Monetary Policy and Inequality in Mexico

Villarreal, Francisco G.

UN Economic Commission for Latin America and the Caribbean, Graduate School of Government and Public Policy, Tec de Monterrey

30 June 2014

Monetary policy and inequality in Mexico

Francisco G. Villarreal *

Subregional Headquarters in Mexico
Economic Commission for Latin America and the Caribbean
United Nations

Graduate School of Government and Public Policy
Tecnológico de Monterrey

June 30, 2014

Abstract

Despite growing interest regarding the distributive impact of macroeconomic policies, the relationship between monetary policy and inequality has received relatively little attention in the literature. This is partly explained by the fact that the workhorse model used for monetary policy analysis summarises the demand-side of the economy by means of a representative agent, whose welfare is the normative criterion of optimal policy. However, alternative formulations using incomplete market models which feature heterogeneous agents, indicate that monetary policy does have an effect on the distribution of income, consumption and wealth, which potentially has implications for the design and conduct of optimal policy. The document empirically investigates the nature of the relationship between monetary policy and household income inequality in Mexico. The ultimate purpose is to uncover certain regularities which characterise the relationship, which can eventually serve as stylised facts for the design of theoretical models. The response of household's income inequality, and its components, to monetary policy shocks indicate that unanticipated increases in the nominal interest rate are correlated with a reduction of household income inequality in the short run, and that the effect dissipates over a two-year horizon. The results are robust to the particular measure of inequality used, as well as the procedure used to identify the policy shocks.

Keywords: Monetary Policy, Income Distribution, Small Open Economy
JEL Codes: C1, D3, E5

*Email: francisco.villarreal@cepal.org

Contents

1 Introduction **3**

2 The evolution of inequality in Mexico **3**
2.1 Data . . . 4
2.2 Household income dynamics . . . 5

3 Identification of monetary policy shocks **6**
3.1 Dynamic Stochastic General Equilibrium Model . . . 7
3.1.1 Data . . . 8
3.1.2 Estimation . . . 8
3.1.3 Results . . . 9

4 Impact of monetary policy shocks on inequality **10**
4.1 Results . . . 11
4.2 Robustness of results . . . 12
4.2.1 Methodology . . . 12
4.2.2 Estimation . . . 13
4.2.3 Results . . . 14

5 Concluding remarks **14**

A Tables and figures **18**
A.1 Tables . . . 18
A.2 Figures . . . 21

1 Introduction

Although the analysis of the distributional consequences of certain macroeconomic policies, such as fiscal policy, is an integral part of their analysis and formulation; the distributional consequences of monetary policy have received relatively little attention in the literature. This partly reflects the fact that current practice favours the use of representative–agent monetary policy models. Thus, while welfare considerations are the normative criterion for the evaluation of monetary policy, the existence of a single representative agent precludes the analysis of the distributional consequences of policy.

The assumption of a representative agent is equivalent to the assumption that markets are complete. However, empirical evidence suggests that agents have differentiated access to certain key markets, such as the one for financial services, which result in substantial heterogeneity in the capacities of households to insure against idiosyncratic shocks, such as spells of unemployment.

Brzoza-Brzezina et al. (2013) and Guvenen (2011) review recent attempts to introduce heterogeneous households into monetary policy models. In general the evidence suggests that not only is heterogeneity affected by the conduct of monetary policy, but what was optimal policy under the representative agent framework is no longer optimal under heterogeneity.

The purpose of this paper is to identify empirical regularities regarding the impact of monetary policy shocks on the inequality of households in Mexico, which could eventually serve as reference in the development of models for monetary policy analysis and formulation for small open economies. In an ideal setting the focus would be on the effect of policy on inequality across the across the household's budget constraint, beginning with the effect on hours worked and finalising with the effect on household's wealth. Of particular interest is the effect on consumption. However, not all the relevant data is available, such as data on wealth; and some of it is not available only sporadically, such as the data on consumption which is only available on a biennial basis. Thus the focus is placed on the impact of monetary policy on labour income.

In general terms, the paper follows the approach used by Coibion et al. (2012) to analyse the impact of monetary policy shocks on inequality in the United States. That is monetary policy shocks are identified using a structural model, and inequality is measured from household survey data. Then the effect of shocks on inequality is evaluated using a time–series model.

In contrast to the results found for the United States, which indicate that contractionary policy increases household inequality; in Mexico the evidence suggests that, under the inflation targeting regime adopted in 2003, unanticipated increases in the interest rate have actually reduced household labour income inequality. A possible explanation is offered by Lee (2014), who develops a theoretical model where the presence of financial frictions imply threshold effects.

The rest of the document is organised as follows: Section 2 analyses the evolution of inequality of household's labour income, after controlling for the effect of observable characteristics. Next, section 3 details the procedures used to identify monetary policy shocks, and discusses the robustness of the findings to alternative specifications. Section 4 evaluates the impact of monetary policy shocks on inequality, and discusses the regularities found. Finally, section 5 concludes with some thoughts on the possible role of financial frictions in reconciling the results found for Mexico and the United States.

2 The evolution of inequality in Mexico

Although inequality remains one its distinctive features, in contrast to the rising trend observed in the developed world (Krueger et al., 2010), inequality in Latin America has fallen significantly since the late 1990s (López-Calva and Lustig, 2010; Ferreira and Ravallion, 2009).

In the case of Mexico, income inequality has exhibited a declining trend that began in 1994 and continues to date. The findings of Esquivel et al. (2010) indicate that over half of the observed reduction of income inequality can be explained by the reduction of inequality of labour income. In particular, after decomposing the reduction of labour income inequality into changes in the observable characteristics of the workforce, and changes in the returns to these characteristics, Campos et al. (2014) concludes that although changes in characteristics increased inequality of income, the dynamics of their returns compensated their effect and explain the dynamics of inequality reduction.

In view of this, and considering the availability of data the paper focuses on the impact of monetary policy shocks on the distribution of household labour income in Mexico. Before describing the main features of the recent evolution of inequality, the next subsection briefly describes the data sources as well as the procedures used to obtain the inequality measures which will be used throughout the paper.

2.1 Data

Considering the time frame in which monetary policy shocks propagate through the economy, it is necessary to use data at sub-annual frequencies. With this in mind, inequality is measured using labour income drawn from the Mexican Labour Force Survey, which is available on a quarterly basis.

The data set contains observations from the first quarter of 1995 through the fourth quarter of 2012. The start date was chosen to coincide with the adoption of a flexible exchange regime by the monetary authorities in Mexico. In addition, it roughly corresponds to the beginning of the period of declining inequality.

The observations from the period 1995.I through 2004.IV are drawn from the National Urban Employment Survey (ENEU) (INEGI, 2001), while the data corresponding to the period 2005.I through 2012.IV, come from the National Survey of Labour and Employment (ENOE) (INEGI, 2007). Since the ENEU survey is only representative at the urban level, in order to splice the data from both surveys, observations from the ENOE survey are restricted to those corresponding to urban areas which were persistently surveyed over the period 1995-2012[1]: Mexico City, Guadalajara, Monterrey, Puebla, León, San Luis Potosí, Mérida, Chihuahua, Tampico, Veracruz, Acapulco, Aguascalientes, Morelia, Toluca, Saltillo, Villahermosa, Tijuana, Culiacán, Hermosillo, Durango, Tepic, Campeche, Cuernavaca, Oaxaca, Zacatecas, Colima, Querétaro and Tlaxcala.

At the individual level, labour income is computed for workers aged 25-65 who are regular residents of the household surveyed, that worked a positive number of hours during the week previous to the survey, and that do not report working on the street in exchange for tips. In order to reduce the bias introduced by extreme observations, the sample is further restricted to those individuals which report real hourly wages of less than 2,000 pesos.

Household labour income includes income by all household members who are older than 14 years of age, however for the computation of summary statistics only households whose head is aged between 25 and 65 are considered. Equivalised household income is calculated by adjusting each household member's labour income by a factor of 1 for the household head and 0.5 for individuals of at least 14 years of age.

Individual hourly wages are obtained by dividing the reported monthly income over the product of reported weekly hours worked times a factor of 4.33. All nominal income measures are deflated using the consumer price index with base corresponding to the second fortnight of December 2010.

[1]The splicing algorithm is partially based on the procedure used by Alcaraz and Nakashima (2013)

Unless otherwise noted, all summary measures are computed using survey sampling weights. In the case of hourly wages, the weights used are the product of sampling weights times weekly hours worked.

2.2 Household income dynamics

Despite the importance of labour force characteristics and their returns in the determination of the dynamics of household income and its inequality, as documented by Binelli and Attanasio (2010), even after controlling for them significant fluctuations remain unexplained.

Under the assumption that at least over the short run, changes in worker's characteristics and their returns are independent of monetary policy shocks, the focus of the study is on the evolution of inequality once the effect of observable characteristics have been taken into account. Thus unless noted otherwise, all subsequent references to household's labour income and their corresponding inequality measures, are based on residual measures.

In order to obtain residual measures of income and its dispersion, labour supply and income measures (weekly hours worked and hourly wages perceived by individuals, as well as the aggregate and equivalised household income) are regressed on a square polynomial on age, and a set of dichotomic variables which indicate whether the individual is a female, finished high–school, has access to social security services, the urban area where the household is located, and the industry where the individual works. In the case of households, the regressors correspond to those of the household head.

Figure 1 shows the evolution of selected regressor coefficients in the determination of the four income measures used. The first feature to note is that the coefficients of the four regressors are statistically significant, and have the expected signs.

Due to structural characteristics of the labour market, women work less hours and perceive a lower hourly wage, which result in lower household incomes of households headed by women, even after controlling for the size of the household. Although the bias that affects female hours worked strengthened during the early 2000's, the effect has been compensated by an improvement in hourly wages. This explains the slight reduction in the gender gap in household income over the period studied, which nonetheless has remained constant once the household's structure is taken into consideration.

With respect to age, which is generally used as a proxy for experience, it can be confirmed that it has a positive effect on all measures, and although not shown the coefficient on its squares indicates that the effect is marginally decreasing. As in the case of the female coefficient, albeit with the opposite sign, the effect of increased returns on experience in hours worked is cancelled by declining returns in wages, which result in a roughly constant effect on both measures of income.

It is interesting to note that the return of having finished high school on hours worked is relatively constant and negative. This is related to lower-skilled workers working for longer to compensate for lower wages. With respect to wages, consistent with the findings of Campos et al. (2014) the results imply that returns to education have been on a downward trend since the mid 1990's. The dynamic of schooling returns on income at the household level largely reflects the declining return in hourly wages.

For its part, the return of being a beneficiary of social security, which is a proxy for working in the formal sector, has exhibited a general U shape over the study period across labour supply and income measures, decreasing from the high level attained as a result of the 1995 balance of payments crisis, then remaining roughly constant through the early 2000's and rising again as the economic conditions deteriorated as a result of the onset of the international financial crisis in 2009. The results suggest a strong cyclicality of the premium attached to working in the formal sector.

Figure 2 shows the evolution of median residual labour supply and income. Despite significant fluctuations, a generally declining trend can be observed in weekly worked hours. For its part, after a steep decline following the 1995 crisis, the hourly wage has been on a generally positive trend, roughly duplicating its value in constant terms over the period studied. Despite the falling trend in hours worked, in general the dynamics of median household income have followed the fluctuations of wages.

In order to provide a robustness check for the estimation of inequality, two alternative measures are used: the Gini coefficient[2], and the standard deviation of logarithms. Figure 3 summarises the evolution of inequality of labour supply and income under both inequality measures[3].

