

Monetary policy and inequality in Mexico

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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 findings are robust to the particular measure of inequality used, as well as the procedure used to identify the policy shocks.

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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. The evidence suggests that heterogeneity is affected by monetary policy. Moreover, the transmission of monetary policy is also affected by the particular nature and magnitude of heterogeneity that characterises the model economy under study. Of particular interest to the design and implementation of policy, is that the rules that are optimal under the representative agent framework are 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 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 1995-2012¹: Mexico City, Guadalajara, Monterrey, Puebla, León, San Luis Potosí, Mérida, Chihuahua, Tampico, Veracruz, Acapulco, Aguascalientes, Morelia, Toluca, Saltillo,

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

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.

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

Although the changing nature of the features that characterise the labour force explain a significant portion of the fluctuations observed in households' income and its distribution, as documented by Binelli and Attanasio (2010) even after accounting for the changes in characteristics and their respective returns a large part remains 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 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. To account for the bias that results from self-selection of women into the labour force, instead of including a dichotomic variable to identify women, the residuals for women were estimated separately using the selection model proposed by Heckman (1976).

Figure 1 shows the evolution of selected regressor coefficients in the determination of the four income measures used, where the confidence intervals were constructed from robust standard errors estimated by clustering at individual industry-city pairs. In general terms, the coefficients have the expected sign and are statistically significant.

With respect to the coefficients on age, which are generally interpreted as returns on experience, they are found to be positive and marginally decreasing for hours, wages and total household income. It is interesting to note that across time there is an upward sloping trend in the returns of experience on hours worked, which is offset by a negative trend on hourly wage, resulting in a relatively flat profile for returns on household income across time. For the case of equivalised income, which accounts for the composition of the household, the returns on experience are not statistically significant.

For its part the coefficients on having finished high-school, which is a broad proxy for skill, is positive and statistically significant for wages and both measures and income. The negative sign on hours worked probably reflects the longer hours worked by lower-skilled individuals in an effort to compensate for significantly lower wages. It is worth highlighting that across time the evolution of the returns of skill are differentiated between men and women. For the case of hours, the returns for women have been relatively flat, whereas returns for men have been upward sloping. This means that while the gap between men of different skill levels has closed over the study period, it has remained constant for women.

With respect to the returns on wages, which in turn drive the results observed for both income measures, a downward trend is observed for men, whereas the returns for women at the end of 2013 are broadly similar to those observed at the beginning of the sample in 1995. This signifies that the skill return gap between men and women has diminished over the sample period. Moreover, considering their relative weight in the sample, the downward sloping trend for men also implies a reduction in the skill return gap for the full sample. As documented by Campos et al. (2014), this has been the main determinant of the observed reduction of inequality of income in Mexico over the last two decades.

Turning to the coefficients on whether the individual has access to social security, which proxy for the returns of working in the formal sector, these are positive and statistically significant across hours, wages and income measures. Moreover, the returns observed for women are significantly higher than those for men. In a similar fashion to the returns on experience, whereas the returns have been increasing in hours worked, they have been on a downward trend for wages, which result in a relatively flat profile for household income.

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, with a steeper decline observed for women. For its part, after a steep decline following the 1995 crisis, the hourly wage has been on a generally positive trend for both sexes. Notwithstanding the falling trend in hours worked, in general the dynamics of median household income have followed the fluctuations of wages. However, in contrast to wages which have roughly duplicated their value in constant terms over the period studied, despite a rapid expansion during the first year of the 2000's, household income has not managed to return to the levels observed in 1995. It is interesting to note that while individually, hours worked and hourly wages are relatively similar for men and women after controlling for observable characteristics, a sizeable gap remains for the median income of households headed by men and women.

Figure 3 summarises the evolution of inequality of labour supply and income under two alternative income measures: the Gini coefficient and the standard deviation of logarithms².

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

The dynamics of inequality under both measures is similar both across time and between sexes.

It is important to note that the Gini coefficients for income are significantly lower than what is usually reported for Mexico³. The lower value reported in this paper 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.

After declining significantly in the aftermath of the 1995 crisis there is an important rebound in the inequality of hours worked, which is measured off the right axis, starting in the early 2000 that continues to date. In contrast, inequality of wages and household income have decreased over the study period. The decomposition by sex indicates that while the reduction of inequality of wages is similar between men and women, the level of inequality of income of those households headed by women is larger than those headed by men, which is likely explained by the higher inequality observed for hours worked by women.

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 decile⁴. The top panel summarises the evolution of inequality for men and the bottom panel for women.

With respect to hours worked, the decomposition shows that the increase of inequality observed after 2003 is mainly explained by the inequality of hours worked by men, and to a lesser extent women, in the bottom half of the distribution. For the case of wages and household income, it can be seen that the main reduction of inequality has occurred in the top half of the distribution, which is related to the relative dispersion of hours worked. This is best exemplified by what happens to women in the bottom half of the distribution, who despite lower inequality of wages, not only have higher levels, but increasing inequality of household income, even after controlling for household size.

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 even in the presence of nominal frictions if a rate change is fully anticipated, the effect of the actual change on

³For example Esquivel et al. (2010) report, based on data from the income and expenditure survey the Gini coefficient for labour income peaking at 0.56 in 1995.

⁴It 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.

economic aggregates will be negligible. However if the actual change is different from expectations, depending on the sign and magnitude of the discrepancy as well as of the nature of nominal frictions, 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.

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 model is characterised by a representative agent which forms expectations rationally, a continuum of monopolistic intermediate goods producers which provide inputs to a representative competitive final goods producer. Trade takes place in the context of a small open economy. A crucial assumption of the model is that intermediate goods producers face restrictions on the frequency at which they can adjust their prices.

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:

$$y_t = E_t y_{t+1} - [\gamma + \alpha(2 - \alpha)(1 - \gamma)](R_t - E_t \pi_{t+1}) - \rho_A \Delta A_t - \alpha [\gamma + \alpha(2 - \alpha)(1 - \gamma)] E_t \Delta q_{t+1} + \alpha(2 - \alpha) \frac{1 - \gamma}{\gamma} E_t \Delta y_{t+1}^* \quad (1)$$

where the structural parameters are: γ 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 non-stationary technology process A_t are considered exogenous⁵.

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:

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

where $\bar{y}_t \equiv -\alpha(2 - \alpha)(1 - \gamma)/\gamma y_t^*$ denotes the level of output that would be observed in the absence of nominal frictions. In addition to γ and α , structural parameters include the

⁵As 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.

