_{Y_U 1}$ is the adjustment coefficient to the first long-run relation (error correction) in the unreported GDP equation; $\alpha_{Y_U 2}$ is the adjustment coefficient to the second error correction described above (ECM_2) in the unreported GDP equation and so on. The results indicate that the equation for $\Delta y_{R,t}$ contains no information about the second long-run relationship since the second cointegration vector does not enter into this equation (this is also consistent with the fact that regular GDP is zero restricted in this ECM_2). An interesting dynamic aspect to note concerns the speed of adjustment to disequilibrium. Coming to our estimates, the coefficient $\alpha_{Y_U 2}$, which represents the speed at which unreported production is equilibrium-correcting, indicates that 18% of the disequilibrium in the long-run determinants of tax evasion is removed in a quarter, more than 70% in a year. With respect to the first stationary relation (the spending multiplier), the intensity of the adjustment of the hidden production is similar, though this has no influence on restoring the equilibrium since the unreported production is restricted to zero in ECM_1. However, the high estimated value for the loading factor suggests that unreported production is very responsive to the economic conditions occurring in the regular economy.

The speed of the adjustment of public spending is significant with reference only to the first ECM, and contributes to the adjustment toward the equilibrium in the public spending multiplier vector.[23]

By contrast, the tax rate is responsive to both the long-run equilibriums, with a faster speed of adjustment for the second cointegration relation (the tax evasion equilibrium) compared to the first one (the public spending multiplier). The positive value for $\alpha_{\tau,2}$ would imply that an excess in the concealed tax base necessarily leads to higher fiscal pressure. However, despite the very large loading factor, the ultimate increase in tax rate is quite small: it would take almost three years to accomplish the adjustment (12.5 quarters). Conversely, when the adjustment takes place through unreported production, only one year would be necessary to restore the long-run relationship, since 18% of the disequilibrium is removed within a quarter. Notice that in the second stationary relationship, the loading factor for the public expenditure is zero ($\alpha_{G,2} = 0$), as in Italy public expenditure is extremely rigid downwards.[24]

[23] Indeed, public spending enters this ECM with a coefficient of -0.7, whereas the tax rate enters with a zero coefficient.

[24] In this respect, it should be recalled that the tax rate is the only variable which is not log-transformed. Therefore, the loading factors provide a percentage variation for all the endogenous variables with the single exception of the tax rate, where the reaction to the ECM is measured as an absolute change. For instance, for a given level of public expenditure, when actual tax evasion exceeds the target defined by the fiscal variables, namely $y_{U,t-1} > 7.4 + 0.26G_{t-1} + 0.02\tau_{t-1}$, to keep the target y_U must be reduced and/or the effective tax rate must be increased. For instance, if the disequilibrium in the long-run determinants of tax evasion amounts to, say, 3%, the short-run adjustment in the tax rate and unreported production are, respectively:

$$\Delta\tau_t = \alpha_{\tau,2} * ECM_{2,t-1} = 3.99 * 0.03 = 0.12 \quad \text{percentage point,}$$

Interestingly, the tax rate is equilibrium-correcting to the estimated public spending multiplier (the first stationary relationship). That is, a restrictive fiscal policy is managed when there is excess public spending compared to the level consistent with the size of regular production. The fiscal adjustment occurs not only through a reduction in spending, but also through a slight increase in the tax rate so as to finance at least part of the excess spending through (hopefully additional) tax revenues.

4. Structural analysis in a VECM framework

4.1 Choleski decomposition and Structural VECM

In the previous section, we discussed the identification of the long-run structure by embedding the two stationary relations in a dynamic equilibrium correction system. In this section we illustrate the results of impulse response analysis. In order to proceed, however, we need to identify also the short-run structure. We impose short-run zero restrictions by using a Choleski decomposition of the residual covariance matrix.

We order the variables in the estimated system as follows: regular GDP, unreported production, government spending and effective tax rate. Since we do not adjust the tax rate for the automatic response to the business cycle, as in Blanchard and Perotti (2002), the effective tax rate as well as public spending have no immediate effect on real variables, whereas they are affected by GDP. Consistently with the literature we set the tax rate to follow government spending, since during the sample period under consideration, Italy experienced sustainability problems and the tax rates were usually managed to run a balanced public budget. In addition, average tax rates are affected not only by government spending and business cycles (which have an immediate impact on the tax base), but also, of course, by tax compliance. As to the ordering of tax evasion and GDP, given that we focus mostly on regular GDP, we claim that the undeclared VAT base is plausibly affected by decisions taken in the regular economy and not vice versa. Therefore we order unreported production after regular GDP.

The decision to identify the short-run structure without imposing structural restrictions on the residuals, except for adopting the Choleski decomposition of the covariance matrix, is justified by two orders of reasons. The first concerns our purpose of investigating the effects of fiscal shocks on GDP after having explicitly included unreported production in the analysis. The latter, as well as regular production value added, is a macroeconomic aggregate mostly pertaining to the aggregate supply, whereas fiscal variables are mainly related to aggregate demand. Starting from the paper by Blanchard and Quah (1989), a plausible empirical identification relies on restricting the long-run effect of the demand shocks on output to zero. This restriction can also be effective in the very short run (instantaneous relations between the variables), allowing us to order demand shocks after supply shocks. This supports the identification based on the recursive ordering of the supply side variables followed by the demand side aggregates. Public expenditure and tax rate respond to economic

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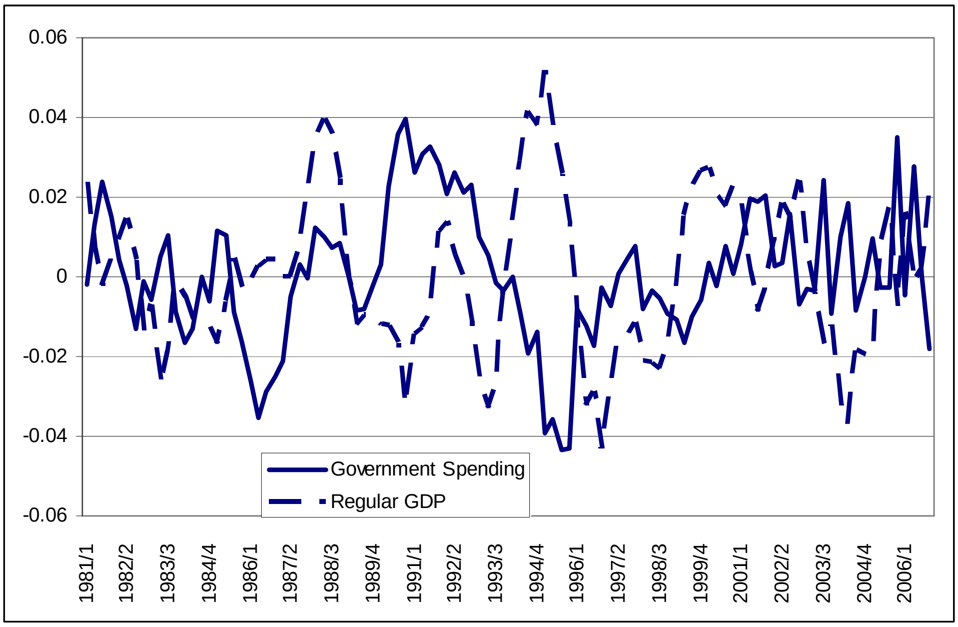
$$\Delta Y_{U,t} = \alpha_{Y_{U,2}} * ECM_{2,t-1} = -0.18 * 0.03 = 0.54 \quad \text{percent.}$$

It should be noted that the contribution of the short-run adjustment to the Error Correction mechanism is mediated by the beta coefficients estimated for each endogenous variable. For example, for the tax rate the final effect amounts to 0.02*0.12=0.0024, whereas for unreported production it is 1*0.18=0.18.

conditions, here described by the temporal evolution of private production which is separated into two components, regular and hidden production. Ordering public spending after GDP is consistent with Favero (2002), who investigates the effect of monetary and fiscal policy assuming the existence of a Taylor rule for fiscal policy. Figure 5 reports the HP cyclical components of public spending and regular production, and shows that in the short run fiscal policy is managed as a function of the business cycle (stabilization policy).[25] The second motivation we adopt to justify the Choleski decomposition is emphasized by Breitung at al. (2004) and Lutkepohl (2009): although imposing structural restrictions may resolve the non-uniqueness problem of innovations, it also raises the same order of criticisms already stressed by Sims (1980) with reference to econometric simultaneous equation models.[26]

Figure 5. Regular GDP and public spending: HP filtered series, percentage changes.



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[25] Favero (2002) investigates the behavior of monetary and fiscal authorities in the Euro area by modeling a fiscal reaction function, constructed by assuming that the behavior of fiscal authorities is determined by an output and a debt stabilization motive. The existence of a Taylor rule for fiscal policy was suggested for the US economy by Taylor in several papers (1996, 1997, 2000).

[26] We also carried out several identification schemes in a VECM structural framework (see, for instance, Breitung et al., 2004 and Juselius 2006). In modelling the Structural VECM , we use an order of the variables which is more consistent with the standard literature on fiscal policy effects, i.e. G, Y, TE, τ. According to cointegration analysis, we choose two temporary structural innovations, spending and tax evasion. In addition, to identify permanent shocks, we also restrict the tax rate to display only temporary effects on public spending; finally, the further short-run restriction is set in such a way that the first shock, public spending, does not have an instantaneous effect on private GDP. The picture we get from the structural impulse response confirms the fiscal policy multipliers, the significant impact of tax rate on tax evasion, though the opposite channel is no longer working. Finally, an increase in tax rates reduces public spending, whereas a spending shock causes a positive reaction in tax rate. The complete analysis is available upon request.

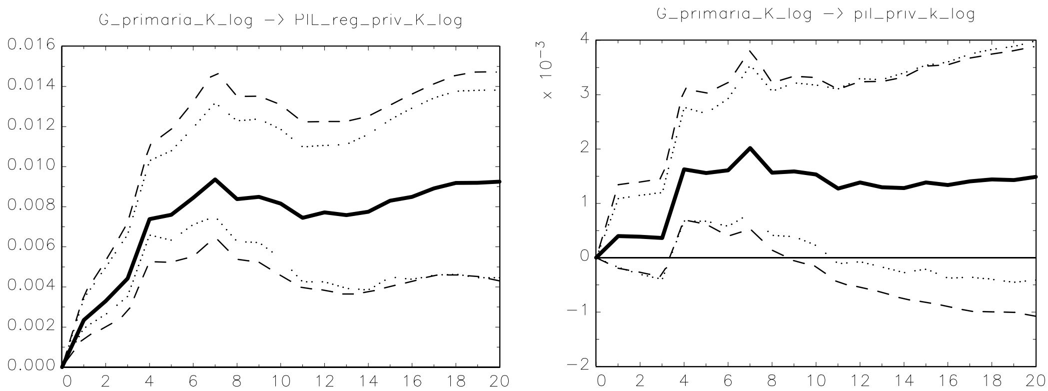
4.2 Fiscal policy multipliers

4.2.1 The public spending shock

The first remarkable result of our analysis concerns the interaction between fiscal policy and unreported production. In Italy the GDP statistics are an exhaustive measure of macroeconomic activity, accounting for the underground economy, as required by Eurostat. This has substantial consequences when estimating the size of fiscal multipliers with a measure of the macroeconomic activity that includes underground economy. In order to investigate this issue, we examine the responses of GDP to shocks to government spending with two different alternative measures of GDP: in the benchmark model we consider the reaction to the shocks by regular GDP (i.e. private GDP net of evaded VAT base), while in the control model we refer to private GDP (i.e. total GDP net of government spending).

