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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's labour income inequality in Mexico. The results indicate that inequality of aggregate household's labour–income increases as a result of an unanticipated increase in nominal interest rate. However, the result is differentiated across labour markets, as well as across the distribution of income, with inequality declining among households in the bottom half of the distribution, whose head is employed in the informal labour market. The findings 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
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1 Introduction

Although the study 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 models for its study. Thus, while welfare considerations are the normative criterion for the evaluation of monetary policy, the utilisation 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 ones for employment and financial services, which result in substantial heterogeneity in the capacities of households to insure against idiosyncratic shocks, such as spells of unemployment Blundell et al. (2008).

Brzoza-Brzezina et al. (2013) and Guvenen (2011) review recent attempts to introduce heterogeneous households into monetary policy models. The surveyed 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 household 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 developing countries. 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 for the Mexican case, such as data on wealth; and some of it is only available sporadically, as is the case for data on consumption which is published on a biennial basis. Thus, the focus is placed on the effects 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.

The findings of the model indicate that unanticipated increases in the nominal interest rate rise overall household's labour income inequality. The result reflects the dynamics of labour income of households in the top half of the distribution whose heads work in the formal sector. In contrast, labour income inequality decreases in response to contractionary monetary policy shocks among households located in the bottom half or the income distribution, whose heads work in the informal sector. However, it must be stressed that the reduced inequality reflects lower income levels, so it is by no means a socially desirable outcome.

While the findings are broadly in line with results found by Coibion et al. (2012) and Gornemann et al. (2015), the specific responses of wages and income between the formal and informal sector, are at odds with the mechanics of the income composition channel. A possible explanation could be that, as documented by Campos-Vazquez (2010), as a result of adverse macroeconomic shocks young and unskilled workers are the most likely to be forced to migrate from the formal to informal sector. Since this group of workers has benefited the most from the fall in the returns to skill over the last two decades, their migration from the formal to the informal sector could account for the rise (fall) in inequality in the formal (informal) sector.

The paper's contributions are threefold. Empirically, to the best of my knowledge the paper is the first to explore the effects of monetary policy on labour income inequality for the case of Mexico, highlighting the heterogeneous effect between households whose heads work in the formal or informal labour market. Second, from a methodological perspective the paper demonstrates that in order to account for informality it is not sufficient to measure an overall formality premium, through the use of indicator variables. Instead the evidence indicates, that the returns of other observable characteristics, such as age and schooling, are statistically different between the formal and informal labour markets. Third, the paper documents the evolution of cross-sectional facts of labour income in Mexico, accounting for observable characteristics, thus updating the work of Binelli and Attanasio (2010).

The rest of the document is organised as follows. Section 2 analyses the evolution of labour income inequality in Mexico over the period 1995–2014. Next, section 3 details the benchmark shock–identification procedure, while the results are discussed in section 4. Section 5 concludes.

2 The evolution of labour–income inequality

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). Notwithstanding the recent decline, inequality remains elevated across Latin America.

In the case of Mexico, income inequality has exhibited a declining trend that began in 1994 and continues to date, although at a substantial lower rate since 2011. 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.

To put the importance of labour-income in context, table 1 summarises the main sources of household income by decile from the most recently available survey data. On average two thirds of household's income stems form labour earnings, although its relative importance is very heterogeneous across the distribution ranging from 38.5% in the first decile to 69,8% in the ninth decile. Within labour income, the composition between salaried work and independent work earnings is also heterogeneous with independent work earnings above the mean for deciles I through V. This is relevant because as discussed by Porta and Shleifer (2014) the vast majority of independent workers work in the informal sector, which is characterised by very low productivity levels and has little, if any, access to social security and formal financial services. As expected, transfers represent an important proportion of current income for the poorest households, with that proportion reaching 36.4% in the first decile. However it should be noted that even for households in the top decile, transfers still account for 15.3% of income. Finally, in contrast to the relevance of capital income, which increases monotonically with income levels, the proportion represented by imputed rent income declines as income increases.

Before summarising 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 labour–income 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 third quarter of 2014[1]. 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 2014.III, 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 1995-2014[2]: 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, as the activity is not considered employment for official figures. 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.