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:

$$\pi_t = \Delta e_t + (1 - \alpha)\Delta q_t + \pi_t^* \quad (3)$$

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:

$$R_t = \rho_R R_{t-1} + (1 - \rho_R)(\psi_\pi \pi_t + \psi_y y_t + \psi_e \Delta e_t) + \varepsilon_t^R \quad (4)$$

where in principle it is assumed that the policy parameters ψ_j for $j \in \{\pi, y, e\}$ are non-negative. The term ρ_R is a smoothing term included to match the persistence commonly observed in interest rates. The term ε_t^R 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) 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 $\theta = [\psi_\pi, \psi_y, \psi_e, \rho_R, \alpha, \beta, \kappa, \gamma, \rho_q, \rho_a, \rho_{y^*}, \rho_{\pi^*}, \sigma_r, \sigma_q, \sigma_a, \sigma_{y^*}, \sigma_{\pi^*}]'$, 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 filter⁶. 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 (γ) 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 variables⁷. 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 is more than proportional to the deviation observed.

Second, it is interesting to note that when the exchange rate policy parameter is not

⁶The Kalman filter recursion is initialised using a diffuse prior as in Jong (1991)

⁷Given 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.

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 under the benchmark and alternative 2 prior specifications, given the importance of imported goods and services in the domestic consumption basket. This could be related to the relatively high value of the point estimate for parameter for the exchange rate in the policy rule, which effectively dampens the pass-through of fluctuations in the price of imported goods.

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. 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 reflecting differences in the implied cost of adjustment for a given inflation reduction.

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.

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 model will be misspecified and the estimated responses 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 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 :

$$Y_{t+s} = \alpha^s + B_1^{s+1}Y_{t-1} + B_2^{s+1}Y_{t-2} + \dots + B_p^{s+1}Y_{t-p} + u_{t+s}^s \quad (5)$$

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 $\widehat{IR}(t, s, d) = \widehat{B}_1^s d$, 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 through 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 shown in figure 7 where the top panel summarises results for men and the bottom panel for women⁸.

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, however the response is differentiated by sex.

Whereas the results summarised in the first column of the top panel of 7 indicate that as a response of unanticipated increases in the nominal interest rate, hours worked by men increase significantly over six 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; the response of hours worked for women peaks between the second quarter, reflecting the quick recovery of wages after the negative contemporaneous shock.

From the third column it appears, however, that the observed increase in hours worked by men is not sufficient to keep household income from falling as a response to contractionary monetary policy shocks. In contrast the positive response of wages for women result in a positive response of both income measures of households headed by women. The rise in women's wages as a result of a monetary policy shock seems puzzling but could be explained by employers exploiting the wage gap in favour of women as a response to the shock.

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 for the full sample, and in figure 9 disaggregated by sex. In both figures the top two rows summarise the responses of respectively, the Gini coefficient and the standard deviation of logarithms to monetary policy shocks; whereas the bottom two rows summarise the responses of the decile ratios.

For the full sample the response of the inequality of hours worked, measured using both the Gini coefficient and the standard deviation of logs, is not statically significant. Once the response is disaggregated using the decile ratios, it can be seen that while inequality declines as a response to a contractionary monetary policy shock in the upper half of the distribution, the effect is offset by the increase in inequality in the bottom half of the distribution. The analysis of the response of hours worked by sex, however points to a contemporaneous increase in inequality across men and women and across the distribution.

Turning to the response of wage inequality, while the evidence for the full sample is only marginally significant, the responses by sex indicate that unanticipated increases in interest rates lead to a statistically significant reduction of wage inequality, which appears to be concentrated in the bottom half of the distribution.

The reduction in inequality as a response to monetary policy shocks is also observed for total and equivalised household income under both inequality measures, when using the full sample as well as for the case of women. For men the point estimates point to a similar effect, however the effect is only marginally significant, which may be explained by the opposite

⁸For reference the dash and dot lines depict the original, non-cumulative, responses to a unit monetary policy shock.

responses found using the decile ratios, which point to a reduction in inequality in the bottom half of the distribution and an increase in inequality in the top half of the distribution. The responses of the decile ratios for women suggest the effect of inequality reduction occurs primarily 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.

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 restrictive⁹. 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:

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

where $\mathbf{Y}'_t = [\Delta y, \pi, R, \Delta e]$ is a vector of variables observed at time t , $B(L) \equiv B_1L + B_2L^2 + \dots + B_pL^p$ 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 $u_t = Z\varepsilon_t$, $E\varepsilon_t\varepsilon'_t = \mathbf{I}$, and $ZZ' = \Sigma$. Typically there exists a multiplicity of Z matrices that represent the data, that is matrices that satisfy $ZZ' = \Sigma$, thus it is necessary to impose restrictions to identify a particular Z .

⁹Alternative restrictions include recursive identification as in Sims (1980), and restriction on the contemporaneous effect of shocks on system variables as in Galí (1992).

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 impulse responses of system (6) by estimating its coefficients subject to constraints 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:

$$A_+ = \begin{matrix} & \begin{matrix} as & ad & mp & fx \end{matrix} \\ \begin{matrix} \Delta y \\ \pi \\ R \\ \Delta e \end{matrix} & \begin{bmatrix} \times & 0 & 0 & 0 \\ \times & \times & \times & \times \\ \times & \times & \times & \times \\ \times & \times & \times & \times \end{bmatrix} \end{matrix} \quad (7)$$

where \times 's imply no restriction.

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

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

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.

For the evaluation of impact of monetary policy shocks on inequality, the following two alternative restrictions are imposed: 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

Figures 10 and 11 summarise the responses of, respectively, the set of restricted and unrestricted variables contained in the structural VAR¹⁰. In both figures the first row summarises the responses to shocks identified using a combination of the long-run restrictions discussed above and contemporaneous sign restrictions consistent with a DSGE-identified monetary policy shocks, and the second row the responses to shocks identified using a combination of long-run restrictions and contemporaneous sign restrictions consistent with DSGE-identified innovations different from monetary policy shocks.

Reflecting the relatively short lag-lengths selected by the Akaike information criterion, 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 two alternative restrictions, the signs of the responses of inflation, interest and exchange rates are consistent with the previous results. That is a positive, contractionary, monetary policy shock contemporaneously decreases inflation, increases the nominal interest rate and causes an appreciation of the exchange rate. Although the response of GDP growth is negative when the combination of long-run and sign restrictions to the response to a monetary policy shock; under the combination of long-run and sign restrictions on shocks different from monetary policy shocks, the contemporaneous response of GDP growth is positive, which although somewhat surprising is consistent with the findings of Uhlig (2005).

Turning to the responses of the Gini coefficient of hours, wages and equivalised income summarised in figure 11, although the responses of the inequality of hours worked and wages are not congruent under the alternative shock identification restrictions, the responses of the Gini coefficient of equivalised income are negative and statistically significant on impact under the two specifications.

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

¹⁰The corresponding confidence intervals were bootstrapped using the OLS residuals of the VAR estimated under the selected factorisation.

and the methodology used estimate their impact, unanticipated increases in nominal interest rates reduce the inequality of equivalised household labour income.