Figure 6 displays the impulse responses to one standard deviation shock to government spending.[27] In Table 1 the original impulse responses are transformed such as to give the GDP response (both, regular and private) to a one-percentage point of total GDP shock to government expenditure.

Figure 6. The response of private GDP (right) and regular GDP (left) to one S.D. shock to public spending



Throughout the paper we define as "statistically significant" those estimates for which the error band, identified by the fifth and the eighty-fifth (ninety-fifth) percentiles, does not include zero (dotted and dashed line, respectively). Figure 6 shows a largely similar dynamic of the fiscal multipliers between the two models. However, the reaction in GDP to a positive shock in public spending is quicker and much more intense in the model with the regular economy, suggesting that it is the regular economy which drives the expansionary effect.

[27] As government spending and GDP are both measured in logs, the variations in the plots can be read as percentage changes of the GDP consequent to one standard deviation shock to government spending.

After one quarter, the output response is weaker for private GDP. Regular GDP exhibits a substantial and long-lasting increase in response to a government spending shock. The spending multiplier is, after one quarter, +1.2% and peaks after one year at +3.7%. The effects are still potent and statistically significant after 5 years. It should be kept in mind, however, that these digits correspond to a reaction of total GDP of 0.7% after one quarter and +2.1% after one year.

Table 1: the response of GDP to a positive shock to public spending and taxation

	Benchmark (Regular GDP and unreported GDP)	Control (Private GDP)
Shock to G by 1% of GDP		
Impact	0	0
+1	1.2%**	+0.2%
1 year (+4)	3.7%**	+0.8%**
3 years (+12)	3.9%**	+0.7%
5 years (+20)	4.7%*	+0.8%
Shock to tax rate by 1 percentage point		
Impact	0	0
+1	-0.3%**	+0.2%
1 year (+4)	-1.06%**	+0.2%
3 years (+12)	-0.4%	-0.00 %
5 years (+20)	-0.9%	-0.3%

** and (*) denote statistical significance, respectively, at the 5% and (15%) level.

These surprising findings show that the public expenditure effects on regular production in Italy are strong and very persistent. The results can, to a large extent, be traced back to differences in the GDP measure used in the models (in the period considered, on average, the total GDP is composed by 22% of direct public spending and 20% of underground production). Moreover, controlling for non-compliance allows us to take into account some crucial effects otherwise neglected: Figure 8 and Table 2, below, and the whole set of the impulse response plots in Appendix C, show that after an increase in public spending (wages, investment, goods and services) there is a positive effect on the regular economy, whereas the underground sector is shrunken, with long-lasting effects. Hence, changes in public spending cause a reallocation from underground to the regular economy with regular GDP which increases more than proportionately to government spending and the underground component of the economy which shrinks considerably.[28] This reallocation effect of the output components contributes to obscure the spending effect on total GDP.

The lesson here should be quite evident. Government manoeuvres involve both the production components. To determine the magnitude of these policies we should decompose their behaviors in order to highlight the short-run substitutability effects.

The intensity of the fiscal multiplier and the quicker and larger response observed in the specification with regular GDP adds important insights to the evidence available for Italy. Giordano et al. (2007) find that the response of Italian private GDP (net of public expenditure component) to an exogenous one per cent (in terms of private GDP) shock to public consumption expenditure is about 1.2 percentage points on impact, with a humped-shaped pattern. Conversely, a very weak effect of the fiscal multiplier is found, in a different identification framework, by Afonso and Sousa (2009), which supports the idea of a crowding-out effect via investment. While results available for the US (Blanchard and Perotti, 2002; Caldara and Kamps, 2008) support a positive effect of government spending on GDP, positive shocks in government spending increase output in several European countries although the effects are different: quite small in Germany (Heppke-Falk et al. 2006), limited to the short-run in Spain (De Castro and Hernandez de Cos 2006), and large in France (Biau and Girard 2005).[29]

Our results show that, in addition to the issues mentioned by these authors to motivate their results, for Italy and many countries with similar characteristics, the size of the government spending multiplier should be carefully analyzed, given the possible incorrect modelling of an important component of private supply, i.e. underground production. As we have seen, the empirical effects of fiscal policy on these two components of GDP, regular and underground, may be very different. By modelling them as a single production aggregate, the literature has so far omitted to consider that tax evasion is a reaction of a "rational agent" to fiscal pressure, providing potentially misleading results.[30]

4.2.2 The tax shock

In order to analyze the effect of an exogenous change in tax policy, in the presence of an unofficial production, we construct a synthetic tax indicator, i.e. the average effective tax rate.
Given the simplifying assumption that there exists a single aggregate statutory tax rate, $\hat{\tau}$, then the total tax revenues, which are paid by compliant tax payers, are given by $\hat{\tau} \cdot (Y_R)$. Therefore, our synthetic tax indicator, the effective tax rate:

[28] A complete picture of the impulse response analysis is reported in Appendix C.

[29] Afonso and Sousa (2009) find that for the US economy the effects on GDP are small, positive, but not significant, whereas the effects of government spending on German GDP are negative, reflecting a fall in private investment. Beetsma and Giuliodori (2010) estimate a VAR model in a panel format with annual data over the sample 1970-2004 for the EU-14 countries, and show that after a one-percent of GDP increase in government purchases, the GDP rises by 1.2% on impact, and peaks at 1.5% after one year. The authors stress that this substantial multiplier effect is consistent with the fact that most countries in the sample have featured only limited exchange rate flexibility against their main trading partners, which in the context of the Mundell-Fleming type model would imply a short-run economic stimulus after a fiscal expansion. An updated survey of fiscal multipliers in the literature may be found in Spilimbergo et al. (2009).

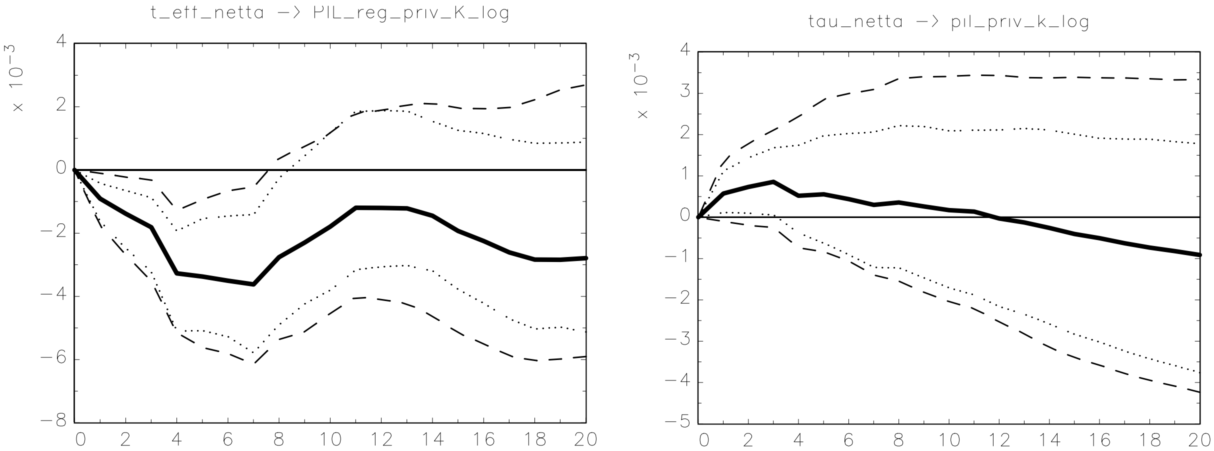
[30] See Busato and Chiarini (2004), and Busato, Chiarini and Marzano (2009).

$$\tau = \frac{\text{TaxRevenues}}{Y_R} = \frac{\hat{\tau} \cdot (Y_R)}{Y_R} \tag{7}$$

is a reliable measure of the aggregate fiscal burden for compliant tax payers. Moreover, we can be confident that every shock to the aggregate economic activity and/or to the level of compliance would not influence the size of this indicator, since, by construction, the numerator and the denominator of the ratio compensate each other.

The impulse responses calculated for a tax shock are shown in Figure 7 and Table 1. As regards the output responses, the disagreement in SVAR literature about the effects of a tax shock is well known. Depending on the identification approach, tax shocks in some cases hardly have any effects on the real economy, whereas in others they may display important distortionary effects.[31] Our results emphasize that regular output decreases by 0.3% after one quarter in response to a one percentage point increase in the effective (net) tax rate. The drop in regular output becomes serious over time, peaking to -1.1% after one year and driving the regular economy to a lower equilibrium (still -0.9% after five years, though not significantly different from zero). Although the negative pattern is confirmed for the private GDP model in the long run, the intensity and significance of the output responses to a tax shock is substantially reduced, whereas on the impact and in the medium term the model provides a positive effect. The figures emphasize that, when considering broader measures of the GDP, the negative impact of the tax rate tends to vanish.

Figure 7. The response of private GDP to one S.D. shock to the tax ratio (right) and the response of regular GDP to one S.D. shock to the effective tax rate (left)



The negative sign of the effect of a tax increase on the regular GDP represents a diverging result compared to the estimates for Italy provided by Giordano et al. (2007). Although they use net revenues

31 See, for instance, Blanchard and Perotti (2002), Perotti (2005), and Mountford and Uhlig (2005). For a review see Caldora and Kamps (2008).

as a fiscal variable, they find that a tax shock equal to one percentage point of GDP produces a transitory but positive and significant effect on private GDP. This result is not consistent with our evidence about regular GDP, whereas it is interestingly in line with the result we find when the two components of the GDP are not separated (the control model), and obscures the true taxation effect on economic activity. More recently, Afonso and Sousa (2009) found for the Italian economy that a shock to government revenues reduces GDP, although the effect is not persistent as it vanishes after four quarters.

As depicted by our impulse responses, the counterintuitive result of Giordano et al. may well be related to having neglected the underground economy as a component of official estimates of GDP, with the consequence that the effect of tax shocks on the these two components, regular and underground, cannot be disentangled. As shown in Figure 8 (see also Appendix C), an increase in the tax rate pushes up the hidden economy (and tax evasion) entailing a reallocation among sectors which blurs, at best, the total effect. After all, the tax burden is known to be the main cause for the growth of the shadow economy (Schneider and Enste, 2002, Chiarini et al., 2011, among others).

Turning to the evidence available for the US, Caldara and Kamps (2008) compare the effects of fiscal policy for different econometric / identification approaches in a SVAR framework. While their results always support the view that government spending shocks generate an increase in real GDP, they find strongly diverging results as regards the effects of tax shocks. Overall, there is no evidence of any "expansionary fiscal contraction".[32]

4.3 Unreported production and fiscal policy

The high tax burden is usually adjudged to be the main determinant of tax evasion, it being claimed that high taxation encourages taxpayers to expose smaller amounts of their revenues to fiscal authorities. The size of tax evasion varies according to the institutional framework as well as to individual characteristics.[33] Extensive tax evasion implies severe troubles for tax collection, with a loss of revenue that can be managed through a reduction in the spending side of the public budget, larger debt and an increase in tax rates.