[1] Although survey data is available through the fourth quarter of 2015, only data up to the third quarter of 2014 was used because as a result of the constitutional reform that increased the minimum working age from 14 to 15 years of age, the most recent survey waves are not directly comparable to those corresponding to quarters before 2004.IV

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

At the household level, 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 Labour–income dynamics

Figure 1 summarises the evolution of median hourly wages, total and equivalised income over the period 1995.I through 2014.III. It can be seen that labour income dynamics largely reflect fluctuations of hourly wages, which after a sharp decline in the aftermath of the 1994–1995 balance of payments crisis did not recover their pre-crisis levels until the early 2000's. The recovery in labour–income was interrupted by the onset of the global financial crisis of 2008–2009, which caused another decline in households' labour income which stretches to the end of the study period. Reflecting higher wages, households whose head is formally employed, as proxied by access to social security, consistently exhibit higher incomes than those in the informal sector. While the formal–informal gap narrows, in absolute terms, when equivalised income is considered, the gap has remained fairly constant across the study period.

Since the results for total and equivalised household income are very similar, for brevity of exposition in the remainder of the document the analytical focus is placed on the evolution of hourly wages and equivalised income, which are comparable across individuals and households, respectively.

Regarding the inequality of income, the evolution of the Gini coefficient is shown in figure 2, where the downward trend documented elsewhere is confirmed. It is interesting to note that although inequality within the informal sector remains larger than in the formal sector, its decline over time has been steeper. As discussed below this reflects the changing returns to education. It is worthwhile noting that the relative stagnation of the decline in inequality that can be observed towards the end of the period studies, originates in the informal sector.

In order to explore the dynamics of labour income across the distribution, figure 3 shows the evolution of the ratio of the incomes of households in the ninth decile with respect to those in the fifth decile, and the ratio of incomes of households in the fifth decile with respect to those in the first decile. There, it can be seen that regardless of the labour market, the aggregate reduction in inequality has occurred almost exclusively in the top half of the distribution, with a larger reduction observed in the informal labour market. In contrast, inequality within the bottom half of the distribution has remained broadly constant in the formal sector and exhibits a creeping upward trend in the informal sector.

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.

Considering the findings of Fernández and Meza (2015) which indicate that informal employment in Mexico is countercyclical, and inversely correlated to formal employment, the sample is split into formal and informal sectors, where as before formality is proxied by having access to social security.

For each sub–sample, in order to obtain residual measures of households' income and its dispersion, labour income measures (hourly wages and equivalised income) are regressed on a square polynomial on age, which serves as a proxy for experience, a dichotomic variable that indicates whether the individual finished middle–school, which is a broad proxy for ability, as well as a set of dichotomic variables that account for the marital status of the individual, the heterogeneity of the urban area where the households are located, as well as of the industry where individuals work in. 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 the individual's sex, the residuals for women were estimated separately using the selection model proposed by Heckman (1976).

Figure 4 illustrates the evolution of the coefficients on age and schooling in the determination of the income measures used, where the confidence intervals were constructed from robust standard errors estimated by clustering at individual industry-city pairs[3].

Overall, results indicate that the returns of experience on wages and household income have remained roughly constant over time, with the exception of the returns on men's wages which show a downward trend. It is interesting to note that whereas the returns on experience are not statistically different between the formal and informal sectors for wages of men and women, as well as equivalised income of households headed by men, the returns to experience on female–headed household income in the formal sector are negative, perhaps reflecting discrimination in the workplace.

For their part, returns to skill exhibit a clear downward trend for both sexes, with the decline being more pronounced for women. 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. The fact that the returns to skills have been more pronounced for informal workers explains the magnitude of the decline in inequality within this segment of the labour market.

It is important to note that the differences in the evolution of coefficients across labour markets indicate that only including an indicator variable to account for the effect of informality would bias the estimated coefficients as well as the residual inequality measures.

Panel (a) in figure 5 shows the evolution of median residual labour–income and the

[3]To avoid cluttering the graph only the linear coefficient on age is shown, the quadratic coefficient is negative as expected, indicating declining marginal returns to experience.

resulting Gini coefficient. The first feature to note is that the clear association between labour income and the business cycle, which was evident in the unconditional data, dissipates to a large extent. Second, the residual income measures are markedly more volatile, more so for the case of formal labour market income, perhaps pointing to the changing composition of the labour force within the formal sector across the business cycle. Third, the dynamics of aggregate residual median income closely track the evolution of labour income originating in the informal sector, underlining its magnitude. Finally, the gap between formal and informal income widens with respect to unconditional data and becomes more volatile.