In general, the generally declining trend of household income inequality, found in other studies, is confirmed. It can be seen, however, that the rate of inequality reduction has stagnated since the early 2000s, and show a slight uptick towards the end of the sample. It is interesting to note that although the inequality of hours worked started increasing since the early 2000s, the evolution of household income inequality seems to be driven entirely by the evolution of hourly wage inequality.

Despite the differences in relative magnitude across the components of income, the alternative inequality measures provide a consistent story regarding the dynamics of labour supply and income inequality.

In order to provide some insight into the dynamics of inequality within the distribution, figure 4 shows the evolution of inequality according to two additional measures: the ratio of the ninth to the fifth decile, and that of the fifth to the first decile[4].

With respect to hours worked, it is evidently clear that while the rise in inequality is observed across the distribution, the increase in inequality in the bottom half is significantly larger. On the other hand, with respect to wages and household income, it is the reductions of inequality observed in the top half of the distribution that are driving the aggregate reduction of inequality.

Having captured the main features of the evolution of inequality from the mid 1990s through 2012, the next section describes the identification of monetary policy shocks used for the analysis.

3 Identification of monetary policy shocks

Under an inflation targeting regime, such as the one used to conduct monetary policy in Mexico, the policy instrument is a short-term interest rate, which in the case of Mexico is the overnight interbank lending rate.

Under the assumption that economic agents are forward–looking, agents will form expectations regarding the evolution of the policy rate. This means that if a rate change is fully anticipated, the effect of the actual change on economic aggregates will be negligible. However if the actual change is different from expectations, depending on the sign and magnitude of the discrepancy, monetary policy can have significant effects on the economy at large.

Thus, for the analysis of monetary policy the interest lies not on the observed changes in the policy rate, but on its unanticipated fluctuations, which are commonly referred to as monetary policy shocks. In order to identify monetary policy shocks it is necessary to impose some economic structure. To do so a standard open economy dynamic stochastic general equilibrium (DSGE) model is used.

[2] All Gini coefficient estimations were carried out using DASP (Araar and Duclos, 2007)

[3] It should be noted that the Gini coefficients for income are significantly lower than what is usually reported for Mexico. The lower value is explained by the relatively narrow focus on a subsample of the population, and more importantly by the fact that as discussed previously the measures are computed on residual inequality.

[4] It is important to highlight that both figures share common axes, so that the evolution of labour supply and income can be compared across the two halves of the distribution.

3.1 Dynamic Stochastic General Equilibrium Model

The basic model specification is that proposed by Lubik and Schorfheide (2007), which in turn is a simplification of the small open economy extension of the standard New Keynesian DSGE posed by Galí and Monacelli (2005).

The solution to the representative household's optimisation problem yields the following Euler equation, which characterises the equilibrium in the goods–producing sector, and can be thought of as a forward–looking open economy version of the traditional IS–curve:

$$\begin{aligned} y_t &= E_t y_{t+1} - [\tau + \alpha(2-\alpha)(1-\tau)]\,(R_t - E_t\pi_{t+1}) - \rho_A \Delta A_t \\ &\quad -\alpha\,[\tau + \alpha(2-\alpha)(1-\tau)]\,E_t\Delta q_{t+1} + \alpha(2-\alpha)\frac{1-\tau}{\tau}E_t\Delta y^*_{t+1} \end{aligned} \tag{1}$$

where the structural parameters are: τ which stands for the household's intertemporal substitution elasticity, and $\alpha \in (0,1)$ which denotes the import share in consumption, and proxies for the level of trade openness. The endogenous variables are domestic aggregate output y_t, and the consumer price inflation rate π_t. For their part, world output y^*_t, the first difference of the terms of trade Δq_t and the the non–stationary technology process A_t are considered exogenous[5].

Optimal price setting of monopolist producers which face restrictions on the frequency with which they can adjust prices, as in Calvo (1983), leads to the open economy version of the Phillips curve:

$$\pi_t = \beta E_t\pi_{t+1} + \alpha\beta E_t\Delta q_{t+1} - \alpha\Delta q_t + \frac{\kappa}{\tau + \alpha(2-\alpha)(1-\tau)}(y_t - \overline{y}_t) \tag{2}$$

where $\overline{y}_t \equiv -\alpha(2-\alpha)(1-\tau)/\tau y^*_t$ denotes the level of output that would be observed in the absence of nominal frictions. In addition to τ and α, structural parameters include the discount factor β, and κ which determines the slope of the Phillips curve, and is a function of the labour supply and demand elasticities as well as the degree of price-stickiness.

Under the assumption that relative purchasing–power parity holds, the consumer price index can be defined as:

$$\pi_t = \Delta e_t + (1-\alpha)\Delta q_t + \pi^*_t \tag{3}$$

where Δe_t is the first difference of the nominal exchange rate.

The policy block of the model is summarised by a Taylor–type rule which governs the evolution of the nominal interest rate R_t, allowing for the possibility that the monetary authority responds to fluctuations in the exchange rate, in addition to changes in inflation and output:

$$R_t = \rho_R R_{t-1} + (1-\rho_R)(\psi_\pi\pi_t + \psi_t y_t + \psi_e\Delta e_t) + \varepsilon^R_t \tag{4}$$

where in principle it is assumed that the policy parameters ψ_j for $j \in \{\pi, y, e\}$ are nonnegative. The term ρ_R is a smoothing term included to match the persistence commonly observed in interest rates. The term ε^R_t is an exogenous policy shock whose identification is, for the purposes of this paper, the objective of the estimation of the model.

The model is closed by defining the laws of motion which determine the evolution of the rest of the exogenous variables, which as in Lubik and Schorfheide (2007) are assumed to follow AR(1)

[5] As discussed by Lubik and Schorfheide (2007), since firms have a degree of market power, the evolution of international prices is not entirely exogenous. This means that in strict terms, the terms of trade are determined endogenously as the relative price that clears the international goods market. However, according to Lubik and Schorfheide (2007) this imposes overly tight cross–equation restrictions which yield implausible values for the rest of the structural parameters. In view of the above, the evolution of the terms of trade are modeled as following the exogenously determined law of motion described below.

processes:

$$\begin{aligned}
\Delta A_t &= \rho_A \Delta A_{t-1} + \varepsilon_t^A \\
\pi_t^* &= \rho_{\pi^*} \pi_{t-1}^* + \varepsilon_t^{\pi^*} \\
y_t^* &= \rho_{y^*} y_{t-1}^* + \varepsilon_t^{y^*} \\
\Delta q_t &= \rho_q \Delta q_{t-1} + \varepsilon_t^q
\end{aligned}$$

where the ε_t's are stochastic innovations which drive the respective processes.

3.1.1 Data

The identification of monetary policy shocks is carried out using quarterly data from 2003.I to 2012.IV. The start date corresponds to the formal adoption of an inflation targeting regime by the Bank of Mexico (Banco de México, 2007).

Data on the evolution of Mexican gross domestic product and consumer, export and import prices come from INEGI. In the case of GDP, quarterly series expressed at 2003 constant peso prices are used. For consumer prices, the general monthly index, with base equal to the second fortnight of December 2010 is used. The terms of trade index is built from changes in the ratio of monthly export to import unit prices, which have base 1980. In the case of interest rates the nominal quarterly average overnight interbank rate compiled by Banco the Mexico is used. Data corresponding to quarterly GDP for the United States is drawn from the Bureau of Economic Analysis, with base 2009; and monthly US inflation data comes from the all–urban consumer price index series from the Bureau of Labor Statistics.

All monthly series are averaged to obtain quarterly observations. Percent changes in GDP, terms of trade and exchange rates were computed by multiplying quarter on quarter log changes times 100. In the case of consumer prices, the inflation rate is annualised by multiplying log changes times 400. All series were seasonally adjusted using TRAMO–SEATS (Gómez and Maravall, 1994, 2001), and detrended using the Hodrick-Prescott filter.

3.1.2 Estimation

In order to identify the structural monetary policy shocks implied by the model, it is necessary to estimate the vector of parameters $\theta = [\psi_\pi, \psi_y, \psi_e, \rho_R, \alpha, \beta, \kappa, \tau, \rho_q, \rho_a, \rho_{y^*}, \rho_{\pi^*}, \sigma_r, \sigma_q, \sigma_a, \sigma_{y^*}, \sigma_{\pi^*}]'$, from the observables vector $Y_t = [\pi_t, y_t, \Delta e_t, R_t, \Delta q_t]$. To do so, a distribution is defined for each of the parameters to be estimated. Based on this prior distribution, the data is used to update the prior by means of the Kalman filter. Following An and Schorfheide (2007), the posterior distribution is then estimated by generating draws from the posterior form obtained by applying Bayes' theorem to the likelihood function.

Table 1 summarises the priors used for estimation. The benchmark prior takes into consideration estimations of small open economy models found in the literature, however it should be noted that in general the parameter priors are fairly loose.

Based on the results found by Cermeño et al. (2012) for Mexico, the means of the policy–rule parameters on inflation (ψ_π) and output (ψ_y) are set to 1,5 and 0,75 respectively, while the selected mean of the interest rate persistence parameter ρ_R is 0.80. In the absence of relevant information for the Mexican case, the mean of the exchange rate policy parameter (ψ_e) is centered at 0.25, while the mean of the substitution elasticity parameter (τ) is 0.5. Both values correspond to the benchmark prior used by Lubik and Schorfheide (2007) for Canada. Following Best (2013) the mean of the import share (α) is set to 0.5. In the same fashion as Lubik and Schorfheide (2007) the intertemporal discount factor β is parametrised in terms of the steady–state interest rate (r_{ss}). Based on the estimates of Ramos-Francia and Torres (2008) its mean is chosen to be 2. Considering the same results, the mean for the Phillips curve slope coefficient (κ) is tightly centered at 0.02. Following Lubik and Schorfheide (2007) the parameters of exogenous processes are chosen by fitting AR(1) models on the corresponding variables[6]. The priors for the standard deviation of shocks are non-informative.

As a robustness check, two alternative priors are used. The first one is the benchmark prior used by Lubik and Schorfheide (2007) for Canada, and the second one uses the policy parameter priors proposed by Best (2013) for the case of Mexico where in contrast to the benchmark prior, the prior for the policy parameter is non–informative and does not constrain it to have a positive value.

3.1.3 Results

The parameter estimation results are listed in table 2. While a detailed evaluation of the model is beyond the scope of this document, a few comments are in order.

Regarding the policy–rule parameter, across the three alternative specifications, results confirm the Bank of Mexico's adherence to the so–called Taylor principle (Woodford, 2003), where monetary policy response prioritises the response to deviations of inflation from its target, and the magnitude of the response if more than proportional to the deviation observed.

Second, it is interesting to note that when the exchange rate policy parameter is not constrained to be strictly positive, it becomes statistically insignificant. The result is puzzling since it is at odds with the results found by Best (2013). Another result which merits attention is the relatively small value that is estimated for the import share α, given the importance of imported goods and services in the domestic consumption basket. However, a thorough investigation is outside the purview of this work, and is left for future research.

Third, considering the very small value assigned to the prior of the Phillips curve slope parameter κ, the estimated coefficient is significantly smaller than the one found using the Lubik and Schorfheide (2007) prior. As discussed by Lee (2014), a small slope parameter could be evidence of significant financial frictions in the economy. It is important to bear in mind that a flatter Phillips curve implies larger inflation stabilisation costs.