Our analysis allows us to provide important elements of discussion useful for interpreting the composite effect of the evaded tax base on public finances, an issue of considerable interest for policy analysis. The impulse response, displaying the effects of an increase in tax rate and public spending on

[32] Giavazzi and Pagano (1990) were the first to argue that fiscal adjustments (deficit reductions) that are large, decisive and on the spending side could be expansionary. Alesina and Ardagna (1998) and Alesina and Perotti (1995) investigate various episodes of fiscal adjustments.

[33] For instance, the complexity of the tax system is claimed to be an incentive for tax avoidance, evasion and fraud (Veermend et al., 2008; Trandel and Snow, 1999) and tax morale has crucial importance, for many authors, in explaining tax compliance (Torgler, 2007). There is a large body of literature on tax evasion. See the Economic Journal Symposium on The Hidden Economy (1999) and Cowell (1990a; 1990b) for a survey of the tax evasion models. Many of these models are based on the tax-evader-as-gambler model first analyzed by Allingham and Sandmo (1972). However, in order to overcome some of the shortcomings of this approach, another stream of literature focuses on different determinants, for instance institutional or governance quality, and social norms (Torgler and Schneider, 2006; 2007) or law abidance (Orviska and Hudson, 2003; Orviska et al., 2006). For a study of the phenomenon in Italy, see, amongst others, Busato and Chiarini (2004), Bordignon and Zanardi (1997), Cannari and D'Alessio (2007), Chiarini et al. (2009; 2011), the papers in Guerra and Zanardi (2007), the ISAE Report (2006) and the papers on the Italian Revenue Agency web site: http://www1.agenziaentrate.it/ufficiostudi/.

unreported production, is shown in Figure 8 and quantified in Table 2 (of course, only for the benchmark model).

Table 2: The response of unreported GDP (benchmark model) to positive shocks to public spending and taxation

	Shock to G by 1% of GDP	**Shock to tax rate by 1 percentage point**
Impact	0	0
+1	-2.07%**	+1.7%**
1 year (+4)	-6.01%**	+4.2%**
3 years (+12)	-1.9%	+1.1%*
5 years (+20)	-3.7%**	+2.6%**

** and (*) denote statistical significance, respectively, at the 5% and (15%) level.

As expected, larger tax rates imply a larger concealment of production; after a shock to the effective tax rate of one percentage point, the unreported production rises by 1.7% after one quarter, to peak at more than 4% after 1 year. An increase in public spending, reallocating the hidden economy to the regular one, reduces unreported production by 2.1% after one quarter, dropping to about -6% after one year. This strong reallocation may be easily explained by the fact that the demand for goods originating from government necessarily spurs only the regular market: in the short run firms reallocate their resources to this segment. However, it should be kept in mind that since unreported GDP makes up about 20% of total GDP, the sharp fall in the underground component of GDP translates into a reduction in the total GDP of about one percentage point. When summing this effect to that already found for regular GDP, the total impact of the spending shock is about +0.28% after one quarter, rising to +0.9 after one year.

Figure 8. Response of unreported GDP to one S.D. shock of public spending (right) and one S.D. shock of the effective tax rate (left).

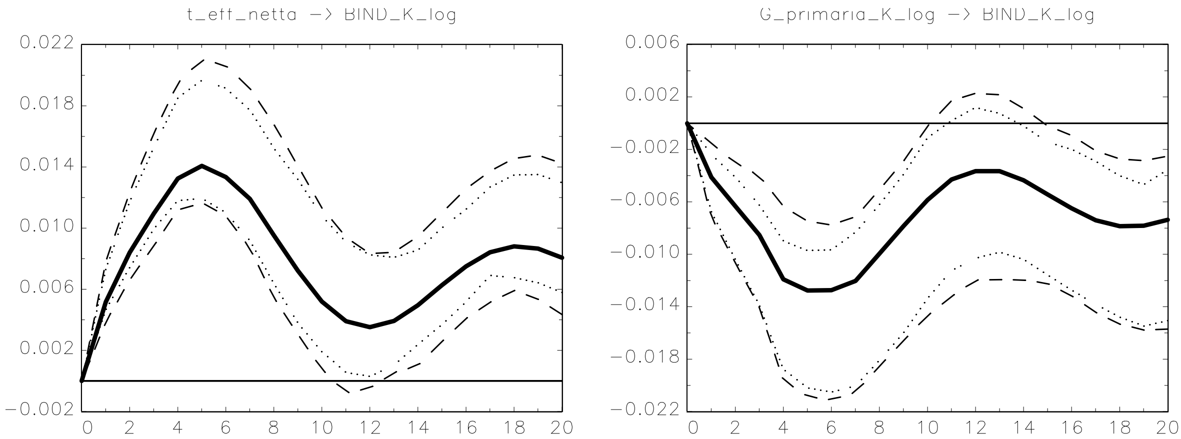
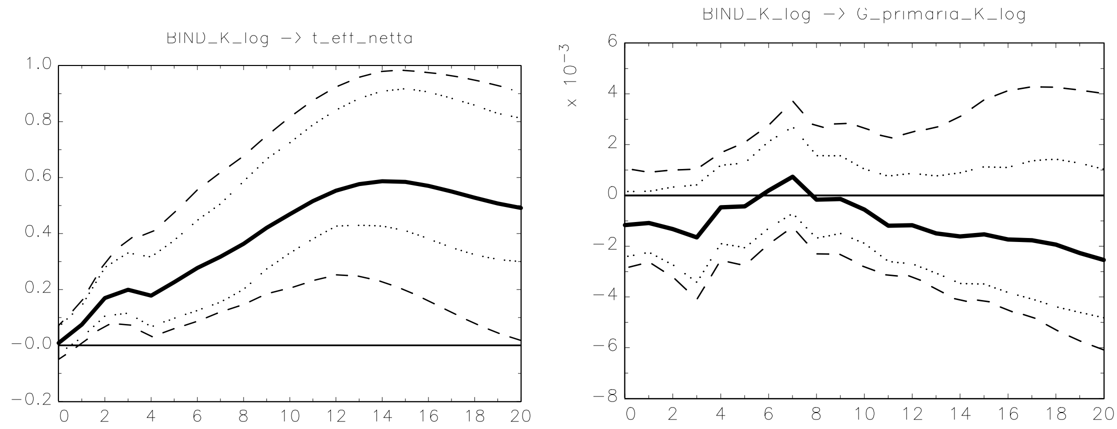


Figure 9 shows the response of the effective tax rates and public spending to an exogenous shock to unreported production. As regards the impact, a larger size of concealed production generates no significant effect on the effective tax rate. However, the long-run effect of a shock of 1% in the unreported production is an increase in the fiscal burden, leading to a rise in the effective tax rate of about .35 percentage points after three years. [34] This increase is considerable since the measure of the fiscal burden here adopted, consistent with the literature, is the net tax burden (revenues minus transfers to households and firms/GDP ratio).

With regard to the response of public expenditure, the data show no significant reaction to an increase in unreported production, suggesting that most of the adjustment of the public budget after a shock to (non-)compliance occurs through the revenue side. Finally, it should be emphasized that we are dealing with direct spending net of transfers, whereas a significant component of public spending is accounted for by transfers, which are included in the net tax rate. Hence, the positive reaction displayed by the tax rate to a shock in evasion could also be consequent to a reduction in government spending in transfers.

[34] Though the effective tax rate is expressed in absolute value and not in logs, one should take into account that the simulated shock is about one standard deviation, i.e., for unreported production, 1.7%.

Figure 9. The response of public spending (right) and effective tax rate (left) to one S.D. shock to unreported production



4.4 Tax and spending: some unexpected interactions

A large body of literature has debated the dynamic relationship between tax revenues and public expenditure. The issue is of major interest in situations characterized by large budget deficits, since it helps to understand the best strategy to reduce imbalances.

As claimed by Hoover and Sheffrin (1992), causality is a slippery concept. However, the causal question at stake in this debate is clear: if it is possible to intervene to control one of the variables (spending or taxation) directly, would that yield control over the other variable.

"**Fiscal synchronization**" of spending and taxation is implicit in many electoral models, based on individual optimal behavior. Under majority vote, the median voter supports the amount of public spending on taxation to satisfy the preferences of the community.[35] A clear direction of causation between spending and taxation is emphasized by the "**spend-and-tax**" hypothesis. Here changes in public spending would lead to taxation. Therefore, in order to cut the public deficit it is mandatory to limit the former. Peacock and Wiseman (1979) argue that even temporary increases in public spending (exceptional events such as war, environmental crisis) eventually lead to a permanent increase in taxation. In the macroeconomic literature, Barro's (1979) influential tax smoothing hypothesis states that the path of expenditures is exogenously given, and taxes are adjusted in order to minimize distorting costs, while the budget is balanced intertemporally.

In the opposite causal structure, the "**tax-and-spend"** hypothesis, the level of spending adjusts to the level of tax revenues available. In this view, larger tax revenues would increase spending and not lower public deficit. Authoritative supporters of this view, such as the economists of the "supply side approach", suggest a reduction in taxes to force subsequent spending cuts. Buchanan and Wagner (1977) share a similar view about the causal ordering but their analysis leads to an opposite interpretation about the sign of the interaction: rising taxes may limit the growth of public spending,

[35] See Musgrave (1966); Meltzer and Richard (1981).

since the taxpayer would perceive too high a price of government provided goods and push for spending reductions.