Regarding inequality, measured by the Gini coefficient, panel (b) in figure 5 shows that once observable characteristics are accounted for the downward trend in aggregate inequality and within the formal market ceases to exist. While the downward trend still characterises inequality within the informal sector, its slope is much flatter than with unconditional data. This is a novel finding, since the downward trend is preserved when informality is accounted for by using a dichotomic variable instead of estimating the coefficients separately for the formal and informal labour market[4]. This underlines the importance of recognising that the differences between the formal and informal labour markets go well beyond a shift in the level of income. The previous point lies behind the fact that aggregate inequality is larger than the inequality within either labour market, implying that the differences between income levels between markets is larger than the differences found within them.

As was the case with unconditional data, in panel (c) of figure 6 it can be seen that the dynamics of aggregate inequality largely reflect the fluctuations in the top half of the distribution. Moreover, in contrast to the evolution of residual inequality within the formal sector, which has remained roughly constant over the study period, labour income inequality within the informal sector is characterised by a clear downward trend. It is interesting to note that among households whose heads work in the informal sector, towards the end of the period the inequality of income in the top half of the distribution is smaller than in the bottom half of the distribution, where the inequality of equivalised income has not kept the pace of reduction of wage income.

Having captured the main features of the evolution of inequality from the mid 1990s through 2014, 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 (Banco de México, 2007), 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 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

[4]See, for example, figure 3 in a previous version of this paper (Villarreal, 2014).

of nominal frictions, monetary policy can have significant effects on the economy at large (Galí, 2008).

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 on the data. To do so, a standard small 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 differentiated 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 sector, and can be thought of as a forward–looking open economy version of the traditional IS–curve:

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

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 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:

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

where $\overline{y}_t \equiv -\alpha(2-\alpha)(1-\gamma)/\gamma y^*_t$ 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

[5]As discussed by Lubik and Schorfheide (2007), since intermediate–goods 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:

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

where Δe_t is the first difference of the nominal exchange rate, and π_t^* is world inflation.

The policy block of the model is summarised by a Taylor–type rule (Taylor, 1993) 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:

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

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 2001.I to 2014.III. The start date corresponds to the adoption of an inflation targeting regime by the Bank of Mexico (Werner and Schmidt-Hebbel, 2002), which as documented by Chiquiar et al. (2010) induced a significant change in the persistence of inflation in Mexico.

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 filter[6]. 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.

Columns 2 through 5 of table 2 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 variables[7]. The priors for the standard deviation of shocks are non-informative.

The parameter estimation results are listed in the last two columns of 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, results confirm the Bank of Mexico's adherence to the so–called Taylor principle (Woodford, 2003), where optimal monetary policy prioritises the response to deviations of inflation from its target, and the magnitude of the response is more than proportional to the deviation observed.

Considering the importance of imported goods and services in the domestic consumption basket, another result which merits attention is the relatively small value that is estimated for the import share α. This could be related to the relatively high value of the point estimate for the exchange rate parameter in the policy rule, which effectively dampens the pass-through of fluctuations in the price of imported goods.

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

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

Taking into account the very small value assigned to the prior of the Phillips curve slope parameter κ, the estimated coefficient's mean is smaller than the 0,3 found by Lubik and Schorfheide (2007) for Canada[8]. While the difference is not statistically significant, as discussed by Lee (2014), a smaller slope parameter could be evidence of significant financial frictions in the Mexican economy.

The impulse responses of endogenous variables to a unit contractionary monetary policy shock are shown in figure 7. In line with standard results found in the literature for small open economies [9], unanticipated increases in the nominal exchange rate contemporaneously reduce output growth, inflation and cause an appreciation of the nominal exchange rate. While inflation and the exchange rate return to their original levels monotonically over a 10 quarter horizon, which is proportional to the response of the nominal interest rate, growth of GDP quickly rebounds to positive territory and then declines monotonically.

The top panel of figure 8 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.

With respect to 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. 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 with a slight contractionary bias, which could be in response to the bouts of inflation volatility that have been experienced in the aftermath of the international financial crisis, which reflect to some degree the evolution of the exchange rate.

In addition, 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. By the same token, considering the gradual reduction of the objective rate from late 2012, the model interprets the non–response to inflation fluctuations experiences over 2013 as shocks.

Having identified monetary policy shocks, the next section explores their impact on the evolution of household labour income inequality.

[8]Using the Lubik and Schorfheide (2007) benchmark prior, with Mexican data yields a posterior mean of 2,5 for parameter κ.

[9]See for example the results of Cushman and Zha (1997), Galí and Monacelli (2005), and Lubik and Schorfheide (2007).