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.

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A Tables and figures

A.1 Tables

Table 1 – Prior distributions

Parameter	Benchmark prior				Alternative prior 1				Alternative prior 2			
	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
γ	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.

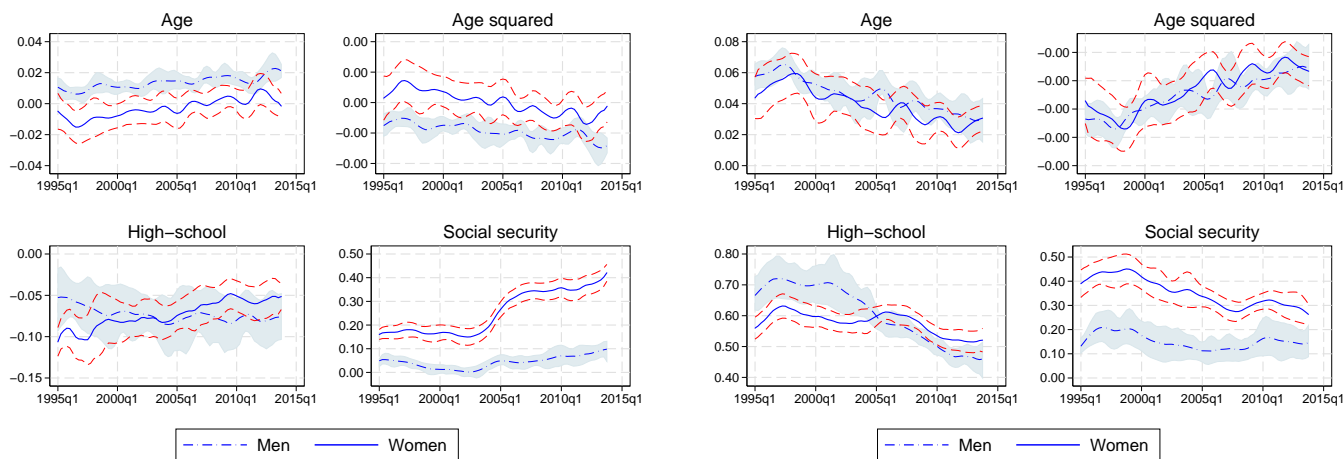
Table 2 – Parameter estimation results

Parameter	Benchmark		Alternative 1		Alternative 2	
	Mean	90% Interval	Alternative 1	90% Interval	Alternative 2	90% Interval
ψ_π	2.1500	[0.9879, 3.1672]	2.2558	[1.5266, 2.9913]	1.6516	[1.2799, 2.0390]
ψ_y	0.6562	[0.1846, 1.0872]	0.2025	[0.0434, 0.3590]	0.4380	[0.0502, 0.7943]
ψ_e	0.2174	[0.0667, 0.3654]	0.1385	[0.0439, 0.2289]	0.1896	[-0.1157, 0.4895]
ρ_R	0.8739	[0.8209, 0.9270]	0.7036	[0.6007, 0.8124]	0.8601	[0.8069, 0.9110]
α	0.0072	[0.0003, 0.0180]	0.0849	[0.0471, 0.1219]	0.0081	[0.0007, 0.0187]
r_{ss}	2.0429	[0.4625, 3.4638]	2.4817	[0.9414, 4.0352]	1.9592	[0.5094, 3.3546]
κ	0.0270	[0.0143, 0.0381]	0.5487	[0.2409, 0.8425]	0.0264	[0.0155, 0.0371]
γ	0.0192	[0.0057, 0.0322]	0.1388	[0.0752, 0.1975]	0.0208	[0.0082, 0.0334]
ρ_q	0.1561	[0.0401, 0.2677]	0.2837	[0.0923, 0.4761]	0.1574	[0.0418, 0.2684]
ρ_A	0.2149	[0.1613, 0.2684]	0.1977	[0.1520, 0.2428]	0.2184	[0.1683, 0.2684]
ρ_{y^*}	0.8081	[0.5893, 0.9711]	0.8430	[0.7835, 0.9037]	0.8806	[0.7799, 0.9725]
ρ_{π^*}	0.1544	[0.0038, 0.2932]	0.4664	[0.2982, 0.6306]	0.1507	[0.0054, 0.2886]
σ_R	0.7833	[0.6356, 0.9303]	1.0799	[0.7861, 1.3689]	0.7627	[0.6235, 0.9078]
σ_q	4.0571	[3.3537, 4.7381]	4.0911	[3.3404, 4.7814]	4.0506	[3.3287, 4.7520]
σ_A	1.1560	[0.9306, 1.3834]	1.2366	[0.9896, 1.4684]	1.1568	[0.9321, 1.3834]
σ_{y^*}	0.6380	[0.2611, 1.0761]	0.7006	[0.3636, 1.0568]	0.5339	[0.2421, 0.8203]
σ_{π^*}	2.9784	[2.4464, 3.4926]	3.2869	[2.6852, 3.8647]	3.0002	[2.4518, 3.5343]

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. For each prior the scale parameter of the jumping distribution’s covariance matrix was adjusted to ensure that the acceptance ratio of the algorithm fell within the 25%-33% range.