The top panel of figure 5 plots the estimated monetary policy shocks across time. For reference, the bottom panel shows the evolution of the overnight interbank lending rate, and the annualised inflation rate over the same period.

There are several features to note. The first is that despite the differences in the estimated coefficients summarised in table 2, the evolution of identified shocks across time is very similar under the alternative specifications, although it should be noted that the shocks identified under alternative 1 tend to be larger in magnitude.

Second, regarding the stance of monetary policy, it can be seen that from the middle of 2003 through the end of 2005, the Bank of Mexico embarked on a tightening cycle in response to the breaches of the upper bound of targeted inflation of 4%. This resulted in a series of positive monetary policy shocks, that is unexpected increases of the nominal interest rate. Once inflation returned to its targeted range, the stance of monetary policy was broadly neutral until the onset of the financial crisis in late 2008, where the reduction of the target for the policy rate resulted in strongly expansionary stance of monetary policy until the beginning of 2010. In the most recent period, monetary policy has remained broadly neutral.

[6]Given the tight relationship between the economies of Mexico and the United States, the latter's GDP and inflation are chosen as proxies for their world counterparts.

Finally, from the figure it becomes clear that changes in the nominal policy rate do not always correspond to the occurrence of monetary policy shocks. For example, note that as a result of the flight to quality resulting from concerns about the sovereign debt sustainability in the Euro zone, the depreciation of the Mexican peso caused an increase of inflation starting from the third quarter of 2011. Despite this increase in inflation, the central bank decided to keep its target rate on hold. This is interpreted by the model as the negative, that is expansionary, monetary policy shock evident in the top panel of the figure in late 2011.

The impulse responses of endogenous variables to a unit contractionary monetary policy shock under the three alternative specifications considered are shown in figure 6. The main difference between the three alternative estimations is explained by differences in the slope of the Phillips curve. As mentioned above a flatter Phillips curve, such as that found under the benchmark and alternative 2 specifications, implies that for a given inflation reduction the central bank must tolerate a larger deviation of output from its potential level. This means that for similarly sized shocks, the response of inflation is larger in magnitude and faster under the alternative 1 prior. Moreover the relative fall in contemporaneous fall in output is smaller and its eventual rebound is faster when the Phillips curve is steeper. Finally, reflecting the sluggishness of adjustment under the benchmark and alternative 2 priors, interest rates are more persistent implying a longer period for exchange rates to return to their steady state level.

Despite the differences in magnitude and speed, the system variables respond in a qualitatively similar fashion. Unanticipated increases in the nominal exchange rate reduce output growth, inflation and cause an appreciation of the nominal exchange rate contemporaneously. While inflation and the exchange rate return to their original levels monotonically over a horizon which is proportional to the response of the nominal interest rate, growth of GDP quickly rebounds to positive territory and then declines monotonically.

In general the responses are consistent with those found in the literature. The next section explores the impact of the alternative monetary policy shocks on the evolution of household labour income inequality.

4 Impact of monetary policy shocks on inequality

Using the evolution of residual inequality measures of household income discussed in section 2, and the monetary policy shocks identified in section 3, it is now possible to investigate the impact of monetary policy shocks on household inequality in Mexico.

To do so, the natural alternative is to first estimate a VAR, and then invert the coefficient matrix in order to compute the impulse response functions. However, unless the true data generating process is well characterised by a VAR, the estimated responses will be biased. Considering this, the impulse responses are instead estimated using the local projection method proposed by Jordà (2005), which is robust to misspecification and approaches the results obtained using a VAR when it is the true data generating process.

Letting, as in (6), Y_t denote the vector of variables of interest, in essence the idea behind obtaining impulse responses using local projections is to estimate the linear projection of the s–step ahead vector Y_{t+s} onto the linear space generated by the information available at time t:

$$Y_{t+s} = \alpha^s + B_1^{s+1} Y_{t-1} + B_2^{s+1} Y_{t-2} + \cdots + B_p^{s+1} Y_{t-p} + u_{t+s}^s \tag{5}$$

where the objects of interest are the coefficient matrices B_i^{s+1} for lag i and horizon $t+s$. Defining the impulse responses as the difference between two forecasts at the same horizon, Jordà (2005) defines the impulse response from the local linear projection (5) as $\widehat{IR}(t,s,d) = \hat{B}_1^s d$, where d is a column vector which defines the shock structure to be investigated.

At its simplest level household income will be affected by monetary policy shocks trough their effects on the aggregate level of production, and on inflation. Thus in order to control for a very general transmission channel, the impulse responses are computed including output growth and inflation in vector T_t in equation 5.

4.1 Results

Before discussing the relationship between monetary policy shocks and inequality it is illustrative to have a look at the relation between policy shocks and the level of the variables. The cumulative responses, and the corresponding single standard deviation confidence interval, of median hours worked, hourly wages and household income to the different monetary policy shocks are summarised in figure 7[7].

Despite the differences in magnitude of the identified monetary policy shocks shown in the top panel of figure 5, as well as of the responses of the rest of the variables in the system summarised in figure 6, the response of median labour supply and labour income measures is very similar across shocks.

The results summarised in the first column of 7 indicate that as a response of unanticipated increases in the nominal interest rate, hours worked increase significantly over the first four quarters following the shock, perhaps as a response aimed at compensating the observed reduction in hourly wages, which are shown in the second column of the figure. From the third column it appears, however, that the observed increase in hours worked is not sufficient to keep household income from falling as a response to contractionary monetary policy shocks. Despite this, as shown in the fourth column, once the composition of households is accounted for the effect of shocks on equivalised household income is only marginally significant. This is probably evidence of household members pooling resources within households.

In a similar fashion to the response of median income, the response of inequality to the alternative shocks is both qualitatively and quantitatively similar. Thus, in the remainder of this section only the results to the monetary policy shocks identified under the benchmark prior specification will be discussed.

The response of inequality to unanticipated increases in the nominal interest rate are shown in figure 8. The top two rows summarise the responses of respectively, the Gini coefficient and the standard deviation of logarithms to monetary policy shocks.

The first feature to note is that the responses are robust to the particular measure of inequality used. Moreover, the results indicate that inequality tends to move in the opposite direction of the levels of the variables. That is, despite the slight increase in the dispersion of hours worked, the observed reduction in inequality of household income appears to be driven by the reduction of inequality of hourly wages across households. This result is significant even after for controlling for household composition.

It is also worth noting, that the effect is statistically significant for roughly six quarters, however it appears that the response becomes indistinguishable from zero not as a result of sampling uncertainty, but as a result of a reversal of the dynamics of inequality towards the end of the two-year horizon.

The bottom two rows of figure 8 show the cumulative responses of the 9th. to 5th., and 5th. to 1st. decile ratios. A quick glance at the pairs of responses reveals that the impact of monetary

[7] For reference the dash and dot line depict the original, non–cumulative, response to a unit monetary policy shock.

policy shocks is differentiated across the distribution of households. While the effect of hours worked across both halves of the distribution is broadly similar, it is evident that the reduction of inequality observed in hourly wages, and consequently on household income, occurs mainly in the top half of the distribution.

In summary, the evidence discussed indicates that an unanticipated increase of the nominal interest rate reduces median household income through a reduction of hourly wages that is not compensated by the observed increase in hours worked. This reduction of income is accompanied by a compression of the dispersion of both wages and income across households. The reduction in inequality seem to be more acute in the top half of the distribution.

Before comparing the results to those found by Coibion et al. (2012) for the United States, the next section evaluates the robustness of the results to the strategy used to identify the policy shocks, as well as the methodology used to estimate their effect on household income inequality.

4.2 Robustness of results

The heavy parametrisation of the NK–DSGE used in the previous section imposes a number of cross–equation restrictions which may not necessarily be supported by the data. With this in mind, they are also estimated by imposing restrictions on the impulse–response functions of a vector autoregressive (VAR) model.

As discussed by Fry and Pagan (2011), from the several types of restrictions that can be imposed on the impulse responses, in principle long–run (Blanchard and Quah, 1989) and sign restrictions (Canova and Nicoló, 2002; Faust, 1998; Uhlig, 2005) are the least restrictive[8]. Considering this, in order to allow the data to "speak" as freely as possible, monetary policy shocks are identified under alternative long–run and sign restrictions using a VAR containing the endogenous variables of the model used in the previous section. In particular a four–variable VAR including GDP growth (Δy), inflation(π), as well as interest rate (R) and exchange rate variations (Δe) is used; from which the following types of shocks can be identified (Peersman and Smets, 2003): aggregate supply (as) shocks, aggregate demand (ad) shocks, monetary policy (mp) shocks and shocks to the exchange rate (fx).

4.2.1 Methodology

To gain some insight into the procedure used to identify monetary policy shocks by imposing restrictions on a VAR, let a reduced–form VAR model of order p be defined as follows:

$$\mathbf{Y}_{t+1} = B(L)\mathbf{Y}_t + u_{t+1} \tag{6}$$

where $\mathbf{Y}_t' = [\Delta y, \pi, R, \Delta e]$ is a vector of variables observed at time t, $B(L) \equiv B_1 L + B_2 L^2 + \cdots + B_p L^p$ is a lag polynomial of order p, and the covariance matrix of the innovations u_t is given by $E u_t u_t' = \Sigma$.

The identification problem can be thought of as the search for a matrix Z which allows the identification of the structural shocks ε_t such that $u_t = Z\epsilon_t$, $E\varepsilon_t\varepsilon_t' = \mathbf{I}$, and $ZZ' = \Sigma$. Typically there exists a multiplicity of Z matrices that represent de data, that is matrices that satisfy $ZZ' = \Sigma$, thus it is necessary to impose restrictions to identify a particular Z.

Using the notation of Rubio-Ramírez et al. (2010), the matrix that summarises the long–run impact of shocks on the system variables can be written as $A_+ = (\mathbf{I} - \mathbf{B})^{-1} A_0$, where $\mathbf{I}$ is the identity matrix, and $A_0 = Z$ is the contemporaneous impact matrix. Restrictions can be imposed on the

[8]Other restrictions include recursive identification as in Sims (1980), and restriction on the contemporaneous effect of shocks on system variables as in Galí (1992).

impulse responses of system (6) by estimating its coefficients subject to constrains of particular elements of the response matrices A_0 and A_+.

Under the assumption that in the long–run, output fluctuations are driven exclusively by aggregate supply shocks, as in Blanchard and Quah (1989), the long–run restriction used to identify monetary policy shocks can be expressed as follows:

$$A_+ = \begin{array}{c} \\ \Delta y \\ \pi \\ R \\ \Delta e \end{array} \begin{array}{c} \begin{array}{cccc} as & ad & mp & fx \end{array} \\ \begin{bmatrix} \times & 0 & 0 & 0 \\ \times & \times & \times & \times \\ \times & \times & \times & \times \\ \times & \times & \times & \times \end{bmatrix} \end{array} \tag{7}$$

where $\times$'s imply no restriction.