4 Impact of monetary policy shocks

As discussed by Coibion et al. (2012), from a theoretical perspective the effect of monetary policy shocks on inequality is a priori ambiguous. Their empirical findings, based on U.S. data, stress the importance of the so-called income composition channel, where monetary policy has heterogeneous effects on different sources of income. The relevance of this channel is corroborated by Gornemann et al. (2015) who, using a DSGE with heterogeneous agents and incomplete markets calibrated to U.S. data, find preliminary evidence indicating that as a result of falling output, the demand for labour as well as wages decrease, which cause a fall in income of households which derive most of their income from wages. In contrast, given the presence of sticky prices, the reduction in inflation lead to higher markups, which in turn imply a greater flow of profits towards households who derive most of their income from the ownership of assets. The result is higher inequality of wages and income.

While informative, the models of Coibion et al. (2012) and Gornemann et al. (2015) do not consider the existence of an informal sector. Using a standard small open economy DSGE, augmented to include formal and informal labour markets, Fernández and Meza (2015) find that a reduction in output decreases both demand for labour and wages in the formal sector, leading to a reallocation of resources towards the informal sector as a result of a greater relative elasticity of labour supply in the formal labour market (Campos-Vazquez, 2010). According to Binelli and Attanasio (2010), this should result in greater wage inequality to the extent that wages in the informal sector are not constrained by labour market regulations.

Thus, for the case of Mexico, the evidence suggests that contractionary monetary policy shocks should increase labour income inequality, and the increase in inequality should be more acute among households whose heads are employed in the informal sector.

In order to investigate the impact of the monetary policy shocks on household inequality in Mexico, the natural alternative is to first estimate a vector autoregresive model (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:

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

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 Y_t in equation 5.

4.1 Response of median labour–income

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, are shown in figure 9. The results in the first row to the full sample, whereas those on the second and third rows correspond to the results of the informal and formal labour markets respectively.

For aggregate data, an unanticipated increase in the nominal interest rate causes a contemporaneous decline in hourly wages and income, which turns into an increase of income after the sixth quarter, which is roughly in line to the impact on GDP growth stemming from the DSGE model. However, the results are differentiated across labour markets. In contrast to the response of hourly wages in the informal sector, which closely resemble the response for aggregate data, wages in the formal sector decline over the short run and do not experience a subsequent recovery until after 6 quarters. Despite the relative better performance of wages, equivalised income within the informal sector declines on impact and remains depressed over the course of about a year and a half. For its part, equivalised income declines marginally on impact, and starts recovering after two years. The decline of equivalised income in households whose head works in the informal sector, might reflect negative externalities stemming from greater vulnerability within the informal sector.

The results are consistent with the notion that in the absence of unemployment insurance mechanisms, the informal market tends to act as a buffer in times of crisis, absorbing workers laid off in the formal sector. While this does not seem to have an effect on wages, it depresses equivalised income since as shown in panel (a) of figure 5, median labour–income is systematically lower in the informal sector. Moreover, results suggest the existence of hysteresis in the informal labour market, evidenced by the fact that in contrast to households whose head works in the informal sector, which do not experience a recovery in equivalised income as economic conditions improve, households whose head works in the formal sector do see an increase in their equivalised income as the effect of the shock dissipates over a two year horizon.

4.2 Response of inequality

The response of the Gini coefficient of residual labour income to unanticipated nominal interest rate increases is shown in figure 10. For the case of hourly wages, the response of aggregate data is not statistically significant. However, looking at the results by labour market, this result seems to be the consequence of opposite effects canceling out. While wage inequality decreases marginally in the informal sector, it increases in the formal sector. This response is consistent with recently unemployed workers, which according to Campos-Vazquez (2010) tend to be mostly young and unskilled, migrating from the formal to the informal labour market.

A similar differentiated response is found when looking at the response of inequality of equivalised income, with inequality marginally decreasing among households whose head is employed in the informal sector, and household income inequality increasing in the formal sector, as a result of larger wage dispersion. In contrast to wages, however, in aggregate the effect does not cancel out, and an increase in household labour–income inequality is observed for the full sample.

Figure 11 shows the response of decile ratios by labour market. From the top panel it is clear that the aggregate results just discussed are to be determined by dynamics in the top half of the distribution. In contrast, inequality in the bottom half for the distribution, for both the full sample and the informal labour market, unequivocally decreases in response to contractionary monetary policy shocks. The small magnitude and lack of significance in the formal labour market are probably reflecting the relatively small number of workers from the bottom half of the distribution who are employed in the formal sector.