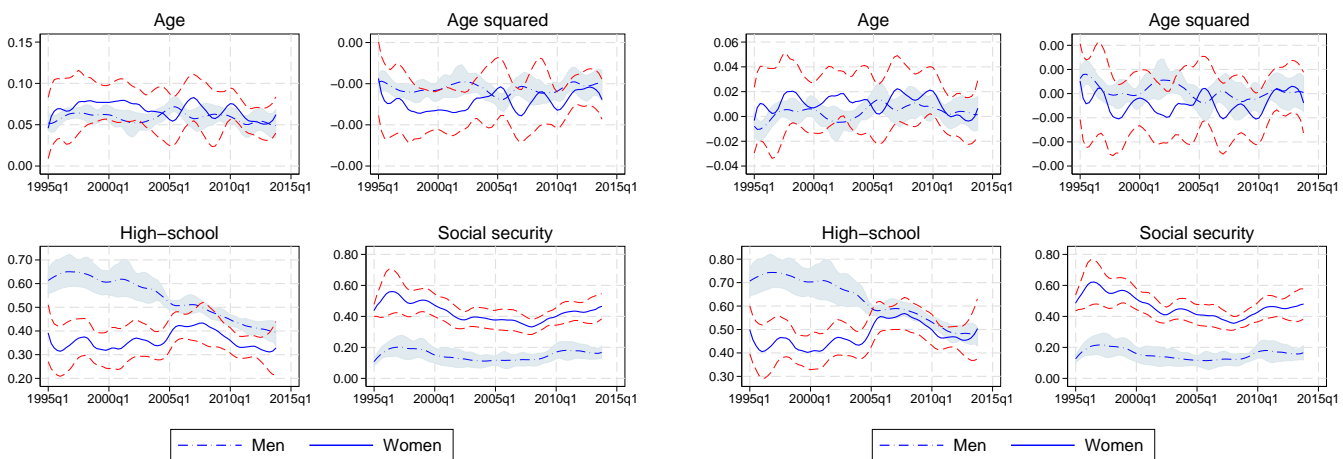
A.2 Figures

Figure 1 – Regression coefficients



(a) Weekly hours worked

(b) Hourly wages

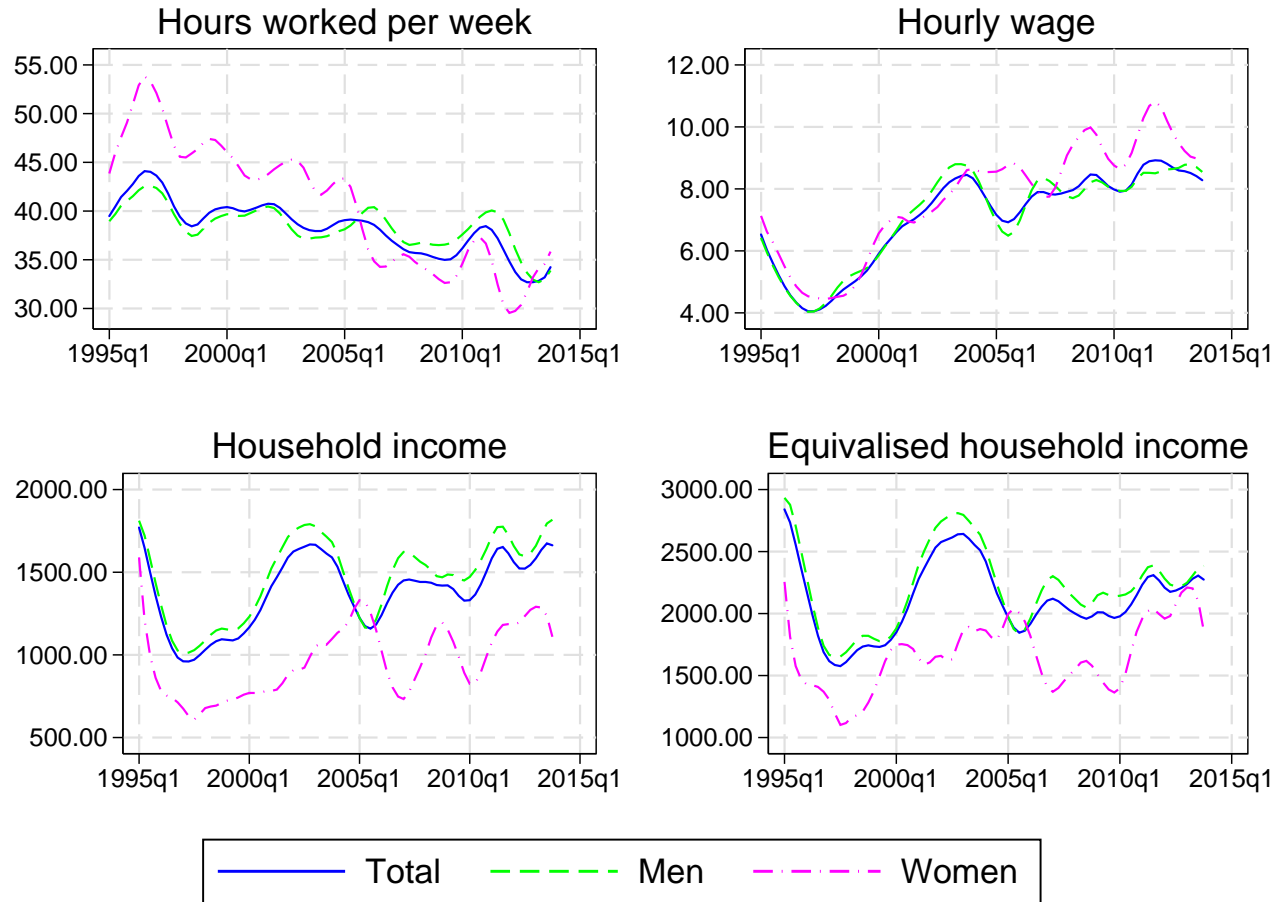


(c) Household income

(d) Equivalised 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 0.15.

Figure 2 – Mexico 1995 – 2013: Evolution of median 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 0.15.

Figure 3 – Mexico 1995 – 2012: Evolution of household income inequality



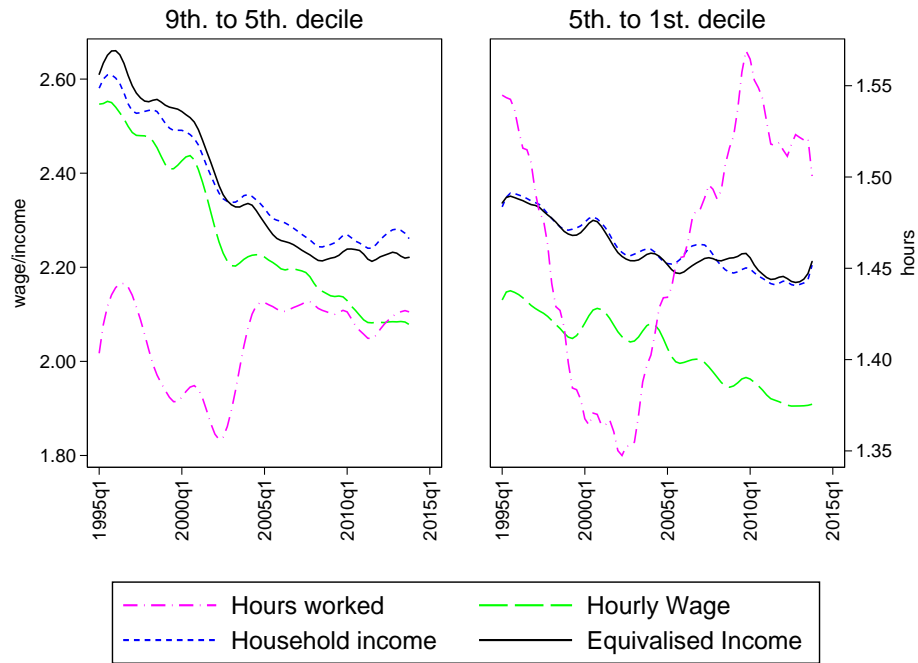
(a) Gini coefficient



(b) Standard deviation of logarithms

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 0.15.

Figure 4 – Mexico 1995 – 2012: Evolution of household income inequality (decomposition)



(a) Men



(b) Women

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 0.15.