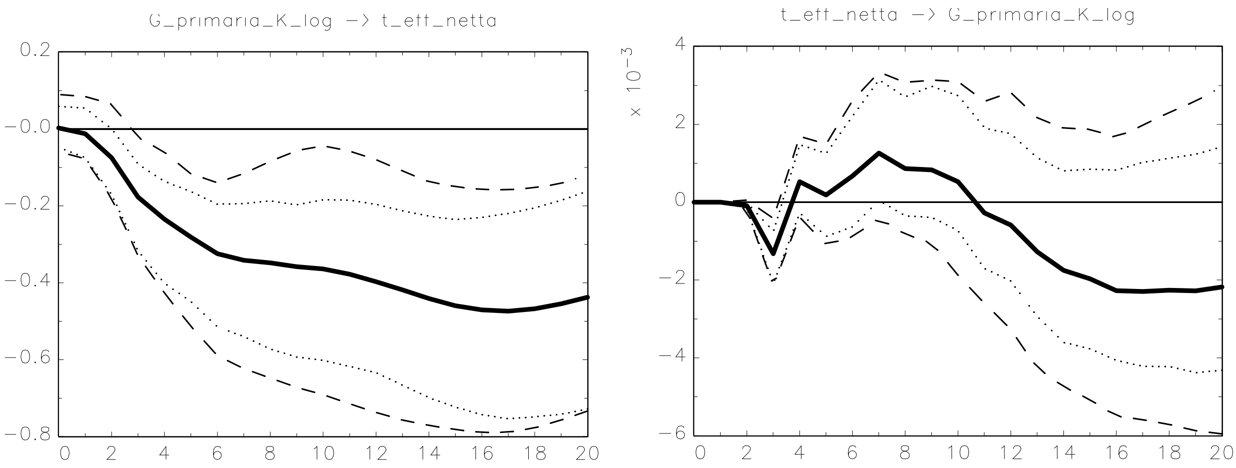
The recent literature using VAR analysis to evaluate the effects of fiscal policy seems to support the “spend-and-tax” hypothesis, since many of the available structural VAR models always order tax revenues after government spending.[36] Including unreported production in the analysis, we enrich the analysis with an issue which is as yet uninvestigated. A warning should be borne in mind: here we are not dealing with tax revenues, but with the effective tax rate. As shown in Figure 10, which reports the reactions of the fiscal variables to their reciprocal shocks, the responses would suggest that there is significant response of the (net) effective tax rate to government direct spending, i.e. spending *does* cause taxation, although the negative sign of this relationship is not the one expected. Why should an increase in public spending depress the tax rate? The answer is twofold, and is related to the role of unreported production and to the way public finances were managed in Italy during the examined period. On the one hand, public expenditure, as expressed in Fig. 8, pushes down the unreported component of the GDP, and, given the impact of this latter on the tax rate (see Fig.9) also the effective tax rate. On the other hand, as mentioned above, the evidence suggests that a sort of active fiscal policy was run in Italy during the examined period. Active fiscal policy means that higher taxes are not expected to fully finance the increase in government spending, i.e. taxes and spending do not adjust in order to stabilize government debt, which, by the way, increases almost throughout the time period examined here. In the Italian case, therefore, the electoral-political process has led to the extreme active fiscal policy rationale, running deficit spending (especially in the 1980s and 1990s) by limiting the raising of tax rates in order to gain more political consensus.[37]

By contrast, there is very weak evidence of the reverse working channel, taxes leading spending, and the dynamic interaction is quite complex (see Figure 10, right hand panel). In the very short run, and after 14 quarters, there is some evidence consistent with Buchanan and Wagner (1977): a positive shock in the effective tax rate generates a decline in spending. This result, already displayed by Giordano et al. (2007), could be related to the pressure for fiscal consolidation, but is not statistically significant. Tax shocks have been related to restrictive policies to manage fiscal budget sustainability, urging the fiscal authorities to adopt more severe schemes also on the spending side.

[36] See Caldara and Kamps (2008) for a review of the various approaches to identify fiscal shocks in structural VAR models, and the implication of the casual ordering of the model variables.

[37] The importance of the interaction between an active fiscal policy along with a passive monetary policy was recently emphasized by Davig and Leeper (2009) to generate positive government spending multipliers. In this perspective, our evidence of appreciable large expansion in regular GDP subsequent to a fiscal stimulus might further support the evidence of Fig.10.

Figure 10. The response of public spending (right) and the tax rate (left) to each other's one S.D. shock

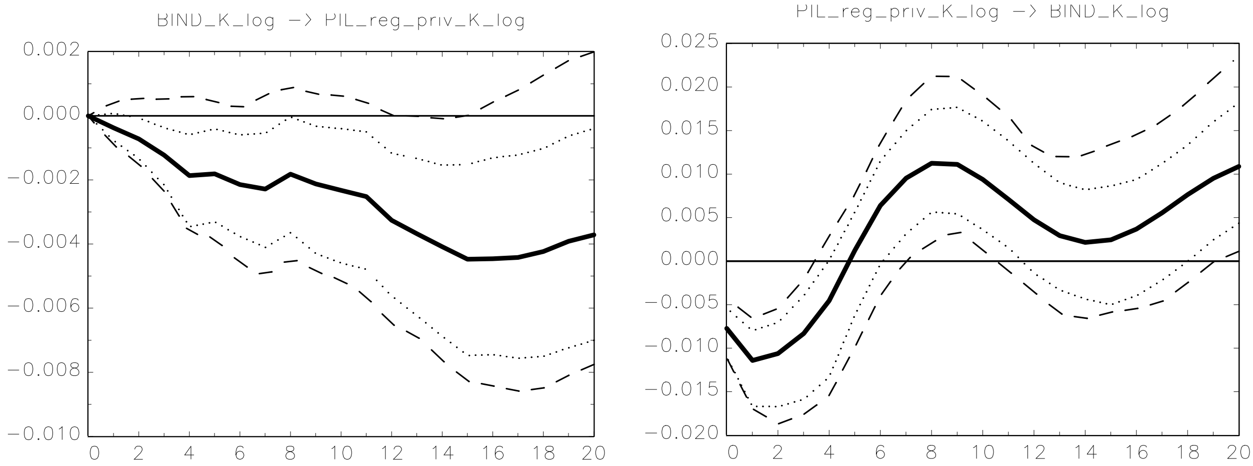


4.5 Unreported and regular GDP: some dynamic aspects

The role of the underground economy and tax evasion in relation to economic growth and the business cycle has attracted extensive attention.[38] Though we are dealing with a statistical model, the data generating process here examined underlines two dynamic aspects of the GDP-tax evasion relationship (Figure 11). First, there is strong evidence for a rise in unreported production to be counterproductive for the regular economy; this can be considered a constraint to the rapid growth of the Italian economy, probably as it generates higher tax rates, fiscal disequilibrium and inequality. Second, when looking at the reverse causality, there is significant evidence of a reduction in unreported production in the short run after a positive shock in the regular economy, approximately for one year, followed by a pro-cyclical pattern for the subsequent years. Thus, after an initial substitution effect between sectors (regular and hidden) there follows substantial complementarity between them, which would indicate that tax evasion is a very structured phenomenon in the Italian economy.

[38] See, amongst others, Busato and Chiarini (2004), Chen (2003) and Onnis and Tirelli (2010).

Figure 11. The response of regular (left) and unreported production (right) to each other's one S.D. shock



This is a striking feature, suggesting that standard models (both theoretical and empirical) omit important aspects of the real world. For example, in models with underground sectors, firms and consumers may be more willing to shift resources out of market activity in response to productivity and policy disturbances than in models without such a sector.[39] By neglecting the economic structure, such models inevitably disregard the implication of resource reallocation between regular and underground production. For instance, Busato and Chiarini (2004) show that underground activities help to mitigate recessions and the cost of high tax burdens by allowing the household to smooth consumption through a proper labor allocation between the two sectors. Our evidence concerning the tax multiplier is consistent with this intuition, since an increase in unreported activities thwarts, to some degree, the recession effects of a rise in taxation. The negative impact of higher taxation on output and income induces firms and households to work more in the underground sector, while highlighting a strong reallocation effect between the two sectors. Figure 7 and Table 1 report that an aggregate shock on taxes pushes the economy into a recession, which is, however, mitigated by the existence of informal production (an underground sector) that offers an *insurance or risk-sharing opportunity*, through labor reallocation toward that sector. As depicted in the right-hand plot of Figure 11, these short-run effects should, however, be considered with the long-run constraint to growth that the informal economy and tax evasion impose on GDP.

[39] The theoretical implications may be relevant here. In a model with a sizeable informal economy and tax evasion, although productivity and policy shocks may affect total hours worked, they may also affect how hours are allocated between the regular and underground sector. The structure of the economy and its parameters that drive the agent's willingness to shift resources between sectors, such as the burden of tax, is relevant to determining aggregate fluctuations. For instance, augmenting the standard stochastic growth model with the informal sector would lead firms and households, like their real world counterparts, to be affected by work and revenue incentives, and therefore push them to undertake intra-temporal allocations of their resources in addition to intertemporal substitutions. See, amongst others, Busato and Chiarini (2004).

5. Concluding remarks and policy issues

By employing a VECM and breaking up GDP into its regular and hidden components we provided striking evidence for Italy on the consequences of a discretionary increase in government spending and tax rate. The relationships between fiscal variables and tax evasion were also investigated.
The effects of fiscal policy in Italy were discussed by referring to a database containing quarterly fiscal variables for the period 1981:1-2006:4. There were two novelties in our data set. First, we elaborated and used long-run quarterly government accrual estimates by making use of a dynamic extension of the disaggregation method currently applied by ISTAT to compute estimates of Quarterly Economic Accounts. Second, we exploited the new yearly time series estimate of the unreported VAT base provided by the Italian Revenue Agency, supplying an accurate quarterly and long enough time series estimate of unreported production. This allowed us to construct and use two important GDP measures in the models: the GDP net of government expenditure referred to as *private GDP*, and *regular GDP* defined as GDP net of government expenditure and evaded VAT base (unreported production).

Decomposing the aggregate production provided us with our first result: we found that the interaction between regular and unreported production demonstrates that the link between the two sectors is very robust, but also very harmful in the long term, since there is strong evidence that shocks to unreported production have long-lasting negative effects for the regular economy. This result warns that any policy which is permissive toward informal production, for its cyclical insurance features, would be a dangerous distraction from growth of the regular economy and fiscal reform plans.

Further important empirical information provided by the model is the estimated equilibrium relationships among the variables and the adjustment to equilibrium following a shock. The key aspect of the cointegration space is that identification of the stationary relationship between public spending and the effective tax rate necessarily requires tax evasion as a further variable to determine a stationary steady state.

Returning to the effects of fiscal policy, our results show that in economies with a sizeable underground sector and tax evasion, the standard aggregate estimate of fiscal effects is not reliable. We found that changes in public spending generate a reallocation from underground to the regular economy with regular GDP which increases more than proportionately to government spending and the underground component of the economy which shrinks. This reallocation effect of the output components contributes to obscure the spending effect on total GDP. Similarly, an increase in the tax rate boosts the hidden economy and tax evasion, entailing a reallocation among sectors which blurs, at best, the total effect. To see the real effect of fiscal policy, we need to decompose regular and hidden production. In this setup the spending multiplier peaks after one year at +3.7% and shows large long-run effects. There is a sizeable short-run drop in regular output (-1.1% after one year) after an increase in the tax rate, and it is associated with a considerable increase in tax evasion.

A question that emerges strongly from these findings is why such striking fiscal multipliers are unable to spur Italian growth. Governments should massively seek to boost the economy with hefty expenditure and reduce the informal component. Our explanation about the awkward gap between the model's predictions and the disappointing growth performance of Italy is based on the following considerations. The first fact to emphasize is that the informal component curbs long-term growth, providing short-run insurance against depression but also mitigating the effects of expansive policies. Public expenditure increases total GDP and sharply reduces its underground component: following an

upward adjustment of expenditure by one percent of total GDP, after one year, regular GDP adjusted by 3.7 points but the reallocation of resources in the informal economy amounted to -6%, with an overall expansionary effect of +0.9 % on total GDP.
A second element to consider is that our impulse response functions were calibrated on a spending shock equal to one percentage point of GDP. In the historical period examined, our measure for public spending (consumption plus investment) grew at a quarterly average rate of 0.5%, which is tantamount to approximately 0.1% of GDP. Clearly, the shock intensities simulated in our Table 1 and 2 are thus well above the average rates of change recorded (by one order of magnitude). Finally, the evidence presented so far is clearly at odds with standard fiscal policy prescriptions: it is hard to justify on economic grounds the increase in public spending subsequent to a positive shock in regular production, and it is similarly difficult to defend the decline in tax rates following the increase in public spending.