4.3 Robustness of results

In order to assess the robustness of results, three alternative specifications are used. The first uses the standard deviation of the logarithms of residual labour–income instead of the Gini coefficient to measure inequality. The second classifies all own–account workers as informal independently of whether they have access to social security. Finally, the third alternative uses shocks identified by imposing different priors on the DSGE model discussed in section 3, as well as shocks identified by imposing sign–restrictions on a VAR model.

4.3.1 Alternative measure of inequality

As an alternative measure of inequality, the standard deviation of logarithms of residual labour–income is used. As can be verified in figure 12, the use if this alternative measure does not affect the results. On the contrary, for the case of the fall in inequality among the informal sector as a result of the occurrence of a contractionary monetary policy shock, the use of the standard deviation of logarithms strengthen both the magnitude and significance of the results.

4.3.2 Alternative definition of informality

According to data from the latest available wave of the micro–enterprise survey (INEGI and STPS, 2013), which is a biennial module of the labour survey, only about 45% of self–reported own–account workers have access to social security. Thus the use of the alternative definition can shed some light on the role of having access to social security. As argued by Levy (2008), access to social security is a key dimension of informality, as it reflects the impact of labour market regulation distortions which create incentives for the existence of a large number of small and inefficient firms in the informal sector.

The results are shown in figure 13. Aggregate results are not robust to the definition of informality, a result which seems to originate in the sample corresponding to the formal labour market, where results using the alternative definition are smaller in magnitude, and not statistically significant. While the results for the Gini coefficient and the 9/5 decile ratio are roughly consistent with the benchmark case results, the finding that contractionary monetary policy shocks reduce inequality among workers in the bottom half of the distribution no longer holds.

This discrepancy in results underlines the fact that having access to social security effectively conditions the response of inequality to monetary policy shocks, perhaps because it is correlated with access to specialised markets, such as the one for formal financial services which might help households to insure against the occurrence of idiosyncratic risks.

4.3.3 Alternative shock identification strategies

Two sets of alternative shock identification procedures are used. The first follows the methodology described in section 3, however it uses the alternative prior distributions summarised in table 3. The first alternative corresponds to 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 exchange rate policy parameter is non–informative and does not constrain it to have a positive value.

The main difference between the alternative estimations is explained by differences in the slope of the Phillips curve. 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 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 to the benchmark case. Moreover, as shown in the first two rows of figure 14, the response of inequality to shocks identified under alternative prior distributions are almost indistinguishable from the benchmark specification[10].

The parametrisation of the DSGE model imposes a number of cross–equation restrictions which may not necessarily be supported by the data. With this in mind, the second set of alternative shocks are identified by imposing restrictions on the impulse–response functions of a VAR model. In particular, following the work of Carrillo and Elizondo (2015), the following specification is used:

$$\begin{aligned} Z_t &= \alpha_Z + \sum_{i=1}^{p} D_i Z_{t-i} + \eta_t \qquad (6) \\ Y_t &= \alpha_Y + \sum_{i=1}^{p} A_i Y_{t-i} + \sum_{i=1}^{p} B_i Z_{t-i} + \epsilon_t \end{aligned}$$

where Z_t and Y_t are, respectively, vectors of exogenous and endogenous variables, α_Z and α_Y are vectors of constants, D_i, A_i and B_i are parameter matrices to be estimated, and the

[10]For brevity of exposition only the responses of equivalent income the full sample are shown. The results for hourly wages, as well as those for the informal and formal labour market sub–samples are available from the author upon request.

vector of errors $[\eta_t \epsilon_t]'$ is assumed to have mean zero, no serial correlation and covariance matrix equal to:

$$\Sigma = \begin{bmatrix} \sigma_\eta & \sigma_{\eta\epsilon} \\ \sigma_{\epsilon\eta} & \sigma_\epsilon \end{bmatrix}$$

The variables used are:

$$Y_t = \begin{array}{c} \text{Output gap} \\ \text{Inflation gap} \\ \text{Producer inflation gap} \\ \text{Real exch. rate depreciation} \\ \text{Real interest rate} \\ \text{Real money growth} \end{array} \begin{bmatrix} y_t - \overline{y}_t \\ \pi_t - \overline{\pi}_t \\ \pi_t^p - \overline{\pi}_t \\ \Delta q_t \\ i_t - \overline{\pi}_t \\ \Delta m_t - \overline{\pi}_t \end{bmatrix}, \qquad Z_t = \begin{array}{c} \text{US Output gap} \\ \text{US Inflation} \\ \text{Oil price inflation} \end{array} \begin{bmatrix} y_t^* - \overline{y^*}_t \\ \pi_t^* \\ \Delta wti_t \end{bmatrix}$$

where the domestic output gap $y_t - \overline{y}_t$ is obtained by computing the log difference of the level of the series y_t with respect to its Hodrick–Prescott (HP) filtered trend $\overline{y}_t$[11]; core consumer price inflation π_t, producer price inflation π_t^p, nominal interest rates i_t[12] and nominal money growth Δm_t[13] are detrended using the HP filtered trend for core consumer price inflation[14][15]. All domestic data, are from the Mexican statistical institute, except for the real exchange rate index which is from the Banco de México. US output and price data comes from the Bureau of Economic Analysis, and oil price data comes from the Energy Information Agency[16].