Figure 5 – Smoothed Monetary Policy Shocks
(DSGE model)

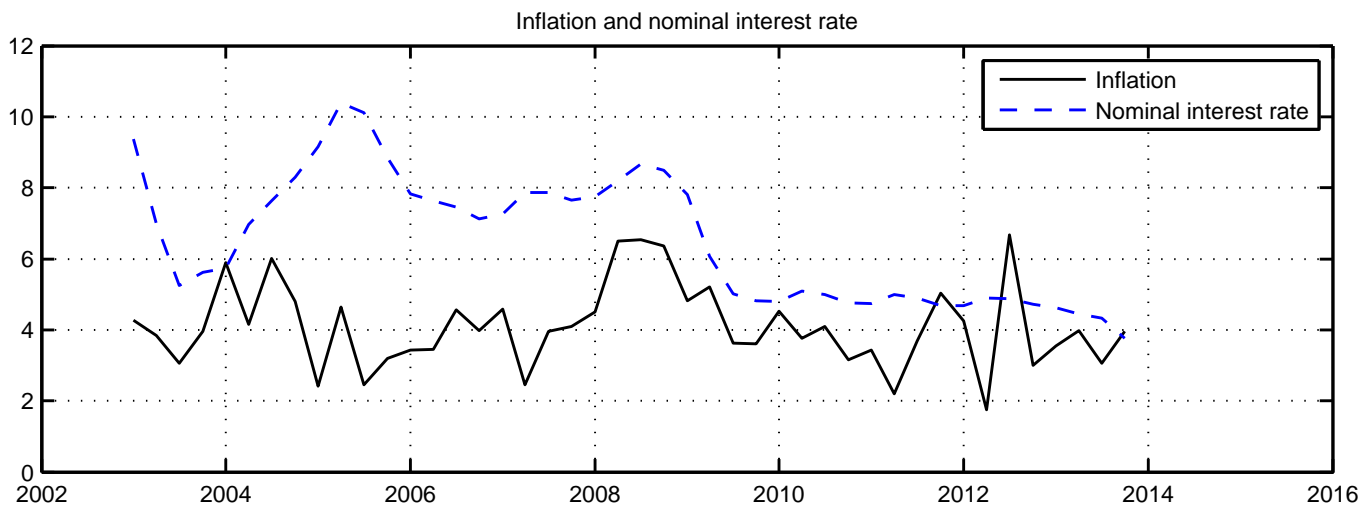
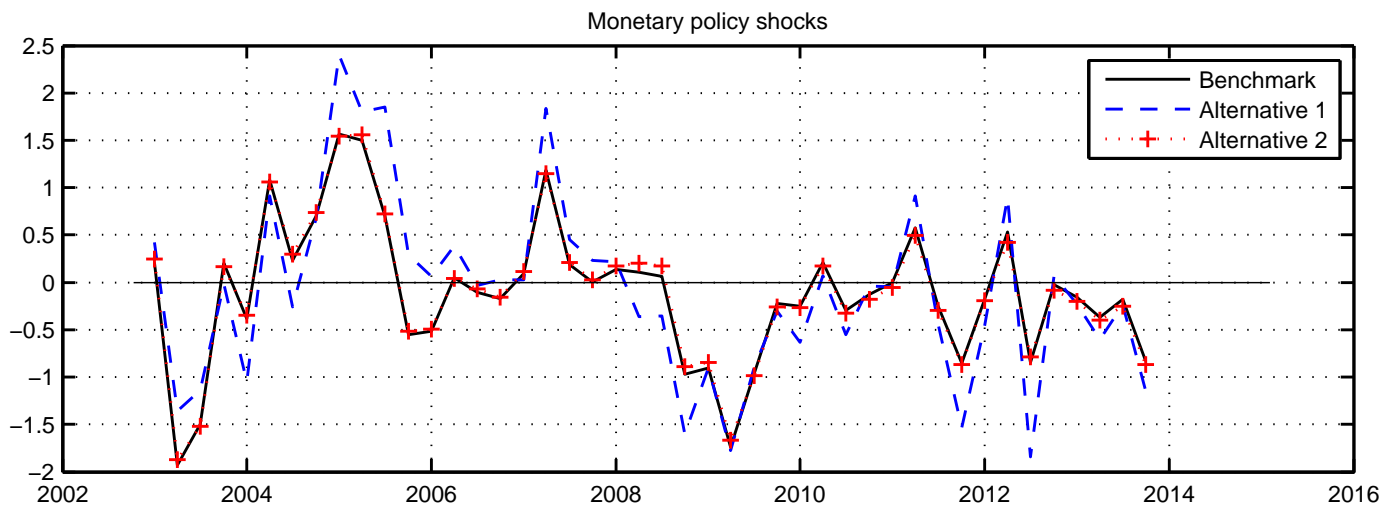


Figure 6 – Impulse Response Functions to a Monetary Policy Shock
(DSGE model)