This puzzling behavior can be traced back to the institutional context, which leaves room for political meddling, influencing spending decisions and tax policy. In the sample considered, public expenditure was intimately linked to political processes, generating a substantial fiscal disequilibrium. Moreover, the problem is that public spending is fixed, difficult to reduce, and an increase in GDP encourages new spending (a particularly robust result in all simulations attempted). Following Talvi and Vegh (2005) the political distortion can be characterized by an endogenous component in government spending, such that decisions on public spending are positively related to the primary surplus. This is consistent with our model prediction about the effects of a positive shock to the regular economy, which leads political leaders, with a short-run horizon, to envisage an increase in tax revenues and hence relax fiscal discipline. Then in good times we will observe an increase in spending, and a procyclical pattern for public surpluses.

When considering the revenue side, the decline in the tax rate following an exogenous increase in government spending suggests that fiscal policy instruments managed to (de)stabilize the economy: an expansionary fiscal policy on the spending side is accompanied by an expansionary policy on the revenue side as well (possibly an increase in transfers for a given tax rate).

Overall, these mechanisms have led to a continuous upward trend in public debt, and provide a typical result found by many authors for many developing countries, that is a procyclical and myopic fiscal policy: when lobbies and unions realize that positive income shocks hit the economy, they immediately apply pressure to obtain an increase in public goods and transfers (see, amongst others, Alesina et al., 2008; Talvi and Vegh, 2005; Battagliani and Coate, 2008).

Which role has tax evasion played in this economy? It is probable that the Italian strong downward rigidity of public spending (that is, it is ungovernable) has led all governments to opt for the easiest way out: a policy of high taxation in the presence of large tax evasion. Moreover, tax evasion has contributed to smooth short-run fluctuations in total GDP (recall that the main criticism leveled at pro-cyclical fiscal policy is excess volatility in output), although it has also caused a downward bias in economic growth.

This paper constituted the first step in an ongoing research project. Natural extensions to search for the transmission mechanism of fiscal policy include the decomposition of both public expenditure and the GDP in consumption and investment components, and analysis of labor market reactions (also using the structural vector autoregression approach to handle sample size restrictions). Finally, following

Favero and Giavazzi (2007), we should consider public debt as an additional endogenous variable in the VAR, in order to fully appreciate the debt dynamics that arise after a fiscal shock and, more importantly, study how taxes, public spending and tax evasion respond to the level of debt.

APPENDIX

Appendix A: Quarterly estimate for the evaded VAT base

In this paper we use quarterly time series for the unreported VAT tax base and VAT tax evasion for Italy over the period 1980-2006, based on quarterly economic and fiscal information from national accounts. As explained in Section 2, yearly data on the unreported VAT tax base were recently estimated for Italy by the Revenue Agency of the Ministry of Economy and Finance over the period 1980-2006. According to Marigliani and Pisani (2007), undeclared VAT taxable amounts are assessed by comparing *theoretical* data (Total VAT base) and actual observations (Declared VAT base) derived from VAT returns.

In order to obtain the quarterly series of the undeclared VAT base we first applied the temporal disaggregation procedure suggested in Proietti (2006) to the yearly series of the Total VAT base and Declared VAT base from the Revenue Agency. Subsequently, we found the quarterly series of the undeclared VAT base as the difference between the two in the same fashion of Marigliani and Pisani (2007).

It is worth spending some words in describing the main characteristics of the temporal disaggregation method used to obtain quarterly estimates for the two mentioned series. The method used consists of a dynamic extension of standard techniques widely used by national statistics institutions to estimate high frequency (monthly or quarterly) series. Indeed, many European national statistics offices make use of indirect methods relying on proxy indicators to compute time series of Quarterly National Economic Accounts. In particular, ISTAT has been using the Chow-Lin (1971) method since the 1980s. The method assumes the existence of a multiple regression model at high frequency *t*:

$$\begin{aligned} y_t &= x_t'\beta + u_t \qquad && t = 1,\dots,n. \\ u_t &= \phi u_{t-1} + \varepsilon_t \qquad && \varepsilon_t : \; NID(0,\sigma^2) \quad -1 < \phi < 1 \end{aligned} \tag{A.1}$$

In our case *t* refers to a quarter. The unknown quarterly variable y_t , of which only yearly observations are available, is regressed on k proxy indicators which can be observed at quarterly frequency and are included in vector x_t'. The model, as presented in equation (A.1), assumes quarterly residuals are generated by a first-order Markov model without drift. The model cannot be estimated at quarterly level due to the lack of information on variable y_t and hence estimation is to be performed at yearly frequency. The model is aggregated by multiplying both sides of the equation referring to y_t for a suitable aggregation matrix. In particular, as dealing with flow variables, the model is aggregated by using a matrix which allows quarters for each year to be summed. As a consequence of this time constraint, the sum of computed quarterly estimates of variable y_t in each year is equal to the corresponding yearly observation, so that the mean disaggregation error is zero by construction.

After the revision of national accounts in 2005, ISTAT added the Fernàndez (1981) method to the Chow-Lin benchmark. Unlike the Chow-Lin method, this more recent model implicitly assumes the y_t series and the k indicators are not cointegrated, or if they are, that the model is not the cointegration one. In line with this assumption, quarterly residuals are supposed to be generated by a random walk by setting ϕ =1 in equation (A.1).

Both the methods provide quarterly BLU estimates which consist of two components: the first includes the effects of indicators, while the second consists of the so-called *adjustment part*, which operates the smoothing of the estimated yearly residuals across the corresponding quarters. Different hypotheses on residuals lead to different shapes of the smoothing matrix, which is a function of parameter ϕ, and thus lead to different quarterly estimates.
The broader Autoregressive Distributed Lag (ADL) model of first order in equation (A.2) nests both the Chow-Lin and the Fernandez methods:[40]

$$\begin{aligned} & y_t = \phi y_{t-1} + x_t^{'}\beta_0 + x_{t-1}^{'}\beta_1 + \varepsilon_t \qquad t = 1,...,n. \\ & \varepsilon_t : \ NID(0,\sigma^2) \end{aligned} \tag{A.2}$$

The two static aforementioned models can be considered as a restricted version of the ADL model. It can be shown that under the following restriction:

$$\beta_1 = -\phi\beta_0 \tag{A.3}$$

the ADL model yields the Chow-Lin model while, under both the restrictions in equation (A.3) and (A.4), the model yields the Fernàndez approach:

$$\phi = 1 \tag{A.4}$$

The estimation of the broader model instead of the Chow-Lin version avoids any potential biases in the coefficients of indicators whenever the common root restriction in (A.3) does not hold. Use of a more general specification thus allows selection of the proper model to describe the relationship between the series to be disaggregated and indicators.
The approach suggested in Proietti (2006) also proposes a new estimation procedure which relies on the state space representation of the ADL model, allowing the distribution problem to be transformed into one of unknown observation. The procedure also provides the computation of the standardized Kalman Filter innovations which are useful for checking the statistical properties of quarterly estimates. In line with recent practice adopted by ISTAT, parameters are computed by using the maximization of likelihood function.[41]
Our application consists in a simplified and more parsimonious version of the general procedure considering k indicators, as we rely on a single indicator selected according to the economic meaning of variables.
In order to disaggregate the yearly time series of the Total VAT base we use a quarterly indicator computed as the sum of GDP and net imports. The indicator assesses the amount of economic resources produced at national level and it is thought to be a good proxy for the theoretical VAT base. In the case of the Declared VAT base we used the quarterly series of net indirect taxes (net of

[40] As shown in Proietti (2006), when the ADL model is reformulated according to first differences of y_t and x_t it yields the Litterman (1985) model, which is the third model included in the so-called Chow-Lin family.
[41] After the 2005 review, ISTAT moved from the Barbone et al. (1981) estimation strategy, which relied on feasible generalized least squares, to the maximum likelihood estimation, which is still performed within an iterative grid search procedure to estimate the first autoregressive parameter in the restricted range.

contributions on production), which includes actual VAT returns. The quarterly indicators are from the Con-Istat quarterly database, containing Quarterly National Accounts time series. Data, which are available from the first quarter of 1980, are at nominal prices and seasonally adjusted with trading day correction at source.

For each yearly series we estimated three models: i) the Autoregressive Distributed Lag model (henceforth ADL) in equation (A.2), with the addition of a constant term and a linear trend, ii) the nested Chow-Lin model (henceforth CL) found by imposing restriction (A.3) into the ADL specification, with the addition of a constant term and a linear trend, iii) the Fernàndez model (henceforth FE) found by imposing restrictions (A.3) and (A.4) into the ADL specification. In the latter case, the model is estimated by using first differences, considering only the constant term.

The selection of the proper model is based on the stochastic properties of smoothed estimates derived from their standardized Kalman Filter innovations. The latter are computed by the filter as the one-step ahead forecast error. Standardization is made on the basis of the standard deviation.[42] Although filtering is useful for diagnostics, we consider smoothed estimates in order to exploit the whole information contained in the estimation set.

Further, we base the selection of the proper specification on the ability of the models to minimize the revision error associated to quarterly estimates. According to the European System of National Accounts, the revision error can be defined as the difference between quarterly extrapolations for year *t+1*, obtained by the model using information up to time *t*, and the revised quarterly estimates computed once the *t+1* yearly observation becomes available. To this end, we performed a rolling forecast experiment consisting in re-estimating the three models with the addition of one more observation at a time, starting from 1992 up to 2006. At the beginning of each year we made dynamic projections for the four subsequent quarters by using only the observations up to the previous year. We then computed the updated quarterly estimates by adding the subsequent yearly observation to the estimation set. In this way, we obtained two values at each quarter over the period 1993-2006: projections, based on old yearly information, and the revised estimates incorporating the new yearly observation. We computed the revision error at each quarter as the difference between the two values. Afterwards, we took the mean for each quarter over the 14-year-period, by obtaining four mean revision errors, assessing respectively the error occurring on average in the first, second, third and fourth quarter.

The main results are summarized in the following three tables. As regards the Total VAT base, Table A.1 shows that the quarterly indicator referring to the amount of economic resources exerts significant effects both at contemporaneous and one-period lagged.

[42] For a more detailed discussion on the topic see Proietti (2006).