For its part, the following two sets of sign restrictions are used:

$$A_0^{(mp)} = \begin{array}{c} \\ \Delta y \\ \pi \\ R \\ \Delta e \end{array} \begin{array}{c} \begin{array}{cccc} as & ad & mp & fx \end{array} \\ \begin{bmatrix} \times & \times & - & \times \\ \times & \times & - & \times \\ \times & \times & + & \times \\ \times & \times & - & \times \end{bmatrix} \end{array}, \qquad A_0^{(ex-mp)} = \begin{array}{c} \\ \Delta y \\ \pi \\ R \\ \Delta e \end{array} \begin{array}{c} \begin{array}{cccc} as & ad & mp & fx \end{array} \\ \begin{bmatrix} + & + & \times & - \\ - & + & \times & - \\ - & + & \times & - \\ - & - & \times & + \end{bmatrix} \end{array} \tag{8}$$

In the first, it is assumed that the sign of the contemporary impact of a monetary policy shock, increases the interest rate and induces a reduction in output and unemployment, as well as an appreciation of the nominal exchange rate. In the second, the signs of the responses of the variables to aggregate supply and demand, and foreign exchange shocks are specified, while no constraints are imposed on the signs of the responses to monetary policy shocks. The signs of the restrictions are drawn from those observed in the NK–DSGE model in the previous section.

The following three alternative restrictions are imposed: monetary-policy shock sign restrictions, long–run and monetary policy shock sign restrictions, and long–run and shocks different from monetary policy sign restrictions. Since the restrictions imposed are not sufficient to uniquely identify the shocks, the simulation approach described in the following section is adopted.

4.2.2 Estimation

The estimation is carried out using the generalisation of the Rubio-Ramírez et al. (2010) algorithm due to Binning (2013). The starting point for the algorithm is the estimation of the innovation covariance matrix $\hat{\Sigma}$ from a reduced–form VAR as in (6), where the lag length of the VAR is selected according to the Bayesian information criterion. The Choleski factorisation of matrix $\hat{\Sigma}$ is then multiplied by an orthonormal random matrix in order to randomise the impact matrix, and thus initialise the simulation procedure. The corresponding orthonormal matrix is obtained by carrying out a QR-decomposition, using Householder transformations, on a random matrix drawn from a multivariate standard normal.

Next, the algorithm searches for a 'rotation' matrix that satisfies the long–run restrictions (See Rubio-Ramírez et al. (2010) for details). Once such a matrix is found, the impulse responses to the shocks are computed and the sign restrictions are verified. A draw is kept if all the restrictions are met, and discarded otherwise. The algorithm proceeds iteratively until 1,000 successful draws are obtained.

Selection of a particular draw to recover the evolution of the structural shocks, is carried out using the median target criterion proposed by Fry and Pagan (2005), which basically solves a least squares minimisation problem to find the draw which is closest to the median distribution across all the impulse responses in the system.

The effect of the identified shocks is computed by including the measures of inequality of hours worked, hourly wages and equivalised income in the system used to identify the shocks, without imposing any additional sign restrictions on them.

4.2.3 Results

The results of the estimation are summarised in figures 9 through 11. The top row in each figure shows the response functions of GDP growth, inflation, interest rate and the exchange rate to the identified shocks; while the bottom row shows the responses of inequality in hours worked, hourly wages and equivalised household income. The corresponding confidence intervals were bootstrapped using the OLS residuals of the VAR estimated under the selected factorisation.

Reflecting the relatively short lag–lengths selected by the Bayesian information criterion, a single lag in most cases, the responses exhibit some kinks in the very short run. Moreover, reflecting the relative looseness of the procedure used the confidence intervals about the responses are relatively broad. As a consequence, in comparison to the results found using the DSGE model above the responses are shorter lived.

Nevertheless, under the three alternative restrictions, the signs of the responses of GDP, inflation, interest and exchange rates are consistent with the previous results. That is positive monetary policy shocks contemporaneously decrease growth and inflation, increase the nominal interest rate and cause an appreciation of the exchange rate. As before, GDP growth quickly rebounds and exhibits a hump-shaped profile.

Although the response of the inequality of hours worked is only positive under the shocks identified under the combination of long–run and sign restrictions corresponding to shocks different from monetary policy; the responses of the inequality of hours worked, and equivalised household income, are negative and statistically significant on impact.

Considering the relatively short time-series available, the results are interpreted as evidence that at least over the short run, irrespective of the strategy used to identify shocks and the methodology used estimate their impact, unanticipated increases in nominal interest rates reduce the inequality of household labour income. The evidence suggests the effect operated through a compression of the hourly wages.

The next section concludes by comparing the results with those obtained by Coibion et al. (2012) for the case of the United States.

5 Concluding remarks

In contrast to the results found for the case of Mexico, where at least under the inflation targeting regime in place since 2003, positive monetary policy shocks reduce the inequality of equivalised household income, Coibion et al. (2012) find the opposite is true for the case of the United States. A possible way to explain the differences in results is the existence of threshold effects.

In a recent paper, Lee (2014) introduces heterogeneity in an otherwise standard sticky price general equilibrium monetary policy model by introducing financial frictions which preclude full risk insurance. As mentioned above, the presence of financial frictions imply a flatter Phillips curve, which increases the cost of inflation stabilisation. In addition, in the model household income heterogeneity is a function of current and past inflation. This means that financial frictions increase the weight placed on inflation, relative to that placed on output, in the central bank's loss function. Lee (2014) shows that there exists a financial friction threshold beyond which the benefits of stabilising inflation exceed its costs.

Considering the different level of financial development between the United States and Mexico, and more importantly the vastly different levels of financial access and use in both countries, it is conceivable that the level of financial frictions in Mexico is such that inflation stabilisation is welfare enhancing, whereas the opposite occurs in the United States.

References

Adjemian, S., Bastani, H., Karamé, F., Juillard, M., Maih, J., Mihoubi, F., Perendia, G., Pfeifer, J., Ratto, M., and Villemot, S. (2011). Dynare: Reference manual, version 4. Dynare Working Papers 1, CEPREMAP.

Alcaraz, C. and Nakashima, C. (2013). Inlfación y pobreza en méxico (1993-2009). In Mendoza, A., editor, *Aplicaciones en Economía y Ciencias Sociales con Stata*. Stata Press, College Station, TX.

An, S. and Schorfheide, F. (2007). Bayesian analysis of DSGE models. *Econometric Reviews*, 26(2-4):113–172.

Araar, A. and Duclos, J.-Y. (2007). DASP: Distributive analysis data package. Technical report, PEP, World Bank, UNDP and Université Laval.

Banco de México (2007). *Informe sobre Inflación Julio–Septiembre 2007*. Banco de México, México DF.

Best, G. (2013). Fear of floating or monetary policy as usual? a structural analysis of mexico's monetary policy. *The North American Journal of Economics and Finance*, 24:45–62.

Binelli, C. and Attanasio, O. (2010). Mexico in the 1990s: The main cross-sectional facts. *Review of Economic Dynamics*, 13(1):238–264.

Binning, A. (2013). Underidentified SVAR models: A framework for combining short and long-run restrictions with sign-restrictions. Working Paper 14, Norges Bank.

Blanchard, O. J. and Quah, D. (1989). The dynamic effects of aggregate demand and supply disturbances. *The American Economic Review*, 79(4):655–673.

Brzoza-Brzezina, M., Kolasa, M., Koloch, G., Makarski, K., and Rubaszek, M. (2013). Monetary policy in a non-representative agent economy: A survey. *Journal of Economic Surveys*, 27(4):641–669.

Calvo, G. A. (1983). Staggered prices in a utility-maximizing framework. *Journal of Monetary Economics*, 12(3):383–398.

Campos, R., Esquivel, G., and Lustig, N. (2014). The rise and fall of income inequality in mexico, 1989-2010. In Cornia, G. A., editor, *Falling Inequality in Latin America. Policy Changes and Lessons*, UNU-WIDER Studies in Development Economics, pages 140–163. Oxford University Press, Oxford, UK.

Canova, F. and Nicoló, G. D. (2002). Monetary disturbances matter for business fluctuations in the g-7. *Journal of Monetary Economics*, 49(6):1131–1159.

Cermeño, R., Villagómez, F. A., and Polo, J. O. (2012). Monetary policy rules in a small open economy: an application to mexico. *Journal of Applied Economics*, 15(2):259–286.

Coibion, O., Gorodnichenko, Y., Kueng, L., and Silvia, J. (2012). Innocent bystanders? monetary policy and inequality in the U.S. Working Paper 18170, National Bureau of Economic Research.

Esquivel, G., Lustig, N., and Scott, J. (2010). A Decade of Falling Inequality in Mexico: Market Forces or State Action? In López-Calva, L. F. and Lustig, N., editors, *Declining Inequality in Latin America. A Decade of Progress?* Brookings Institution Press; United Nations Development Programme.

Faust, J. (1998). The robustness of identified VAR conclusions about money. *Carnegie-Rochester Conference Series on Public Policy*, 49:207–244.

Ferreira, F. H. G. and Ravallion, M. (2009). Poverty and inequality: The global context. In Salverda, W., Nolan, B., and Smeeding, T. M., editors, *The Oxford Handbook of Economic Inequality*, pages 599–636. Oxford University Press, Oxford, UK.

Fry, R. and Pagan, A. (2005). Some issues in using vars for macroeconometric research. CAMA Working Paper 2005-19, Centre for Applied Macroeconomic Analysis, Crawford School of Public Policy, The Australian National University.

Fry, R. and Pagan, A. (2011). Sign restrictions in structural vector autoregressions: A critical review. *Journal of Economic Literature*, 49(4):938–960.

Galí, J. (1992). How well does the IS-LM model fit postwar u. s. data? *The Quarterly Journal of Economics*, 107(2):709–738.

Galí, J. and Monacelli, T. (2005). Monetary policy and exchange rate volatility in a small open economy. *The Review of Economic Studies*, 72(3):707–734.

Gómez, V. and Maravall, A. (1994). Estimation, prediction, and interpolation for nonstationary series with the kalman filter. *Journal of the American Statistical Association*, 89(426):611–624.

Gómez, V. and Maravall, A. (2001). Seasonal adjustment and signal extraction in economic time series. In Peña, D., Tiao, G., and R.S., T., editors, *A Course in Time Series Analysis*, chapter 8, pages 202–247. J. Wiley and Sons, New York.

Guvenen, F. (2011). Macroeconomics with heterogeneity: A practical guide. *The Federal Reserve Bank of Richmond Economic Quarterly*, 97(3):255–326.

INEGI (2001). *Documento metodológico de la Encuesta Nacional de Empleo Urbano*. Institito Nacional de Estadística y Geografía, Aguascalientes, México.

INEGI (2007). *Cómo se hace la ENOE. Mt́odos y procedimientos*. Institito Nacional de Estadística y Geografía, Aguascalientes, México.

Jordà, O. (2005). Estimation and inference of impulse responses by local projections. *The American Economic Review*, 95(1):161–182.

Krueger, D., Perri, F., Pistaferri, L., and Violante, G. L. (2010). Cross-sectional facts for macroeconomists. *Review of Economic Dynamics*, 13(1):1–14.

Lee, J. W. (2014). Monetary policy with heterogeneous households and imperfect risk-sharing. *Review of Economic Dynamics*, 17(3):505–522.

López-Calva, L. F. and Lustig, N., editors (2010). *Declining Inequality in Latin America. A Decade of Progress?* Brookings Institution Press; United Nations Development Programme, Washington D.C.

Lubik, T. A. and Schorfheide, F. (2007). Do central banks respond to exchange rate movements? a structural investigation. *Journal of Monetary Economics*, 54(4):1069–1087.

Peersman, G. and Smets, F. (2003). The monetary transmission mechanism in the euro area: More evidence from a VAR analysis. In *Monetary Policy Transmission in the Euro Area. A Study by the Eurosystem Monetary Transmission Networ*, pages 36–55. Cambridge University Press, Cambridge, UK.