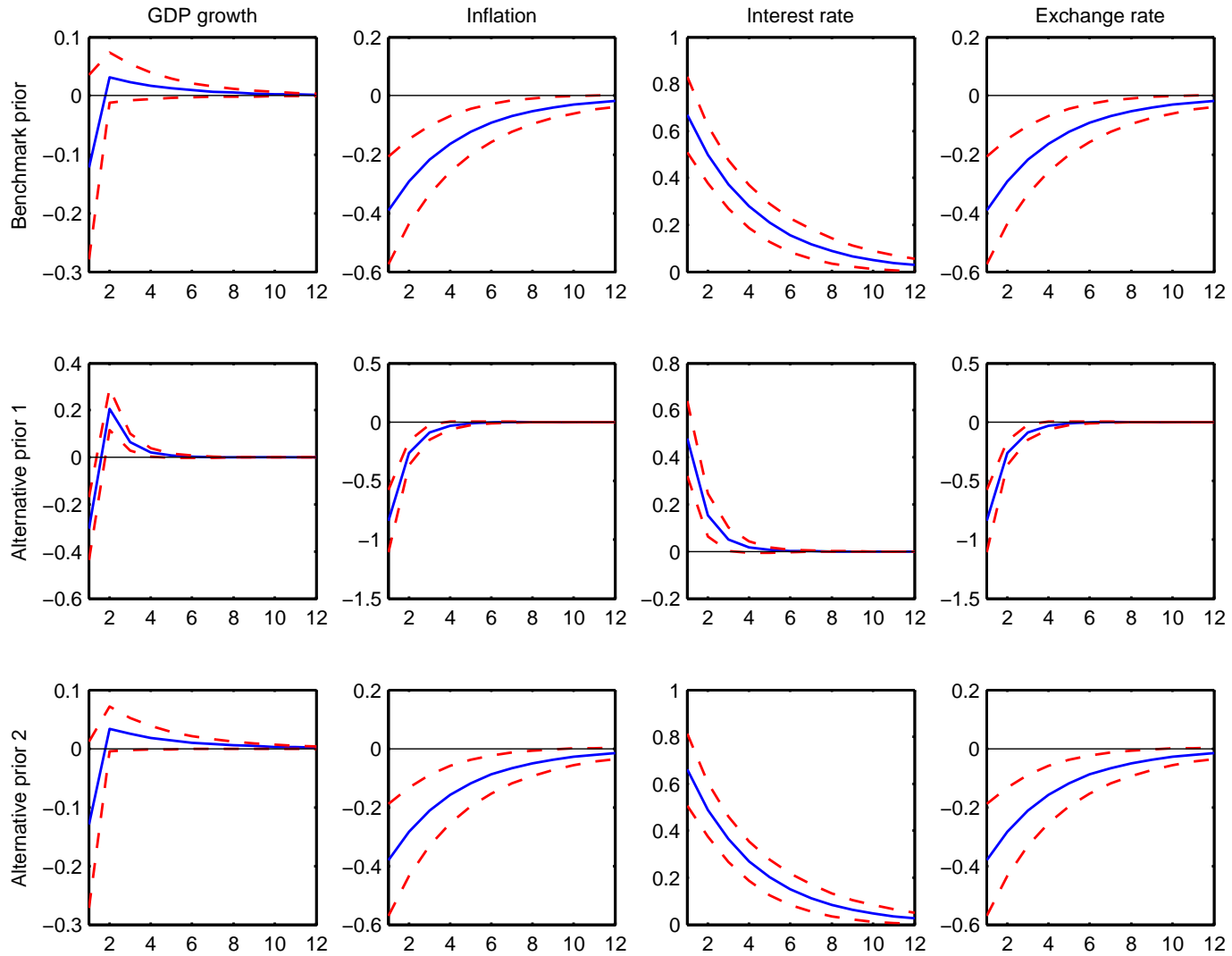
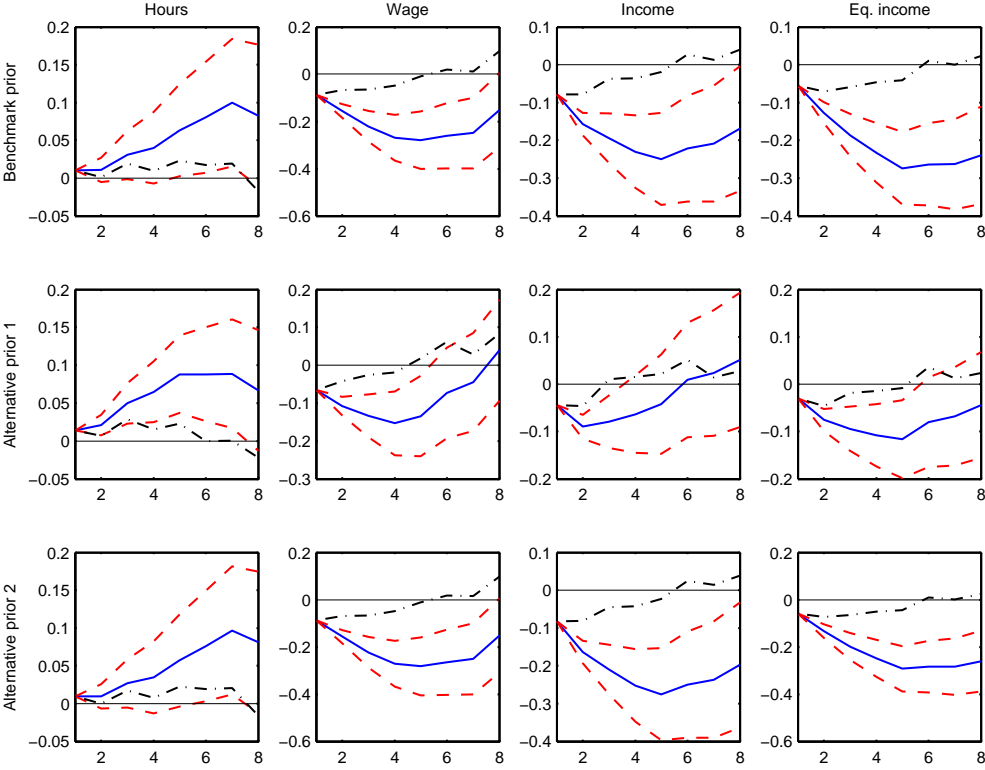
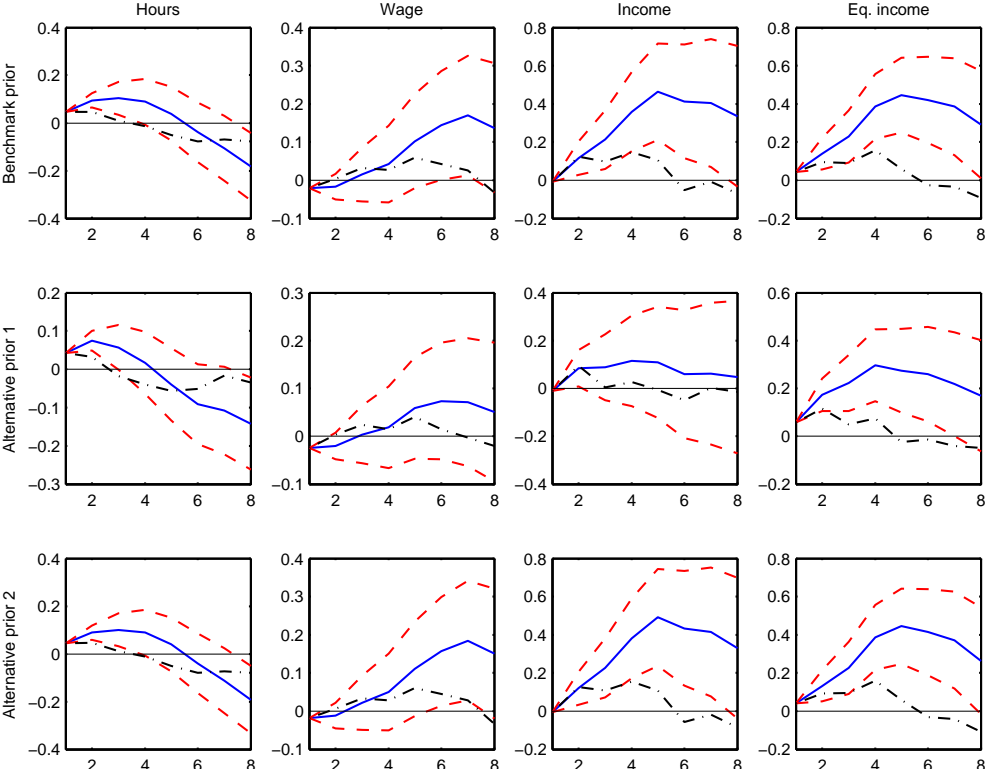