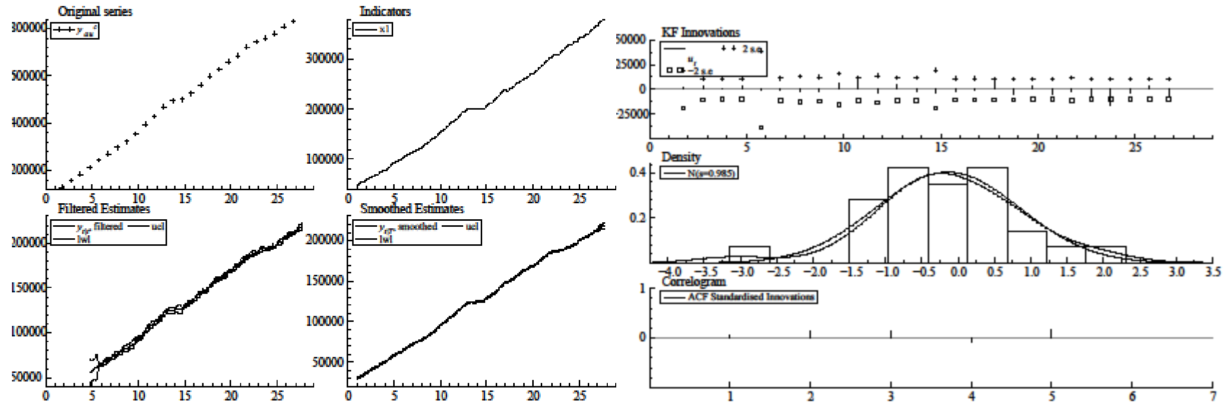
Table A.1 Total VAT base, main parameter estimates

Total VAT base				
	Main parameter estimates			
Method	β_0	β_1	ϕ	$-\phi*\beta 0$
	(Student's t in brackets)			
Adl	0.42 (5.71)	-0.41 (-5.74)	0.97	-0.41
Cl	0.39 (4.97)	…	0.97	
Fe	0.39 (5.19)	…	1.00	

As expected, regression effects are always positive at contemporaneous dates, and negative one-period lagged in the ADL model. Results also suggest the common root restriction in equation (A.3) can be accepted, and that therefore the use of the Chow-Lin or the Fernandez model would be advisable. The shift to restricted versions allows a gain in efficiency due to the reduction in the number of parameters to be estimated.
No major differences can, however, be observed in estimating the three models (ADL, CL and FE), even when considering their growth rates. Smoothed estimates for the FE model are presented in the left-hand panel of figure A.1.

Figure A.1 Total VAT base, Fernàndez model

Panel a | Panel b



According to results from the diagnostic check in the right-hand panel of figure A.1, the estimates show quite good statistical properties. The standardized Kalman Filter innovations can be considered a white noise process; innovations show a symmetric distribution without any problem of Kurtosis and very close to a Normal one. Finally, autocorrelation seems to be absent.
We considered the results obtained from the rolling estimation in order to identify the method which is able to provide the best extrapolation and the minimum mean square revision error.

Table A.2 Total VAT base, revision errors

	1Q	2Q	3Q	4Q
Mean percentage revision error				
Adl	0.11	-0.33	-0.14	-0.23
Cl	0.04	-0.43	-0.27	-0.30
Fe	-0.19	-0.27	-0.33	-0.37
Mean revision Error				
Adl	12.52	-711.11	-401.98	-594.11
Cl	-137.87	-822.82	-659.78	-689.71
Fe	-364.25	-527.17	-647.40	-725.70
Mean absolute revision error				
Adl	1,429.51	1,574.10	1,624.55	1,490.39
Cl	1,307.49	1,265.86	1,543.88	1,474.59
Fe	768.05	1,126.35	1,374.12	1,522.14
Mean root square revision error				
Adl	2,235.36	2,163.27	2,300.88	2,130.05
Cl	1,617.92	1,870.89	2,116.73	2,079.57
Fe	988.79	1,443.93	1,750.97	1,930.33

As shown in table A.2, the FE model is the one whose mean root square revision error is the lowest over quarters. The corresponding smoothed estimates are therefore considered the most correct for constructing the evaded VAT base variable.
As regards the Declared VAT base, estimation showed less satisfactory results than those obtained for the Total VAT base. As expected, regression effects are positive but the quarterly indicator referring to net indirect taxes does not seems to exert statistically significant effects on the variable to disaggregate, especially in the ADL model, as shown by Student's t in Table A.3.

Table A.3 Total VAT base, main parameter estimates

Declared VAT base

Method	β_0	β_1	ϕ	$-\phi*\beta 0$
	(Student's t in brackets)			
Adl	0.13 (0.13)	0.24 (0.25)	0.63	-0.08
Cl	0.88 (1.59)	…	0.69	
Fe	1.45 (1.96)	…	1.00	

The common root condition can be accepted and the CL model can be considered the proper specification instead of the broader ADL specification. For a two-side Student's t test at the 5 per cent level of confidence and with 21 degrees of freedom, the critical value is 2.08. The standard error of β_1 being 0.96, the t statistic for β_1 is:

$$t_{(25-4)} = \frac{0.24 + 0.08}{0.96} = 0.33$$

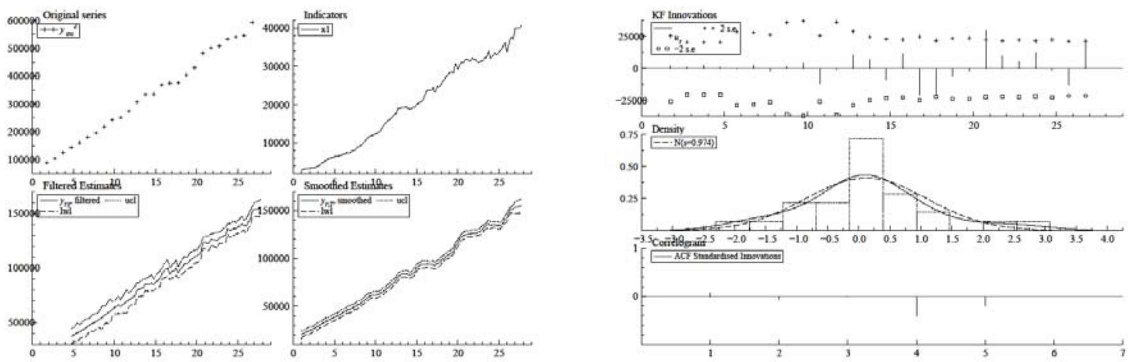
The test clearly suggests accepting the H_0 hypothesis of $\beta_1 = -\varphi*\beta_0$.

However, even when considering the CL model, the statistical significance of the quarterly indicator is not accepted either at the 5 or 10 per cent level. Differently, β_0 is statistically significant at the 10 per cent level in the FE model.

Panel a of Figure A.3 shows smoothed estimates deriving from the CL model and their statistical properties. There clearly emerges a larger confidence interval around estimates than that observed in the previous case (see the lower left-hand panel in Figure A.1 for a comparison). As a major consequence, the smoothed estimates of the Declared VAT base are less precise than that found for the Total VAT base.

Figure A.2 Declared VAT base, Chow-Lin model

Panel a Panel b



However, compared with the estimates found with the ADL and the FE method, the estimates from the CL model are the most precise with the lowest root mean square error.
Results from diagnostic checking on the standardized Kalman Filter innovations suggest the innovations present a symmetric distribution and no significant autocorrelation problems.
The rolling experiment provides additional information. It clearly suggests that the ADL model underperforms with respect to its restricted versions. Also, the comparison between the CL and the FE is not straightforward, except when considering the root mean square revision error. As shown in Table A.4, according to the latter indicator, the CL model on average outperforms that of FE in the second, third and fourth quarters of predictions.

Tab A.4 Declared VAT base, revision histories

	1Q	2Q	3Q	4Q
Mean percentage revision error				
Adl	0.31	0.76	1.27	0.73
Cl	-0.02	0.81	0.64	0.34
Fe	0.16	0.16	0.25	0.33
Mean revision Error				
Adl	-65.54	900.32	1,148.05	710.37
Cl	-160.95	1,001.85	736.09	426.91
Fe	165.88	167.73	273.21	350.29
Mean absolute revision error				
Adl	5,204.70	3,520.27	4,052.53	4,390.08
Cl	4,154.74	2,005.62	3,077.61	3,485.23
Fe	1,961.29	2,862.85	3,378.28	3,712.62
Mean root square revision error				
Adl	7,271.97	4,956.71	5,605.73	6,138.85
Cl	5,421.41	2,933.66	3,733.86	4,289.52
Fe	2,652.06	3,871.95	4,546.15	4,931.49

Hence, as a conclusion of the analysis, we consider the smoothed estimates corresponding to the Chow-Lin model as the correct for constructing the declared VAT base.

As further proof of the reliability of data, we also tried to disaggregate the series of Declared VAT by considering two indicators, net indirect taxes and the ratio of employees over self-employed. The ratio of employees over self-employed is computed by using data from the CON-ISTAT quarterly database and refers to the total economy; it is expressed in full-time equivalent and seasonally-adjusted by the source.

First, the inclusion of this second indicator strengthens regression effects exerted by net indirect taxes in the Fernàndez model: the Student's t associated to contemporaneous effects goes up from 1.96 to 2.26 as shown in Table A.5. Estimation effects of indirect taxes are again not significant either at 5 or 10 per cent in the other two models under consideration (ADL and CL).

The ratio of employees to self-employed exerts positive effects on the declared VAT base as expected. The main parameter estimates shown in Table A.5 suggest the existence of significant contemporaneous and one-period lag effects in the Fernàndez model and significant contemporaneous effects in the ADL model, but only at 0.90 per cent of confidence (Student's t with 21 degrees of freedom is equal to 1.725 in a two-sided alternative test).

Tab A.5 Declared VAT base, main parameters using two indicators

Declared VAT base

Main parameter estimates

	x1		x2			x1	x2
Method	β_0	β_1	β_0	β_1	ϕ	$-\phi*\beta 0$	$-\phi*\beta 0$
	(Student's t in brackets)		(Student's t in brackets)				
Adl	0.63	-0.30	3.66	-3.43	0.67	-0.42	-2.46
	(0.71)	(-0.34)	(1.81)	(-1.61)			
Cl	0.72	…	1.70		0.70		
	(1.33)		(1.64)				
Fe	1.58	…	2.83		1.00		
	(2.26)		(1.99)				

Legend: x1= net indirect taxes; x2= ratio of employees to self-employed.

As before, the common root condition can be accepted by the data. In the case of net indirect taxes, the standard error of β_1 is equal to 0.88, yielding the following t statistics:

$$t_{(25-6)} = \frac{-0.30 + 0.42}{0.88} = 0.1363$$

The test clearly suggests that the H_0 hypothesis of $\beta_1 = -\varphi^*\beta_0$ be accepted even at the 0.99 per cent of confidence. In the case of employees over self-employed, estimation results yield the following t statistics:

$$t_{(25-6)} = \frac{-3.43 + 2.46}{2.136} = 0.454$$

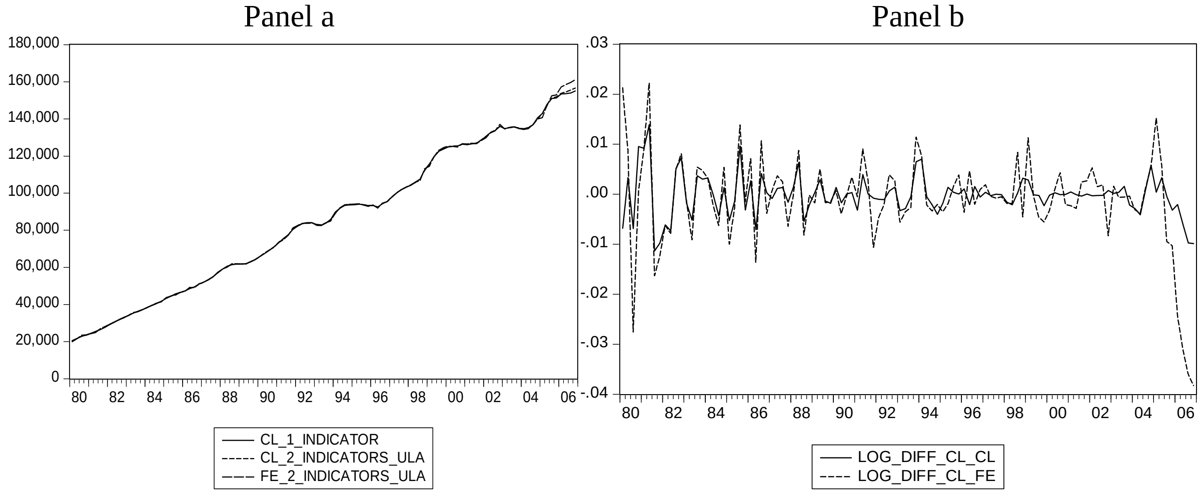
The test confirms the null hypothesis at 0.99 per cent of confidence as well and we can still consider the Chow-Lin as the proper model instead of the broader ADL specification.
According to the rolling experiment, CL estimates still out-perform those obtained from the ADL and FE models, as shown in Table A.5. Hence estimation results lead us to choose estimates obtained with the Chow-Lin model.