Ramos-Francia, M. and Torres, A. (2008). Inflation dynamics in mexico: A characterization using the new phillips curve. *The North American Journal of Economics and Finance*, 19(3):274–289.

Rubio-Ramírez, J. F., Waggoner, D. F., and Zha, T. (2010). Structural vector autoregressions: Theory of identification and algorithms for inference. *The Review of Economic Studies*, 77(2):665–696.

Sims, C. A. (1980). Macroeconomics and reality. *Econometrica*, 48(1):1.

Uhlig, H. (2005). What are the effects of monetary policy on output? results from an agnostic identification procedure. *Journal of Monetary Economics*, 52(2):381–419.

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

A Tables and figures

A.1 Tables

Table 1 – Prior distributions

	Benchmark prior				Alternative prior 1				Alternative prior 2			
Parameter	Density	Domain	μ	σ	Density	Domain	μ	σ	Density	Domain	μ	σ
ψ_π	Gamma	$\mathbb{R}^+$	1.50	0.750	Gamma	$\mathbb{R}^+$	1.50	0.50	Normal	$\mathbb{R}^+$	1.50	0.250
ψ_y	Gamma	$\mathbb{R}^+$	0.75	0.330	Gamma	$\mathbb{R}^+$	0.25	0.13	Normal	$\mathbb{R}^+$	0.50	0.250
ψ_e	Gamma	$\mathbb{R}^+$	0.25	0.130	Gamma	$\mathbb{R}^+$	0.25	0.13	Uniform	[-2,2]	0.00	1.150
ρ_R	Beta	[0,1)	0.80	0.100	Beta	[0,1)	0.50	0.20	Beta	[0,1)	0.80	0.100
α	Beta	[0,1)	0.50	0.200	Beta	[0,1)	0.20	0.05	Beta	[0,1)	0.50	0.200
r_{ss}	Gamma	$\mathbb{R}^+$	2.00	1.000	Gamma	$\mathbb{R}^+$	2.50	1.00	Gamma	$\mathbb{R}^+$	2.00	1.000
κ	Gamma	$\mathbb{R}^+$	0.02	0.006	Gamma	$\mathbb{R}^+$	0.50	0.25	Gamma	$\mathbb{R}^+$	0.02	0.006
τ	Beta	[0,1)	0.50	0.250	Beta	[0,1)	0.50	0.25	Beta	[0,1)	0.50	0.250
ρ_q	Beta	[0,1)	0.15	0.080	Beta	[0,1)	0.40	20.00	Beta	[0,1)	0.15	0.080
ρ_A	Beta	[0,1)	0.35	0.170	Beta	[0,1)	0.20	0.05	Beta	[0,1)	0.35	0.170
ρ_{y^*}	Beta	[0,1)	0.80	0.100	Beta	[0,1)	0.90	0.05	Beta	[0,1)	0.80	0.100
ρ_{π^*}	Beta	[0,1)	0.50	0.250	Beta	[0,1)	0.80	0.10	Beta	[0,1)	0.50	0.250
σ_R	Inv. Gamma	$\mathbb{R}^+$	1.00	4.000	Inv. Gamma	$\mathbb{R}^+$	0.50	4.00	Inv. Gamma	$\mathbb{R}^+$	1.00	4.000
σ_q	Inv. Gamma	$\mathbb{R}^+$	1.00	4.000	Inv. Gamma	$\mathbb{R}^+$	1.50	4.00	Inv. Gamma	$\mathbb{R}^+$	1.00	4.000
σ_A	Inv. Gamma	$\mathbb{R}^+$	1.00	4.000	Inv. Gamma	$\mathbb{R}^+$	1.00	4.00	Inv. Gamma	$\mathbb{R}^+$	1.00	4.000
σ_{y^*}	Inv. Gamma	$\mathbb{R}^+$	1.00	4.000	Inv. Gamma	$\mathbb{R}^+$	1.50	4.00	Inv. Gamma	$\mathbb{R}^+$	1.00	4.000
σ_{π^*}	Inv. Gamma	$\mathbb{R}^+$	1.00	4.000	Inv. Gamma	$\mathbb{R}^+$	0.55	4.00	Inv. Gamma	$\mathbb{R}^+$	1.00	4.000

Note: μ and σ respectively denote the means and standard deviations of the beta, gamma, normal and uniform distributions; and the scale and shape parameters of the inverse gamma distribution.

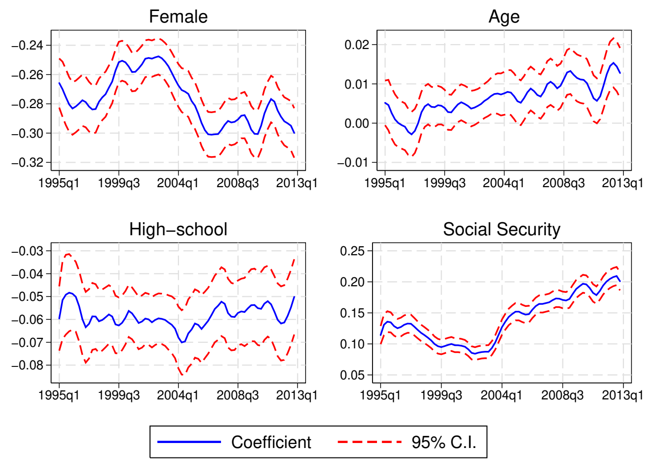
Table 2 – Parameter estimation results

	Benchmark		Alternative 1		Alternative 2	
Parameter	Mean	90% Interval	Alternative 1	90% Interval	Alternative 2	90% Interval
ψ_π	2.1806	[0.9560, 3.2773]	2.1805	[1.3233, 3.0396]	1.5755	[1.1868, 1.9644]
ψ_y	0.6459	[0.2109, 1.0842]	0.1904	[0.0366, 0.3344]	0.4648	[0.0641, 0.7947]
ψ_e	0.0955	[0.0213, 0.1671]	0.0805	[0.0321, 0.1288]	-0.0326	[-0.1273, 0.0611]
ρ_R	0.9018	[0.8554, 0.9507]	0.7974	[0.7189, 0.8826]	0.8693	[0.8201, 0.9215]
α	0.0104	[0.0024, 0.0177]	0.0594	[0.0371, 0.0814]	0.0085	[0.0016, 0.0152]
r_{ss}	1.9752	[0.5139, 3.3931]	2.5225	[0.9435, 4.1154]	1.9946	[0.5151, 3.4675]
κ	0.0237	[0.0137, 0.0338]	0.4502	[0.1662, 0.7270]	0.0234	[0.0130, 0.0331]
τ	0.0712	[0.0299, 0.1108]	0.3934	[0.2423, 0.5428]	0.0739	[0.0289, 0.1223]
ρ_q	0.1672	[0.0445, 0.2895]	0.3915	[0.1677, 0.6238]	0.1599	[0.0400, 0.2742]
ρ_A	0.2117	[0.1565, 0.2691]	0.1683	[0.1161, 0.2199]	0.2110	[0.1577, 0.2675]
ρ_{y^*}	0.7712	[0.6157, 0.9379]	0.8610	[0.7904, 0.9331]	0.7542	[0.5875, 0.9275]
ρ_{π^*}	0.2397	[0.0235, 0.4291]	0.5942	[0.4248, 0.7711]	0.2049	[0.0178, 0.3696]
σ_R	0.2298	[0.1819, 0.2744]	0.2442	[0.1811, 0.3056]	0.2174	[0.1745, 0.2596]
σ_q	4.2577	[3.4814, 5.0312]	4.3430	[3.5348, 5.1320]	4.2601	[3.4851, 4.9990]
σ_A	1.3333	[1.0621, 1.6170]	1.2572	[1.0191, 1.4896]	1.3298	[1.0597, 1.5877]
σ_{y^*}	1.0560	[0.2981, 1.8611]	1.7836	[0.5269, 3.1221]	1.4850	[0.2859, 3.6502]
σ_{π^*}	2.6571	[2.1480, 3.1324]	2.9452	[2.3529, 3.5131]	2.6322	[2.1531, 3.0901]

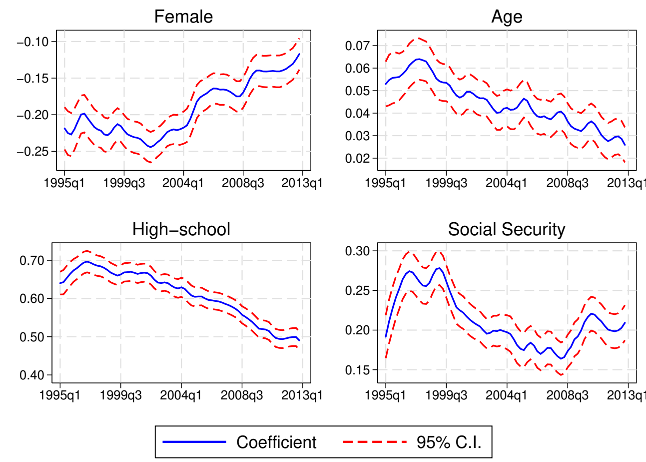
Note: The posterior distribution was estimated using Dynare version 4.4.0 (Adjemian et al., 2011) through 100,000 draws obtained using the Metropolis–Hastings algorithm, dropping 20% of the resulting draws.