Figure 7 – Impulse Response Functions of Median Income to a Monetary Policy Shock (DSGE model – by sex)



(a) Men



(b) Women

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

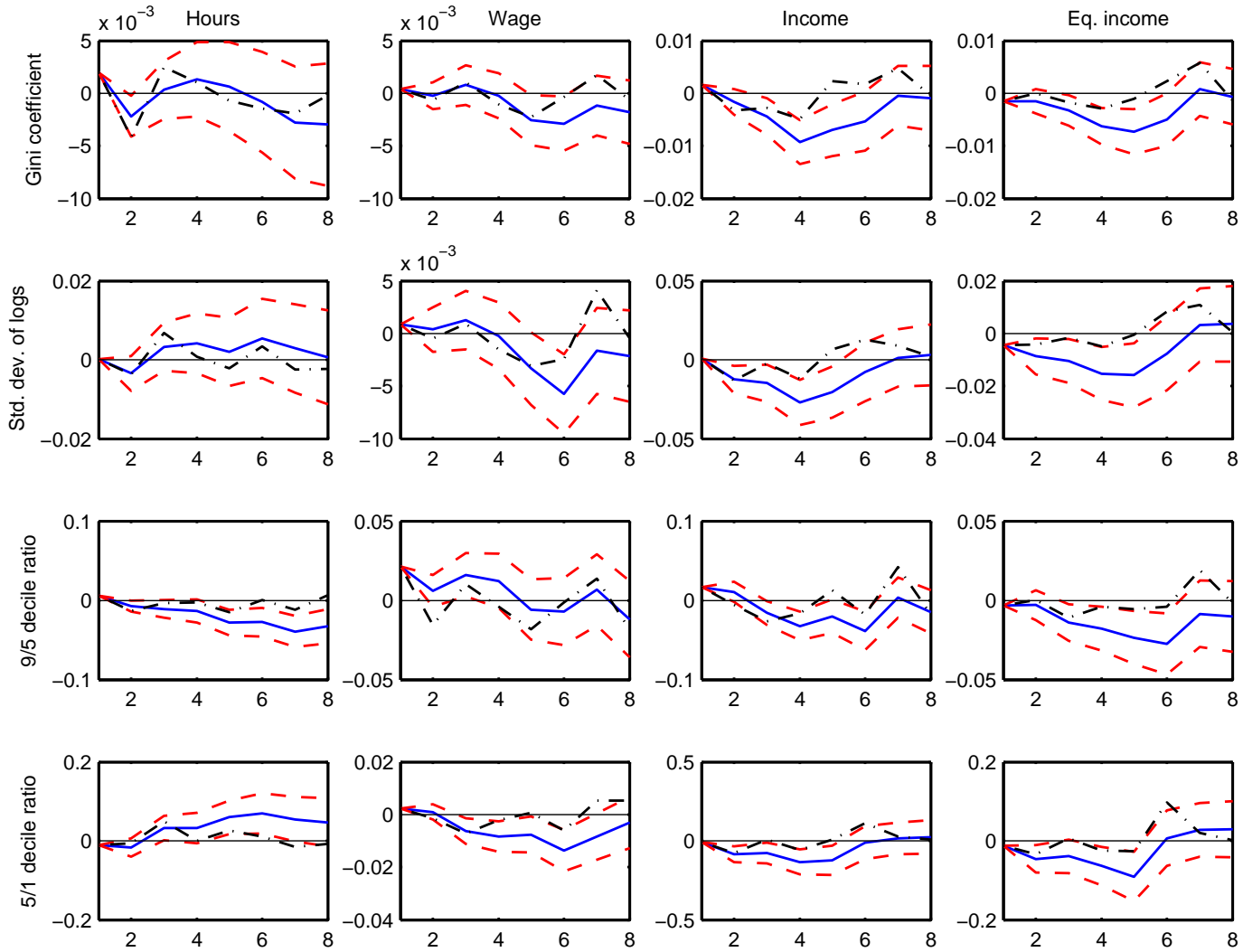


Figure 9 – Impulse Response Functions of Income Inequality to a Monetary Policy Shock (DSGE Benchmark prior – by sex)