Tab A.6 Declared VAT base, revision histories using two indicators

	1Q	2Q	3Q	4Q
Mean percentage revision error				
Adl	0.55	0.82	1.93	1.14
Cl	0.42	0.65	0.85	0.52
Fe	0.11	0.12	0.24	0.33
Mean revision Error				
Adl	23	982	1673	863
Cl	315	786	927	580
Fe	93	107	245	329
Mean absolute revision error				
Adl	5500	3705	4782	5262
Cl	3574	1853	2865	3254
Fe	1876	2783	3323	3624
Mean root square revision error				
Adl	781.552	522.211	643.600	757.973
Cl	489.083	276.940	355.843	416.536
Fe	250.581	375.165	447.452	482.490

However, comparison of estimates obtained with the Chow-Lin method using one indicator and the new ones based on two indicators (net indirect taxes and employees to self-employed ratio) clearly shows that no significant differences emerge between the two time series, as shown in panel a of Figure A.3.

Figure A.3 Declared VAT base, comparison



Panel b shows log-differences between the two series, represented by the solid line. No systematic difference emerges between the two series (the mean is zero) and differences also vary within a very small range: from a minimum of about -0.01 per cent to +0.01 per cent.
In Figure A.3 we also plotted the estimates obtained by using the Fernàndez method with two indicators and found that major differences occur at the end of the estimation sample. The series called “log_diff_cl_fe” in panel b shows log-differences between CL estimates based on one indicator and FE estimates based on two indicators. Also in this case no systematic difference emerges, even if differences show a slightly higher variation range, from -0.04 to 0.02 per cent.
As a final result, we do not expect the use of estimates based on two indicators for the Declared VAT base to be able to make major differences in the analysis. Therefore, as already stated, we consider estimates obtained with the Chow-Lin method and one indicator as the correct procedure for constructing the declared VAT base.

Appendix B: Quarterly fiscal data

Since the introduction of the System of National Accounts, the *accrual principle* for measuring fiscal quantities has been favored by national statistic institutes belonging to UN. The International Monetary Fund and EUROSTAT consider the accrual principle as the appropriate one to handle fiscal variables, because committed expenditures and assessed revenues might have a greater effect on the behavior of economic agents than receipts and payments.

Moreover, models with rational agents show that private sector behavior is not only influenced by cash-basis government finances, but also by expenses and revenues that will be paid or cashed in the future, throughout the anticipation of their effects. Also, the recent empirical literature on macroeconomic effects of fiscal policy has pointed out that the reliability of results is strongly affected by the use of proper fiscal policy indicators, suggesting the use of long-lasting quarterly national accounts data for the general government sector. Consistent with this orientation, and due to the shortage of genuine quarterly national account government data in many countries, most of the literature on macroeconomic effects of fiscal policy has extensively focused on the US and, to a lesser extent, on the UK, Canada, Western Germany and Australia.

The issue is important for Italy, as well as many other countries, where an information gap seems to emerge both at yearly and quarterly frequency. Yearly general government national accounts data based on the European System of Accounts (henceforth ESA95) have been issued by ISTAT in the *Conto Economico Consolidato delle Amministrazioni Pubbliche* (henceforth CECAP), but only since 1980. Quarterly general government data based on ESA95 are published in the *Conto Economico Trimestrale delle Amministrazioni Pubbliche* (henceforth CETAP). However, these fiscal time series have only been available since the first quarter of 1999.

Six variables, referring to direct and indirect taxes, social contributions and social payments have, instead, been published by EUROSTAT since the first quarter of 1991 and can therefore be found in the EUROSTAT NEW-CHRONOS database, which actually extends public finance information from the CETAP account.[43] At a second-level classification, neither ISTAT nor EUROSTAT provide quarterly national accounts data for current taxes on household (direct taxes on labor) and on corporate income (direct taxes on profits).

Thus, an empirical analysis relying on official quarterly general government data, compiled on a national accounts basis, would be feasible only if one considered a few budgetary items, referring to the non-market sector for the spending side and revenues from VAT and other taxes on imports. These time series are available back to 1980:I. As a consequence of this information gap, empirical applications for Italy have relied on either OECD quarterly general government data or quarterly cash-basis data. Two kinds of problems arise, however, when using these figures. On the one hand, OECD quarterly government time series are interpolated without a guide; on the other, cash principle does not seem to be considered as the proper one according to recent international opinion on government accountancy.

The first issue, regarding OECD high frequency data, deals with temporal disaggregation methods. There seems to be quite a broad consensus on the recent criticism that OECD quarterly data are no more informative than annual data.[44] When no official long-run high frequency data are available from

[43] The items refer to current taxes on income and wealth, taxes on production and imports, value-added type taxes, capital taxes receivable, actual social contributions receivable (excluding imputed social contributions) and social payments (social benefits other than social transfers in kind).

[44] See, for instance, Perotti (2004).

the national source, the OECD computes quarterly estimates by interpolating yearly data without any guide, as in the case of Italian quarterly government accounts. The model used to perform temporal disaggregation is exclusively based on mathematical criteria aimed at avoiding unrealistic jumps when passing from one year to another.
With regard to cash-basis data, according to many authors, such data should be considered an incomplete guide for the measurement of counter-cyclical activity effects of the government sector on the private one.[45] Another stream of literature, instead, emphasizes that payments would produce greater effects on private consumption than spending commitments.[46] A recent study of the Bank of Italy (Giordano et al. 2007) relies on quarterly cash-basis data of the public sector. Data for the consolidated public sector are based on genuine cash observations referring to the State sector, local governments, health sector and social security institutions published by the Italian Ministry of Economic and Finance. For the period 1982-1993, the above authors performed data consolidation for the public sector, consolidating intergovernmental flows when possible. On the spending side, only government consumption spending is considered.
As explained in Section 2 of the paper, one of the novelties of this study refers to the use of fiscal policy variables based on national accounting definitions. The estimation of the VECM model relies upon quarterly general government (henceforth government) national accounts estimates, namely government direct spending and net revenues. The two series are taken from a broader set of quarterly government estimates which has been recently constructed for Italy over the period 1980:1- 2007:4.[47] The quarterly database aims to fill the above information gap and provides fiscal policy data fully comparable with statistics from the major industrialized countries for their institutional coverage and accountancy basis.
In the following we briefly report the main characteristics of the temporal disaggregation method used to obtain quarterly estimates for government primary spending and net revenues. The quarterly time series were estimated starting from yearly observations of the CECAP account, which represents the information basis for calculating the net lending requirement defined by the Maastricht Treaty, over the period 1980:1 – 2007:4. The temporal disaggregation method used consists in the dynamic extension of the Chow-Lin procedure suggested in Proietti (2006). The procedure relies on the estimation of a first order Autoregressive Distributed Lag (ADL) model, described in detail in Appendix A, which embeds the three univariate models constituting the Chow-Lin classic approach, namely the Chow-Lin, Fernàndez and Litterman models. The fiscal variables we selected are widely used in the empirical literature on discretional/non-systematic fiscal policy. On the spending side, we focus on direct public spending on goods and services. Public transfers are, however, considered for the definition of net revenues.
The source of quarterly proxy indicators used to derive the unknown pattern of government primary spending and net revenues is the CON-ISTAT quarterly database, containing data from Quarterly National Accounts. For each yearly series we selected one related time series by taking SEC definitions into consideration: quarterly indicators of the expenditure side are supposed to include the unknown government quarterly time series, whereas the quarterly indicators of the revenues side are supposed to replicate the unknown pattern of corresponding taxation base.
Table B.1 provides insight into the input data considered.

[45] Levin (1993).
[46] Auerbach, William (2009).
[47] For more details about the construction of the quarterly dataset see Basile (2009) and Basile, De Arcangelis and Proietti (2010).

Tab B.1 Input data for the construction of quarterly fiscal policy estimates

YEARLY SERIES	QUARTERLY INDICATORS[1]
Gov. consumption expenditure	Consumption expenditure of non-market sector
Gov. fixed capital formation (gross of divestments)	Fixed capital formation of total economy
Current taxes on income and wealth	The sum of estimates for the items of Current taxes on corporate income and Current taxes on household income
Social contributions receivable	Compensations of employees less gross wages and salaries of total economy
Taxes on production and imports	VAT and other taxes on imports
Total revenues net of transfers to households and firms	The sum of estimates for the items of Current tax on income and wealth, Social contributions and Taxes on production and imports

1) Seasonally-adjusted with trading day correction at the source.

-Government spending

Quarterly estimates for government direct spending are found as the sum of quarterly estimates for public consumption spending and gross investments. As shown in Table B.1, the disaggregation of government consumption was based on the quarterly series of non-market consumption spending.[48] Obviously, government consumption spending includes wages and non-wages current purchases of goods and services.

As regards public investments, we consider the series of government gross investments not inclusive of earnings from real estate divestments in order to avoid irregular jumps due to such divestments. According to public accountancy rules, indeed, divestment earnings in real estate enter the budget in the item of government fixed capital formation with the minus sign, leading to strong volatility in both official yearly CECAP figures and quarterly CETAP estimates provided by ISTAT. As explained in Basile (2009), inclusion of such extraordinary events in the yearly series seems to be misleading for research purposes, and for the economic meaning of the fiscal variable itself. The literature focuses on public investments, which are mostly determined by long-term funding programmes and not by contingent events. The revised definition of the yearly series leads to a significant gain in terms of precision of estimates referring to public investments and also allows this information to be used in the analysis of fiscal policy. Gross investments excluding divestments in real estate are disaggregated on the basis of total gross fixed capital formation of the whole economy.[49]

[48] According to the ESA95 accounting system, the non-market sector includes both public and private producers whose market revenues are less than half of production costs. Hence the series represents a very good proxy for the government series.

[49] The choice of an indicator for government investments is not easy and, at present, it appears to be controversial. The use of an indicator referring to part of aggregate demand might produce some bias when using government investment estimates to assess their impact on GDP. We thank Sandro Momigliano for some discussion on this issue.