A.2 Figures

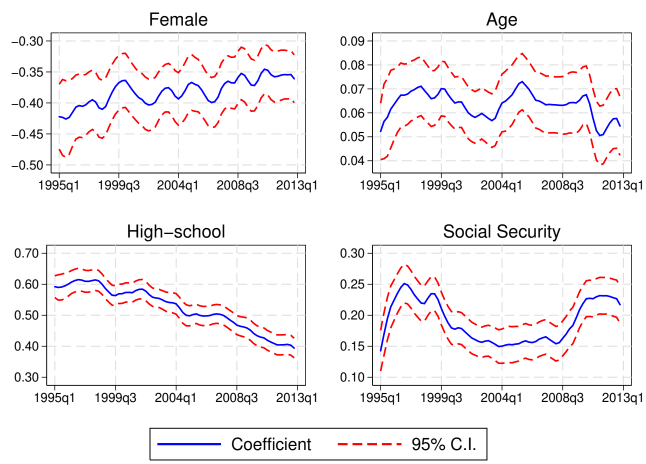
Figure 1 – Regression coefficients



(a) Weekly hours worked



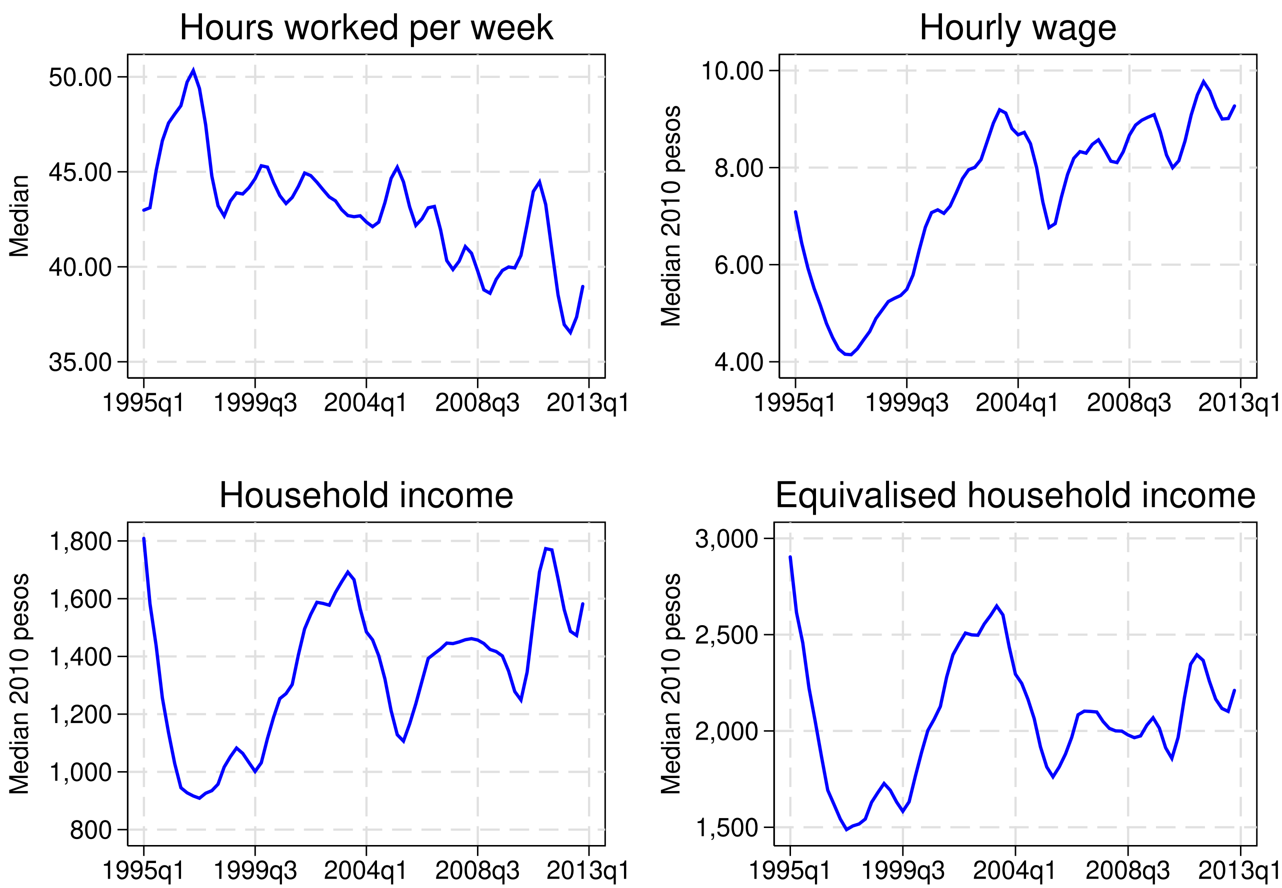
(b) Hourly wages



(c) Household income

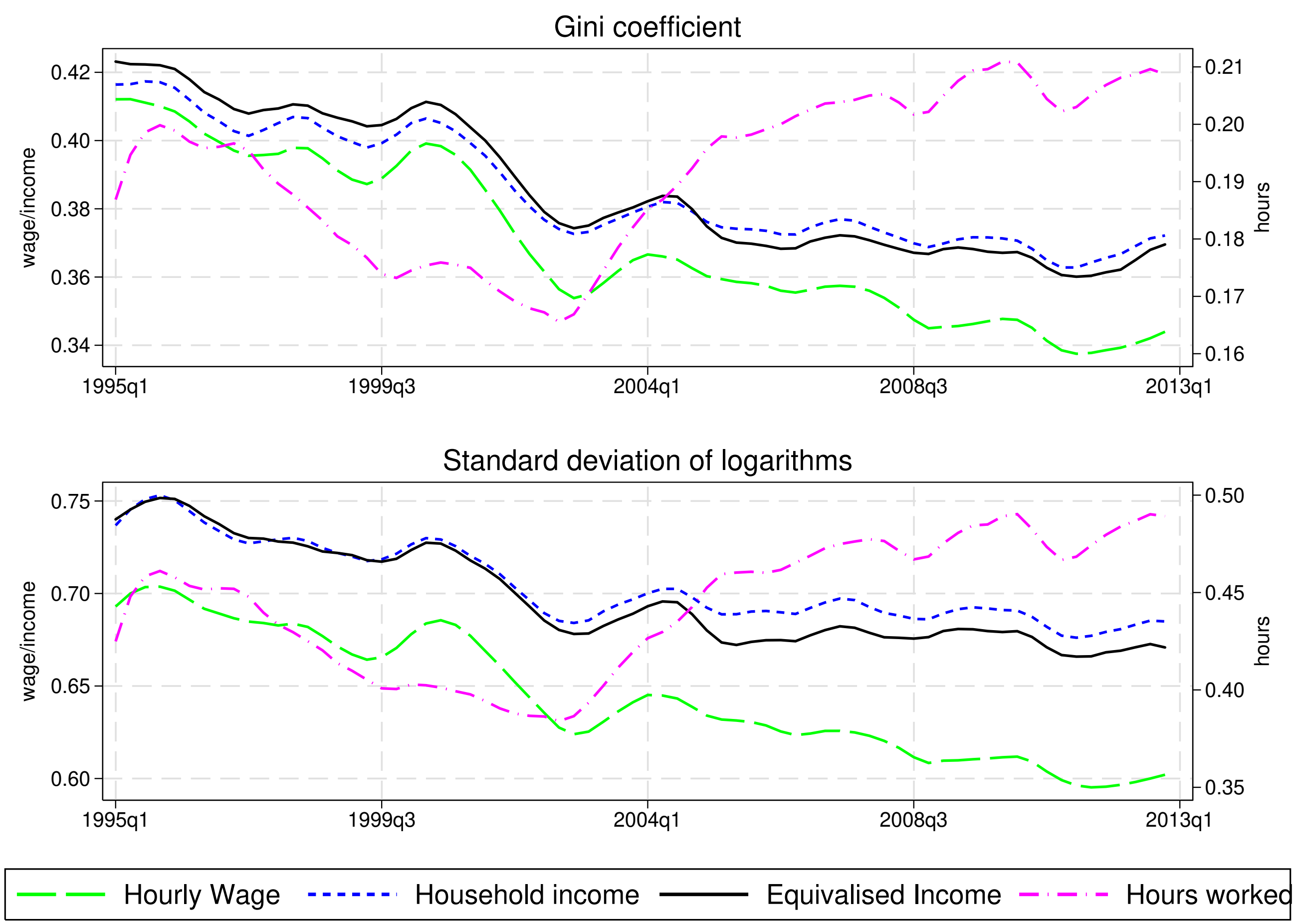
(d) Equivalised household income

Figure 2 – Mexico 1990 – 2012: Evolution of household income



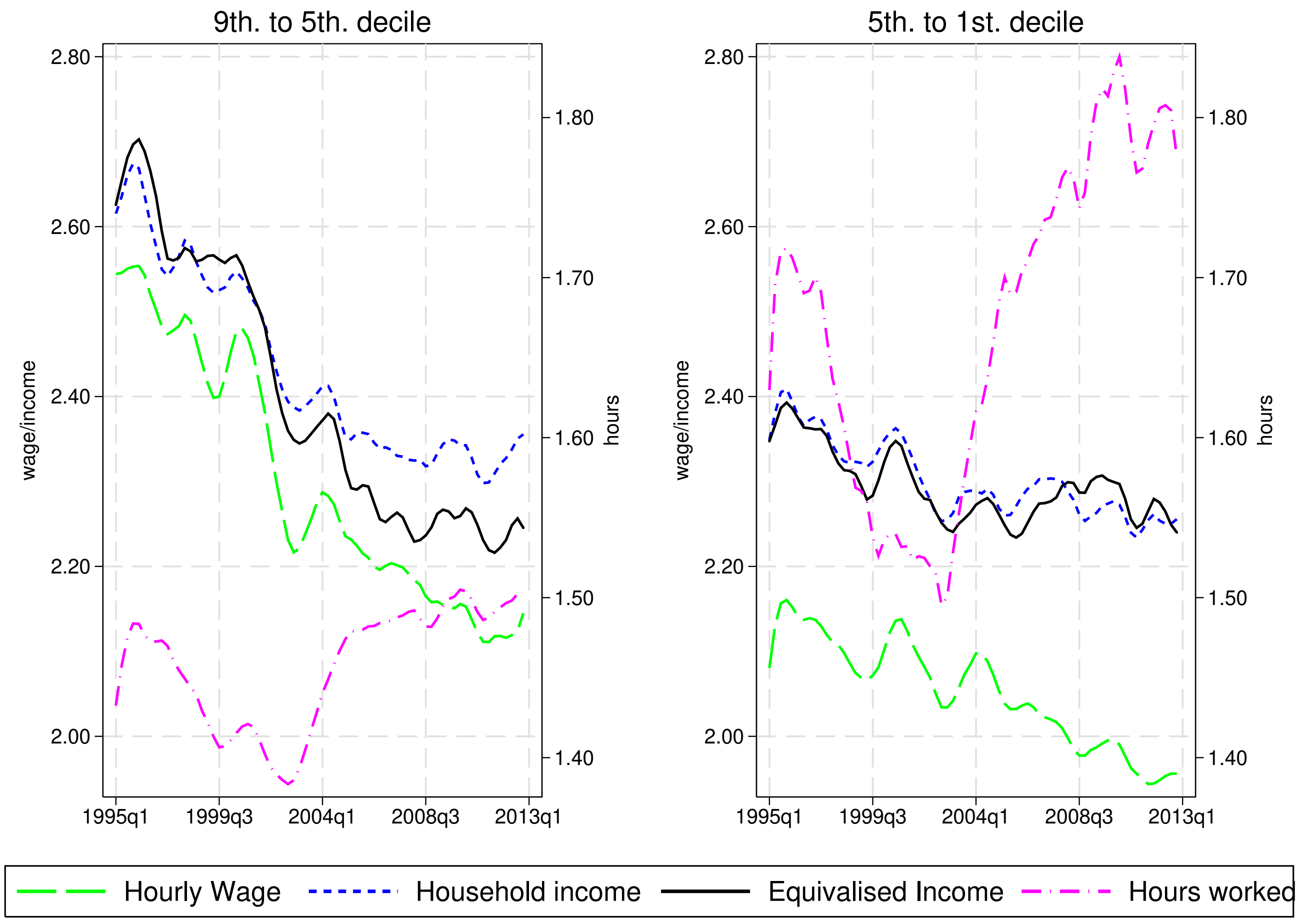
Note: In order to improve the readability of the figure, the series shown were smoothed using a non–parametric locally weighted regression with bandwidth equal to 0.1.

Figure 3 – Mexico 1990 – 2012: Evolution of household income inequality



Note: In order to improve the readability of the figure, the series shown were smoothed using a non–parametric locally weighted regression with bandwidth equal to 0.1.

Figure 4 – Mexico 1990 – 2012: Evolution of household income inequality (decomposition)



Note: In order to improve the readability of the figure, the series shown were smoothed using a non–parametric locally weighted regression with bandwidth equal to 0.1.