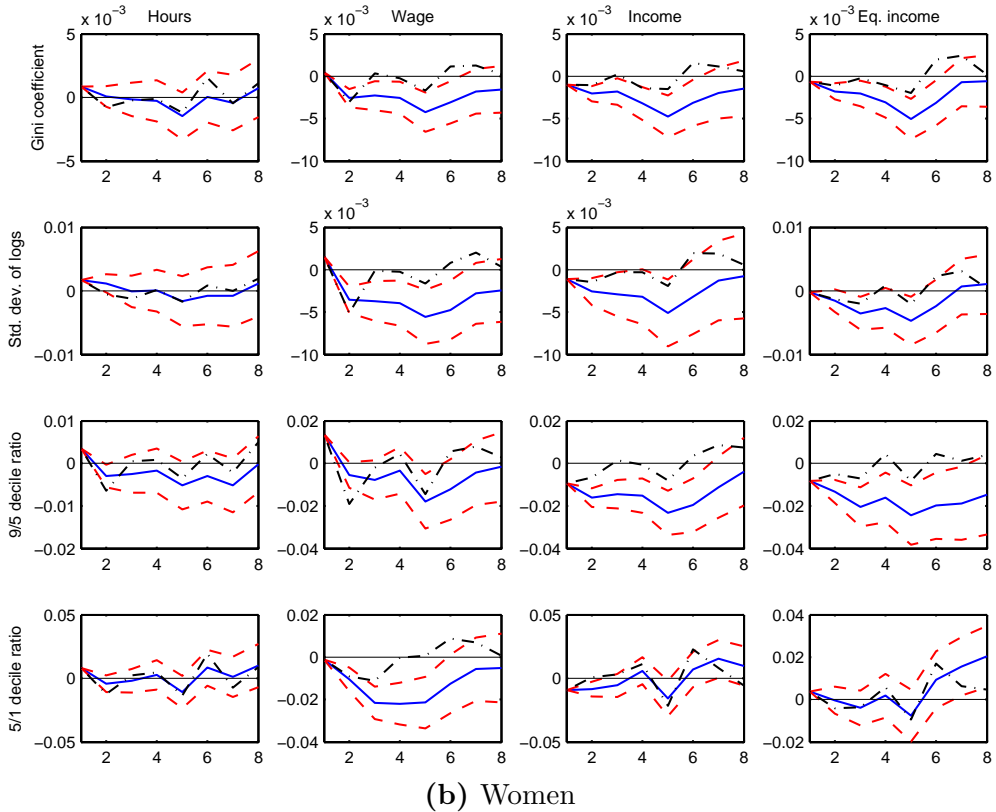
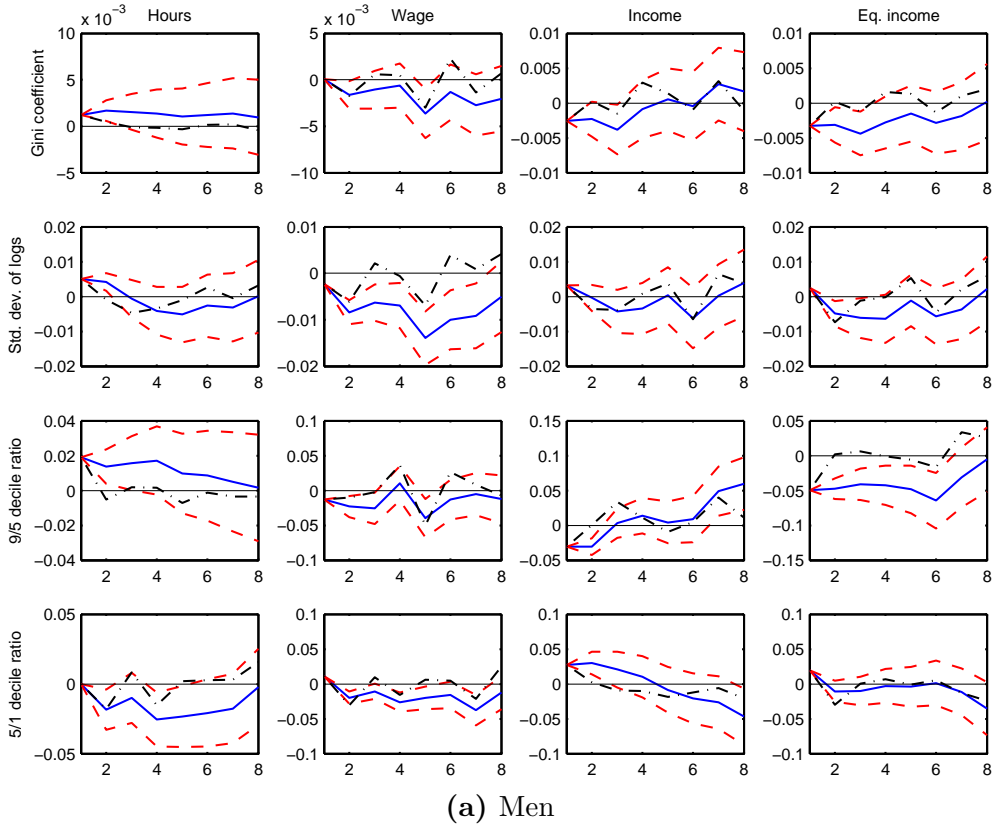


Figure 10 – Impulse Response Functions of Restricted Variables to a Monetary Policy Shock

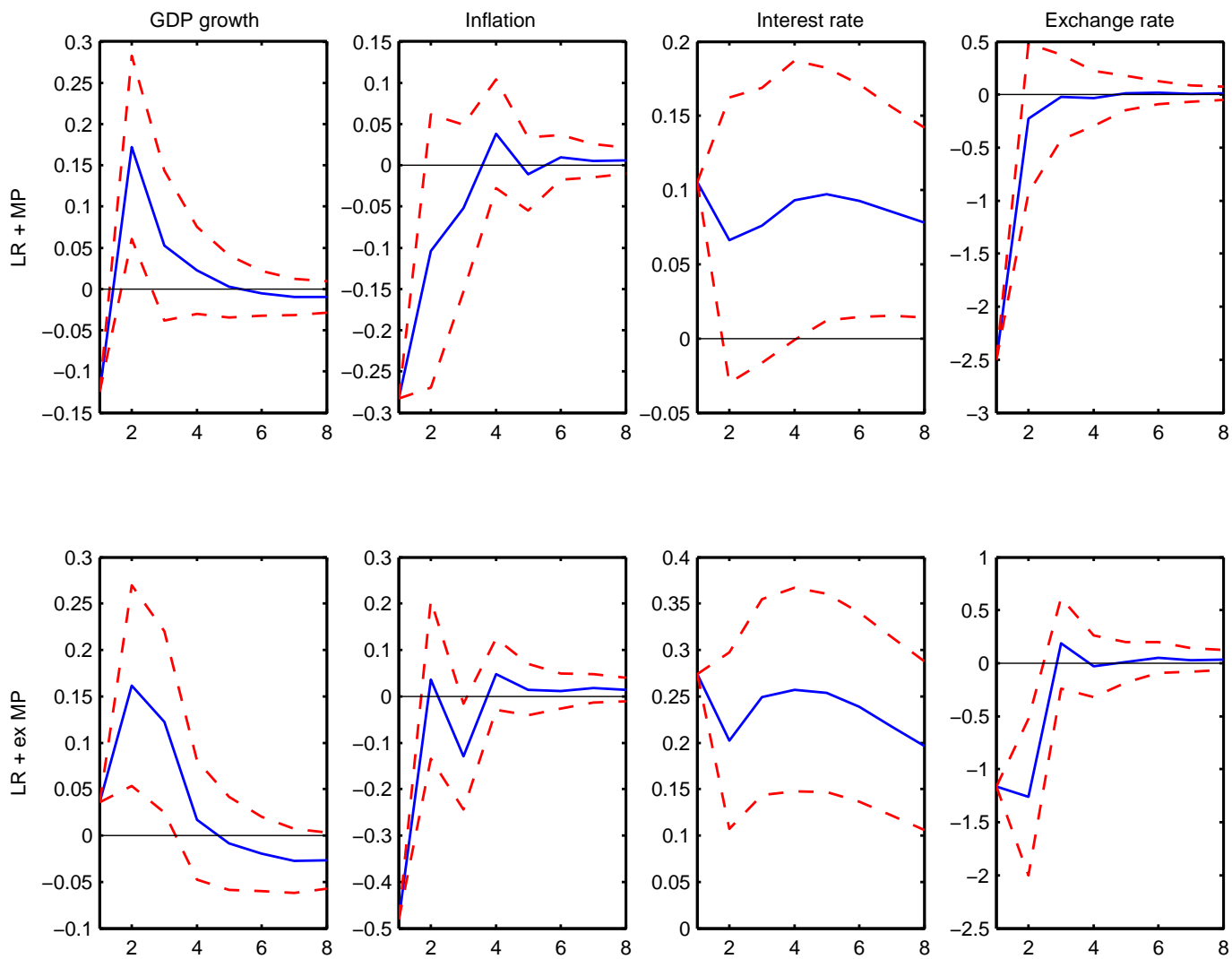


Figure 11 – Impulse Response Functions of Unrestricted Variables to a Monetary Policy Shock