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[17].

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[18].

In particular, two cases are considered. The first imposes (long–run) block exogeneity of the exogenous variables and sign restrictions on the responses to aggregate supply, aggregate demand and monetary policy shocks as summarised in the following matrix:

$$\begin{array}{r} \\ \text{Aggregate supply} \\ \text{Aggregate demand} \\ \text{Monetary policy} \end{array} \begin{array}{c} \begin{array}{cccccc} y_t - \overline{y}_t & \pi_t - \overline{\pi}_t & \pi_t^p - \overline{\pi}_t & \Delta q_t & i_t - \overline{\pi}_t & \Delta m_t - \overline{\pi}_t \end{array} \\ \begin{bmatrix} +1 & -1 & -1 & +1 & \times & \times \\ +1 & +1 & +1 & -1 & +1 & \times \\ -1 & -1 & \times & -1 & +1 & -1 \end{bmatrix} \end{array}$$

[11]Following standard practice, the Hodrick–Prescott filter is used with a smoothing parameter equal to 1,400.

[12]In contrast to the estimation of the DSGE model, where the overnight interest rate was used, the rate for 28–day Mexican treasuries (CETES) is used instead.

[13]The M2 monetary aggregate is used.

[14]All changes are quarter on quarter changes.

[15]Price inflation and changes on the level of GDP and the M2 monetary aggregate are computed on the basis of seasonally adjusted data.

[16]To price of West Texas Intermediate oil, which is the relevant commodity price for the case of Mexico, is used.

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

[18]A brief description of the methodology to identify shocks by imposing restrictions on the VAR impulse responses is available in appendix A

where $\times$'s imply no restriction. The second case imposes (long–run) block exogeneity plus the sign–restrictions corresponding to aggregate supply and demand only.

As in section 4, the responses of inequality to the identified monetary policy shocks are estimated by means of local linear projections. The responses of household's equivalent income for the full sample are shown in the bottom two rows of figure 14. Although the results stemming from the VAR impulse response identification procedure are not as clear–cut as those stemming from the benchmark specification, they are broadly in line with the results discussed in section 4. That is, an unanticipated increase in nominal interest rates causes an increase in inequality of household's equivalent income. Moreover, the impact is differentiated across the distribution as the rise in inequality concentrates on the top half of the distribution; in contrast, among households in the bottom half inequality of labour–income actually declines.

In summary, the response of households' labour–income inequality to unanticipated changes in interest rates are robust to both the particular measure of inequality that is used, as well as to the procedure used to identify structural shocks. The results are however not robust to the definition of informality, underlining the relevance of labour market informality, which beyond excluding households from social security is correlated with limited access to certain key markets, such as the one for financial services.

5 Conclusions

The macroeconomics workhorse model for the analysis and formulation of monetary policy is the New–Keynesian Dynamic Stochastic General Equilibrium Model, which relies on the assumption that the demand–side of the economy can be modeled as a representative household, thus precluding the analysis of the distributive impact of monetary policy. The validity of the assumption of a representative household hinges on the existence of complete markets where households can insure themselves against the occurrence of idiosyncratic risks such as illness or unemployment. The empirical evidence indicates that market incompleteness is significant, thus casting doubt on the validity of the use of a representative household as a modeling device.

Once households are allowed to be heterogenous, from a theoretical perspective the nature of the impact of monetary policy shocks on household inequality, if any, is ambiguous since it depends on the sign and magnitude of alternative transmission channels. Using data for the United States, Coibion et al. (2012) and Gornemann et al. (2015) highlight the importance of the income composition channel, and find evidence that household inequality increases as a result of contractionary monetary policy shocks.