Figure 5 – Smoothed Monetary Policy Shocks
(DSGE model)

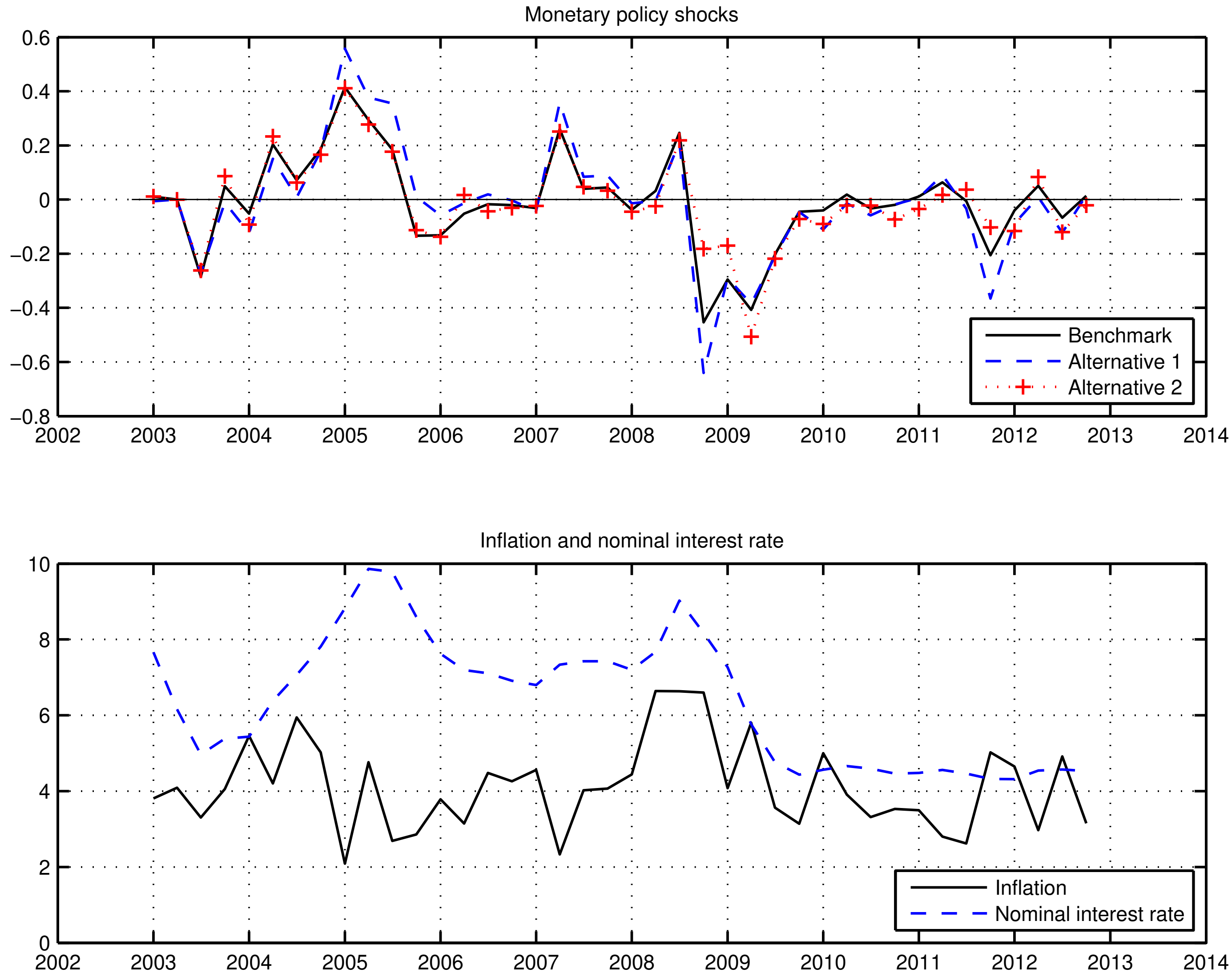


Figure 6 – Impulse Response Functions to a Monetary Policy Shock
(DSGE model)

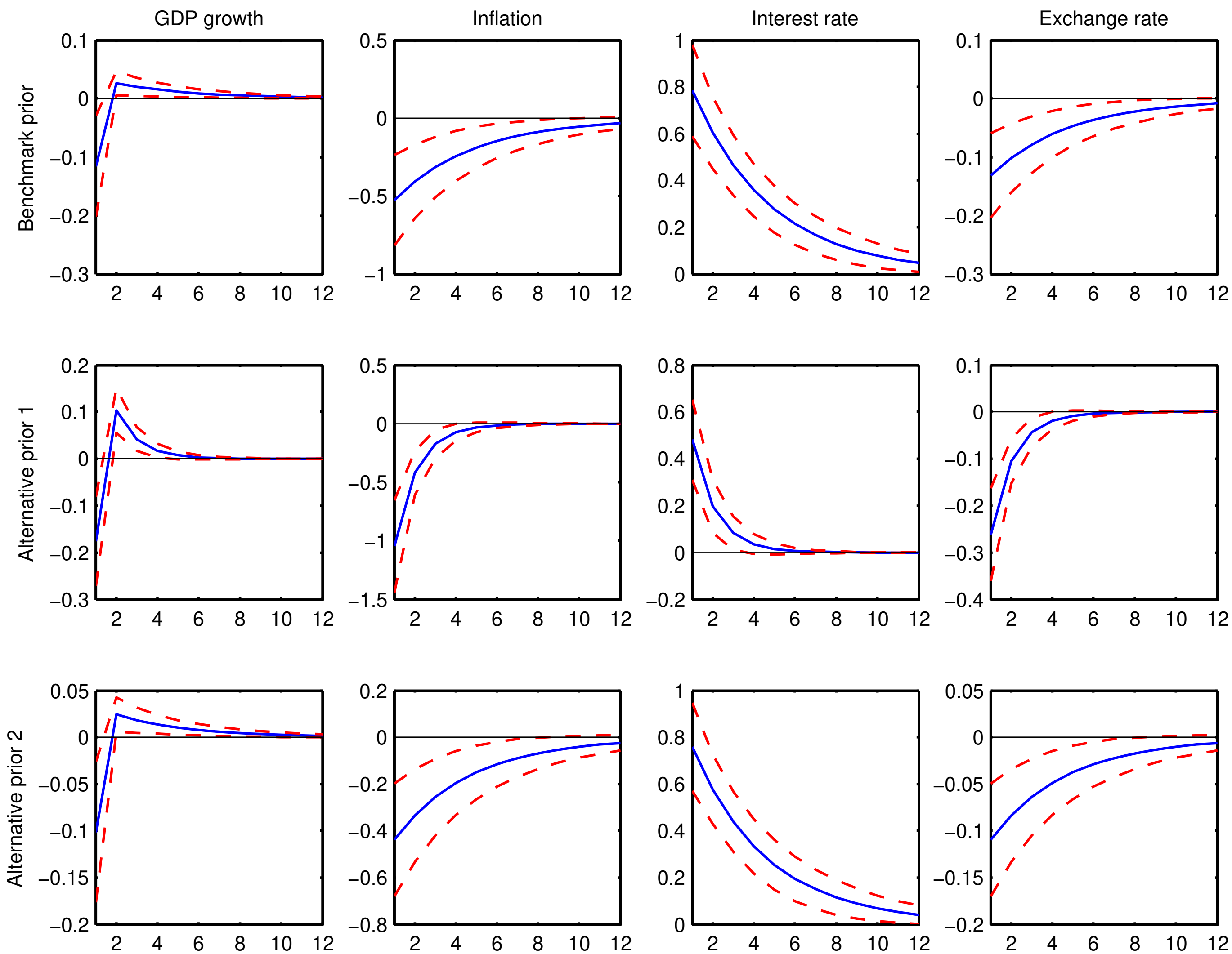


Figure 7 – Impulse Response Functions of Median Income to a Monetary Policy Shock
(DSGE model)

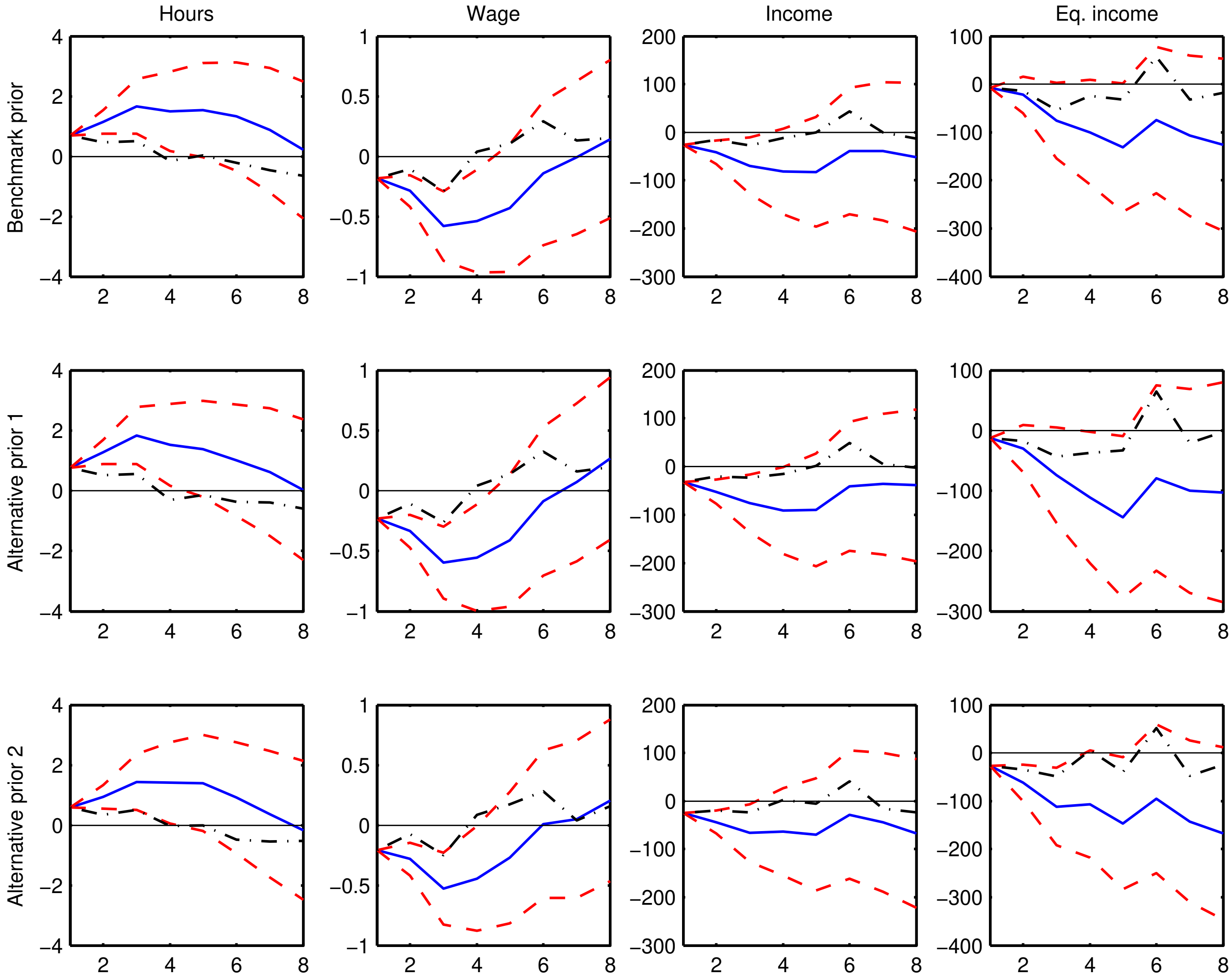


Figure 8 – Impulse Response Functions of Income Inequality to a Monetary Policy Shock
(DSGE Benchmark prior)