-Net revenues

Net revenues are defined as total revenues less monetary social benefits (including pensions and other assistance monetary transfers), "other current expenditure" (including production subsidies), and "other capital expenditure" (including investment subsidies).[50] The variable is widely used in fiscal policy VAR models and in macroeconomic models with credit constraints where fiscal policy is supposed to produce economic effects via the demand-side channels.[51]

The disaggregation of the net revenues was based on a two-step approach. In the first step, we estimated quarterly series for direct taxes, social contributions and indirect taxes. These estimates were then used to proxy the unknown pattern of the whole amount of net revenues.[52] The quarterly indicator used in the disaggregation is, indeed, found as the sum of the quarterly estimates for current taxes on income and wealth, taxes on production and imports, and social contributions.

The quarterly government estimates used in this analysis benefit from the following main characteristics: i) they are time-consistent with the corresponding official yearly data from ISTAT in the sense that the sum of quarterly data is equal to the corresponding year observation, ii) they are consistent with the economic accrual principle set by ESA95 as both yearly observations and quarterly indicators are complied with according to the economic accrual principle, iii) they are fully comparable with official quarterly national accounts data issued by the national statistics institutes of other industrialized countries.

[50] Interest payments were excluded from the definition of net revenues. See De Castro (2004) amongst others.

[51] See, for example, Blanchard and Perotti (2002).

[52] For more details on the disaggregation results and a comparison of estimates with official quarterly estimates issued by EUROSTAT and ISTAT see Basile (2009) and Basile, De Arcangelis, Proietti (2010), which also provide a robustness proof of the methodology for the USA case.

Appendix C: The regular GDP model

In this appendix we first report the statistical characteristics of the time series used in the models, and then the cointegration analysis and the complete set of the impulse response analysis obtained for the regular GDP model.

C1. Unit Root Tests

Table C.3: Augmented Dickey Fuller Test for unit root. Null Hypothesis: unit root

	Level		First Difference	
	Test statistic	Deterministic (lags)	Test statistic	Deterministic (lags)
Government Primary Spending	-2.11	intercept, time trend (3)	-3.63***	Intercept (3)
Net tax revenues/Regular GDP	-1.97	intercept, time trend (2)	-4.55***	Intercept (1)
Unreported GDP	-3.46**	intercept, time trend (1)	-5.41***	Intercept (0)
Unreported GDP (KPSS)^	0.49***	intercept, time trend (1)		
Total GDP	-1.79	intercept, time trend (4)	-4.55***	Intercept (3)
Private GDP	-1.46	intercept, time trend (4)	-5.19***	Intercept (3)
Regular GDP	-2.63	intercept, time trend (6)	-4.54***	Intercept (5)

Thresholds (constant and trend): 1% (-3.96) 5% (-3.41) 10% (-3.13).
Thresholds (constant): 1% (-3.43) 5% (-2.86) 10% (-2.57).
^KPSS test rejects the null of trend stationarity for the undeclared VAT base at the significance level of 1%.

C2. Cointegration analysis

The deterministic variables included in the Johansen test are:

- one mean-shift dummy describing a regime shift starting in 1992:1. This dummy is restricted to lying in the cointegration space.
- three impulse dummies (describing a permanent intervention/shock) for the observations: 1983:4; 1989:1; 1998:1.
- two transitory shock dummies, for the observations: 2003:3-4 and 2005:4-2006:1.
- Constant.

In addition, the rate of growth of the stock of public debt is added as an exogenous variable

Table C2: Johansen Trace Test for Regular GDP, unreported GDP, public spending, net effective tax rate

R0	LR	Pi value	90%	95%
0	95.81	0.0000	56.97	59.58
1	55.85	0.0000	37.72	39.97
2	**23.51**	**0.066**	**22.36**	24.47

Optimal endogenous lags from information Criteria: 5 (AIC); 2 (Final Prediction Error; Hannan Quinn C.; Schwartz C.)

C3. The VECM analysis

We briefly describe the main diagnostic tests for the benchmark model (Regular GDP, Unreported GDP; Public Spending; Net Effective Tax Rate), for which we also plot the complete Impulse Responses. Further elaborations are available upon request.

The order selection criteria suggest using four lags. Hence the model is based on a sample period going from 1982:3 to 2006:4 (98 observations). We use a two-stage estimation procedure: in the first stage the cointegration matrix has to be estimated by the S2S method explained in Lutkepohl (2004). Once an identified form of the estimated cointegration matrix is available, it can be used in the second stage of the estimation procedure. In the second stage, structural and subset restrictions as well as exogenous variables can be accounted for.

The rate of growth of public debt is included in the VECM specifications as an exogenous variable, while the deterministic variables are:

- one mean-shift dummy describing a regime shift starting in 1992:1. This dummy is restricted to lying in the cointegration space. Its difference should be included as an unrestricted permanent impulse dummy in the VAR equations, current and lagged,
- three impulse dummies (describing a permanent intervention/shock) for the observations: 1983:4; 1989:1; 1998:1. A shift in the levels of a variable becomes an impulse dummy in the differenced variable,
- two transitory shock dummies, for the observations: 2003:3-4 and 2005:4-2006:1,
- a further impulse dummy, not linked to the break in the data though accounting for an outlier for the tax rate equation in the VECM, is added for the observation 1996:4. This dummy is necessary in order to provide normally distributed residuals for the tax rate equation in the VECM,
- constant.

C3.1 Residuals analysis: diagnostic tests

Table C3: residual autocorrelation tests for VEC Benchmark Model

Test	$ARCH\_LM_{16}$ (u_1)	$ARCH\_LM_{16}$ (u_2)	$ARCH\_LM_{16}$ (u_3)	$ARCH\_LM_{16}$ (u_4)	LM_5	Multiv. $ARCH\_LM_5$
Test statistic	19.59	22.95	16.81	9.27	73.48	509.57
p-value	0.29	0.11	0.40	0.90	0.68	0.37

Table C4: residual non-normality tests for VEC Benchmark Model

Test	Doornik Hansen(1994)			Lutkepohl (1993)		
	joint	skewness	kurtosis	joint	skewness	kurtosis
Test statistic	3.42	3.09	0.32	2.19	1.06	1.12
p-value	0.90	0.254	0.99	0.97	0.89	0.89

Table C4: residual non-normality tests for VEC Model 2 (follows)

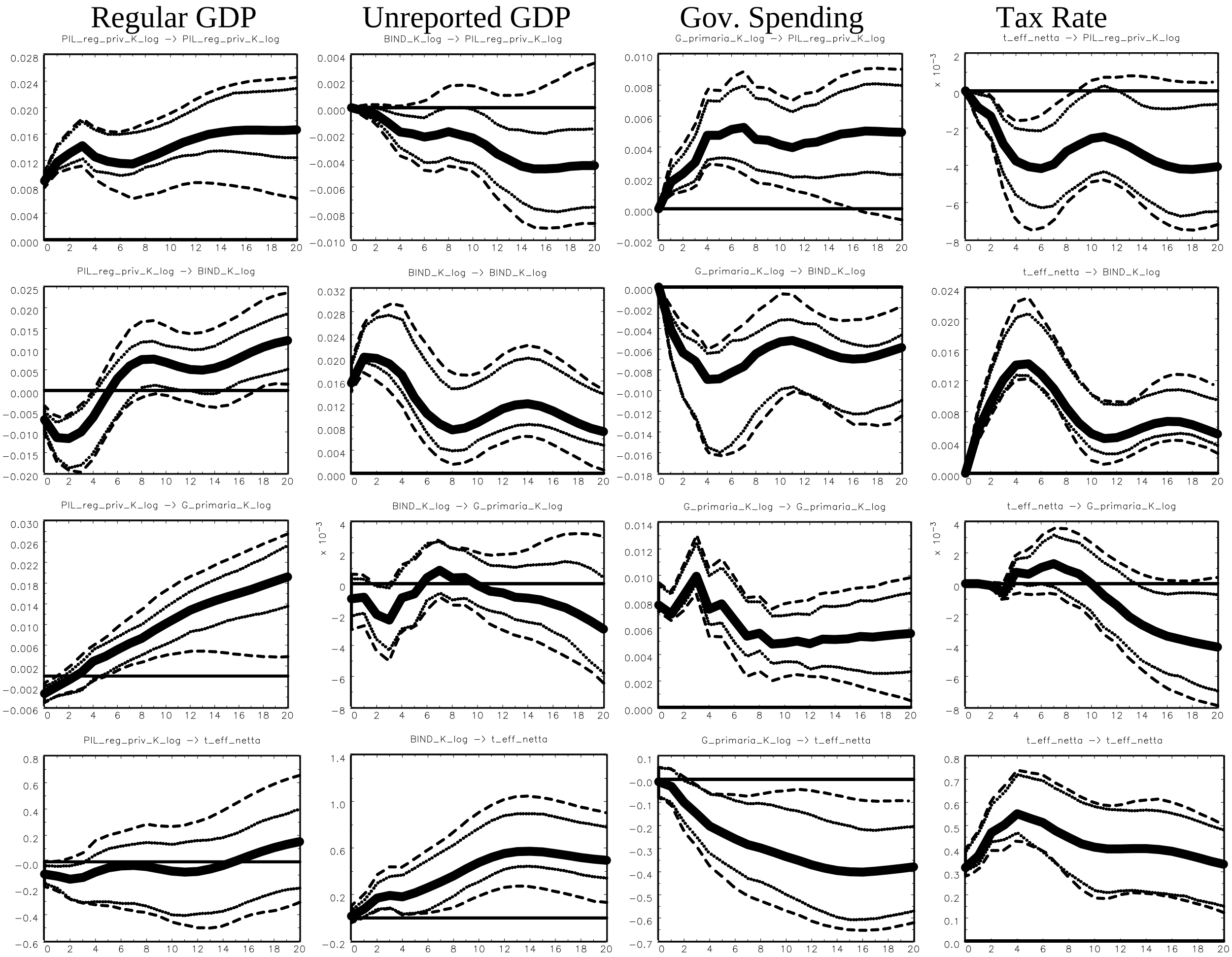
Test	Jarque Bera			
	u_1	u_2	u_3	u_4
Test statistic	0.71	1.88	0.36	1.38
p-value	0.70	0.39	0.83	0.50

Table C5: correlation of VEC residuals, Benchmark Model

	Regular GDP	**Unreported GDP**	**Public Spending**	**Net effective Tax rate**
Regular GDP	1.000000e+00	-4.463771e-01	-4.064290e-01	-2.834976e-01
Unreported GDP	-4.463771e-01	1.000000e+00	6.249293e-02	1.520343e-01
Public Spending	-4.064290e-01	6.249293e-02	1.000000e+00	1.196190e-01
Net effective Tax rate	-2.834976e-01	1.520343e-01	1.196190e-01	1.000000e+00

C3.2 The Impulse Response analysis

Column 1: effects of one S.D. shock to regular GDP on regular GDP, Unreported GDP, Gov. Spending, Tax rate.
Column 2: effects of one S.D. shock to Unreported GDP on regular GDP, Unreported GDP, Gov. Spending, Tax rate.
Column 3: effects of one S.D. shock to Gov. Spending on regular GDP, Unreported GDP, Gov. Spending, Tax rate.
Column 4: effects of one S.D. shock to Tax Rate on regular GDP, Unreported GDP, Gov. Spending, Tax rate.



Solid line: VECM orthogonal Impulse Response; Dashed Line: 95% Hall Percentile Confidence Interval; Dotted line: 85% Hall Percentile Confidence Interval

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