This paper follows the approach of Coibion et al. (2012) to empirically investigate the impact of monetary policy on the distribution of household labour–income for the case of Mexico. Recognising the relevance of the informal labour market, the paper separately identifies the impact on households whose head is formally or informally employed, as proxied by having access to social security. In line with the results found for the United States, this paper finds evidence that unanticipated increases in the nominal interest rate cause an increase in the inequality of household's labour income. Moreover, it finds that the effect is heterogeneous across households depending on whether their head is employed formally or informally. While over the medium term, households whose head is formally employed see a increase in inequality of labour–income, this occurs in the context of rising real income levels. In contrast, while the distribution of labour–income of households whose head is informally employed becomes less unequal, this comes about as real incomes fall.

A possible explanation for the heterogeneous responses found among households stems from the dynamics of employment across the cycle. In the absence of unemployment benefits and very limited access to and use of formal financial services, which could enable households to insure themselves against risks, the informal labour market acts as a buffer across the business cycle, with young and unskilled workers the most likely to migrate from the formal to the informal labour market as a result of shocks. Since this group has benefitted the most from the fall of the skill premium in the labour market, which has been the main force driving down the inequality of labour income over the last two decades, their migration from the formal to the informal labour market can, in principle, explain the (rise) fall of inequality of households whose head is employed (in)formally.

In terms of public policy, the results suggests that the impact of monetary policy on households' labour-income inequality could be attenuated by three sets of policies. The first concerns the availability of unemployment insurance, which could afford workers to search for employment within the formal sector, instead of migrating to the informal sector as a response to unemployment. The second set is related to policies aimed at reducing the size of both informal employment, which to a large extent reflect distortions introduced by labour regulations Levy (2008), as well as the informal sector, which mainly reflect the incentives created by the design and implementation of tax policies. The third set of policies are related to a financial inclusion strategy which enhances the set of tools which households can use to insure themselves against risk.

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6 Tables and figures

6.1 Tables

Table 1 – Mexico 2012: Household Income by Decile
(Current pesos)

	Total	Decile I	II	III	IV	V	VI	VII	VIII	IX	X
	(Proportion of total)										
Total current income	100.00	100.00	100.00	100.00	100.00	100.00	100.00	100.00	100.00	100.00	100.00
Labour Income	66.43	38.49	50.03	57.40	62.60	65.85	66.36	68.99	73.11	69.75	66.32
Salaried workers	53.21	20.44	32.34	43.13	48.40	50.24	53.40	55.92	58.46	58.67	53.74
Own-account workers	10.46	13.58	13.33	10.41	10.75	11.69	9.14	10.35	11.88	8.57	10.54
Other labour income	2.76	4.47	4.36	3.85	3.44	3.93	3.82	2.72	2.77	2.51	2.03
Capital income	4.11	0.93	0.70	0.69	1.02	1.49	1.28	1.59	1.58	2.98	8.39
Transfers	17.25	36.40	31.14	26.10	21.62	18.32	18.92	16.18	13.17	16.16	15.25
Imputed rent	12.09	23.71	17.94	15.62	14.71	14.28	13.41	13.11	12.13	10.97	9.91
Other current income	0.12	0.47	0.20	0.19	0.05	0.06	0.04	0.13	0.02	0.15	0.14

Source: Author based on INEGI (2013)

Table 2 – DSGE model parameters: Benchmark specification

	Prior distribution				Posterior estimation	
Parameter	Density	Domain	μ	σ	Mean	90% Interval
ψ_π	Gamma	$\mathbb{R}^+$	1.5000	0.7500	2.3730	[1.5039,3.2192]
ψ_y	Gamma	$\mathbb{R}^+$	0.7500	0.3300	0.6253	[0.1904,1.0463]
ψ_e	Gamma	$\mathbb{R}^+$	0.2500	0.1300	0.2663	[0.0864,0.4385]
ρ_R	Beta	[0,1)	0.8000	0.1000	0.8341	[0.7860,0.8839]
α	Beta	[0,1)	0.5000	0.2000	0.0090	[0.0005,0.0179]
r_{ss}	Gamma	$\mathbb{R}^+$	2.0000	1.0000	2.0119	[0.4556,3.4739]
κ	Gamma	$\mathbb{R}^+$	0.0200	0.0060	0.0236	[0.0136,0.0336]
γ	Beta	[0,1)	0.5000	0.2500	0.0244	[0.0118,0.0371]
ρ_q	Beta	[0,1)	0.1500	0.0800	0.1627	[0.0463,0.2751]
ρ_A	Beta	[0,1)	0.3500	0.1700	0.2327	[0.1906,0.2803]
ρ_{y^*}	Beta	[0,1)	0.8000	0.1000	0.9048	[0.8561,0.9509]
ρ_{π^*}	Beta	[0,1)	0.5000	0.2500	0.2058	[0.0295,0.3660]
σ_R	Inv. Gamma	$\mathbb{R}^+$	1.0000	4.0000	0.9724	[0.8026,1.1349]
σ_q	Inv. Gamma	$\mathbb{R}^+$	1.0000	4.0000	3.6647	[3.1153,4.2348]
σ_A	Inv. Gamma	$\mathbb{R}^+$	1.0000	4.0000	1.1732	[0.9608,1.3807]
σ_{y^*}	Inv. Gamma	$\mathbb{R}^+$	1.0000	4.0000	1.1444	[0.2399,2.7804]
σ_{π^*}	Inv. Gamma	$\mathbb{R}^+$	1.0000	4.0000	2.9263	[2.4761,3.3725]