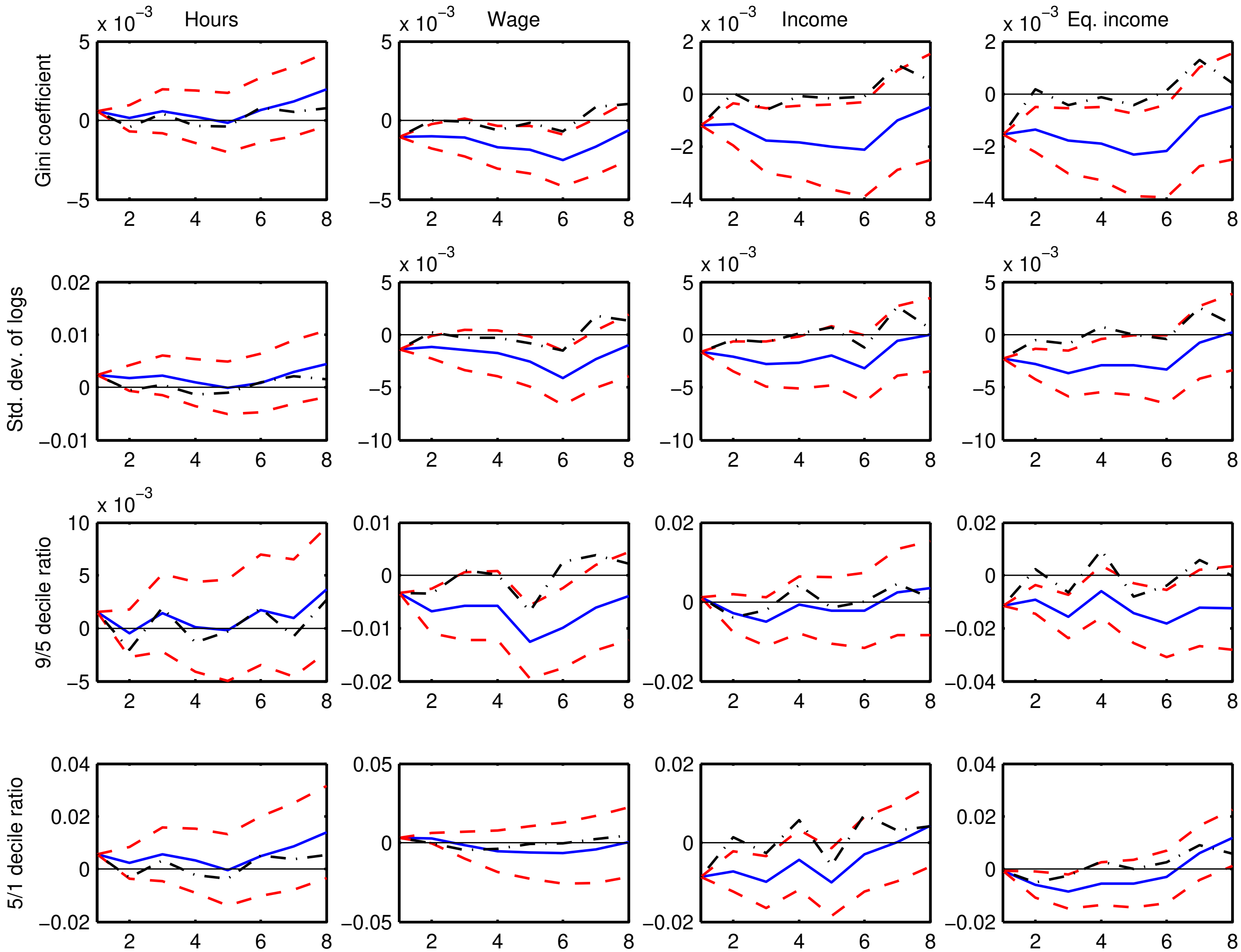


Figure 9 – Impulse Response Functions of Income Gini Coefficient to a Monetary Policy Shock
(SVAR: Monetary policy shock sign restrictions)

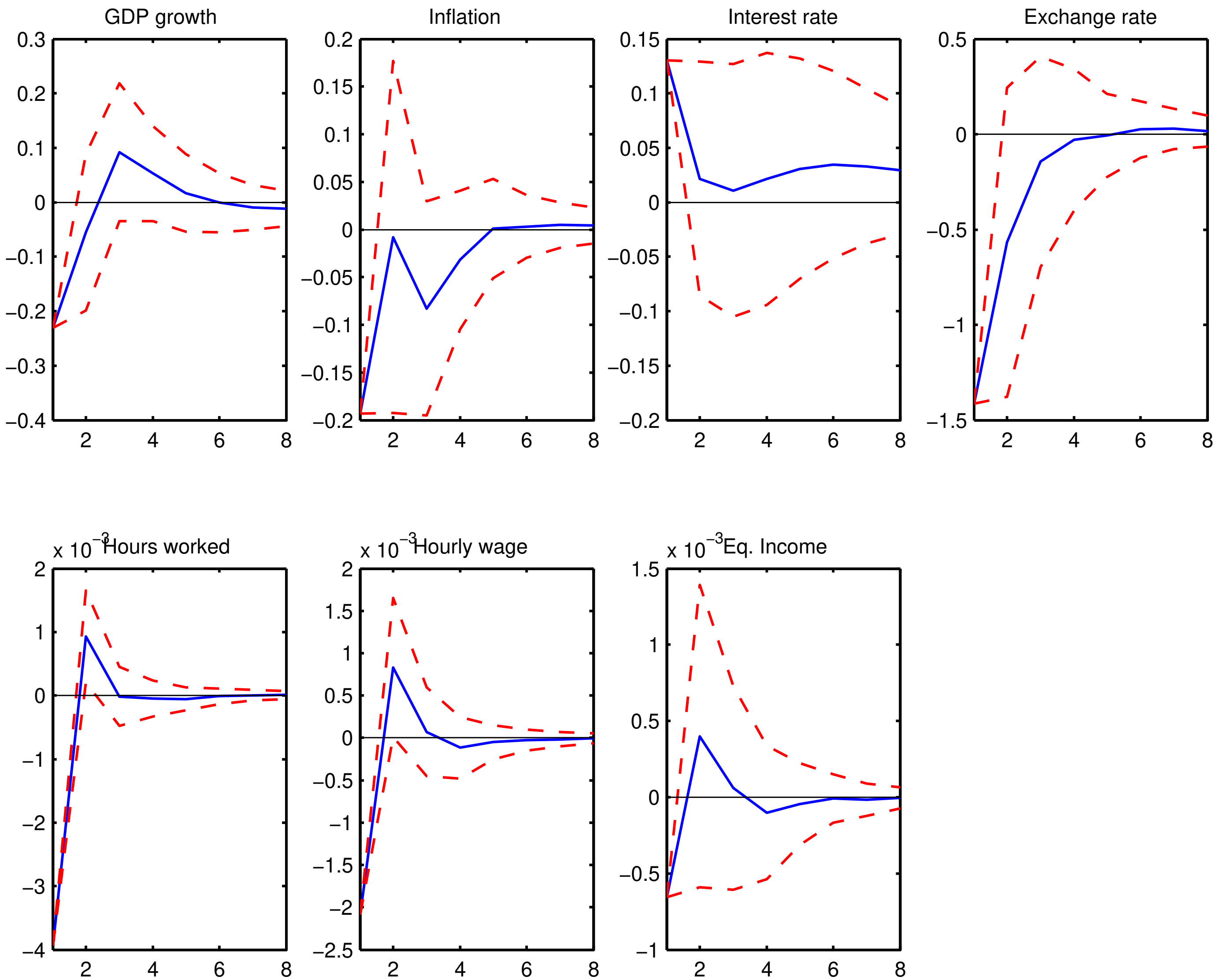


Figure 10 – Impulse Response Functions of Income Gini Coefficient to a Monetary Policy Shock (SVAR: Long–run and monetary policy shock sign restrictions)

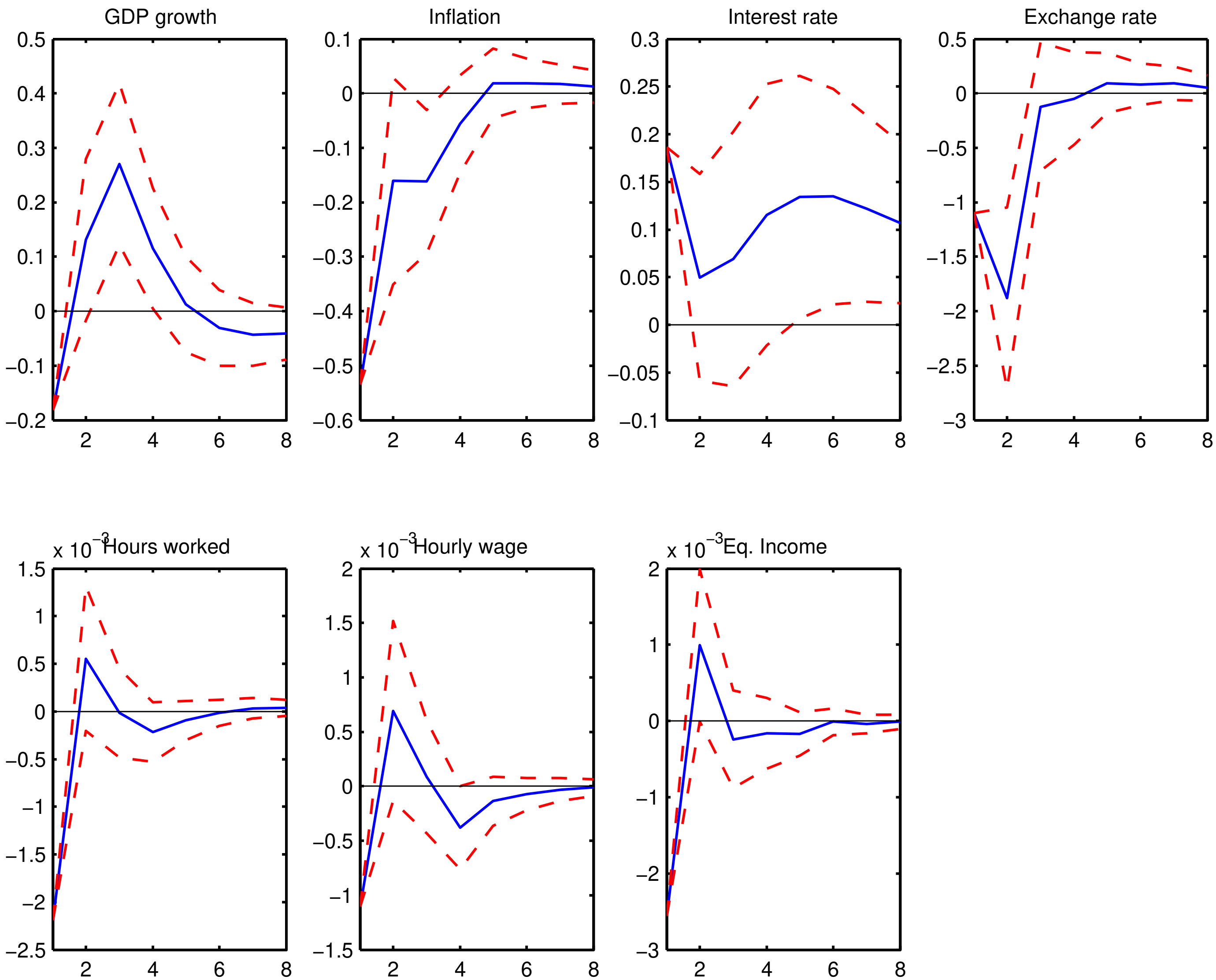


Figure 11 – Impulse Response Functions of Income Gini Coefficient to a Monetary Policy Shock (SVAR: Long–run and rest of shocks sign restrictions)