Notes: 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. The posterior distribution was estimated using Dynare version 4.4.3 (Adjemian et al., 2011) through 100,000 draws obtained using the Metropolis–Hastings algorithm, dropping 20% of the resulting draws. 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.

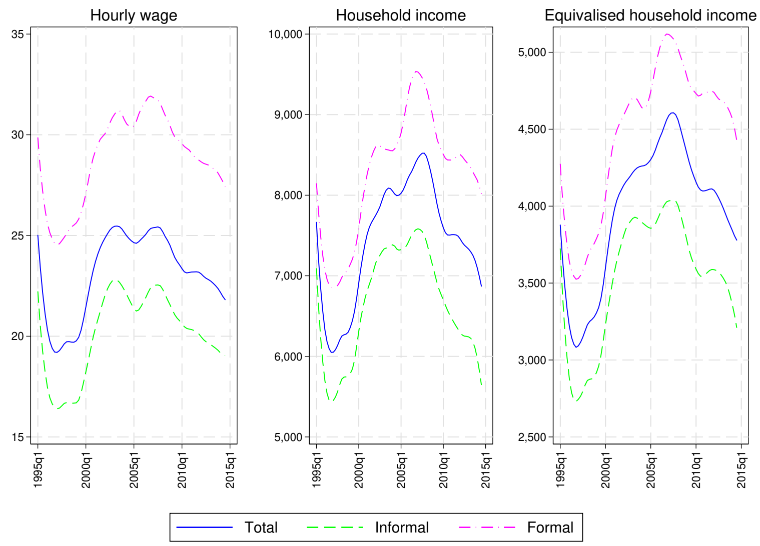
Table 3 – Alternative priors for DSGE model

	Prior Distribution: Alternative 1				Prior Distribution: Alternative 2			
Parameter	Density	Domain	μ	σ	Density	Domain	μ	σ
ψ_π	Gamma	$\mathbb{R}^+$	1.5000	0.5000	Normal	$\mathbb{R}^+$	1.5000	0.2500
ψ_y	Gamma	$\mathbb{R}^+$	0.2500	0.1300	Normal	$\mathbb{R}^+$	0.7500	0.3300
ψ_e	Gamma	$\mathbb{R}^+$	0.2500	0.1300	Gamma	[-2,2]	0.0000	1.1500
ρ_R	Beta	[0,1)	0.5000	0.2000	Uniform	[0,1)	0.8000	0.1000
α	Beta	[0,1)	0.2000	0.0500	Beta	[0,1)	0.5000	0.2000
r_{ss}	Gamma	$\mathbb{R}^+$	2.5000	1.0000	Gamma	$\mathbb{R}^+$	2.0000	1.0000
κ	Gamma	$\mathbb{R}^+$	0.5000	0.2500	Gamma	$\mathbb{R}^+$	0.0200	0.0060
γ	Beta	[0,1)	0.5000	0.2500	Beta	[0,1)	0.5000	0.2500
ρ_q	Beta	[0,1)	0.4000	0.2000	Beta	[0,1)	0.1500	0.0800
ρ_A	Beta	[0,1)	0.2000	0.0500	Beta	[0,1)	0.3500	0.1700
ρ_{y^*}	Beta	[0,1)	0.9000	0.0500	Beta	[0,1)	0.8000	0.1000
ρ_{π^*}	Beta	[0,1)	0.8000	0.1000	Beta	[0,1)	0.5000	0.2500
σ_R	Inv. Gamma	$\mathbb{R}^+$	0.5000	4.0000	Inv. Gamma	$\mathbb{R}^+$	1.0000	4.0000
σ_q	Inv. Gamma	$\mathbb{R}^+$	1.5000	4.0000	Inv. Gamma	$\mathbb{R}^+$	1.0000	4.0000
σ_A	Inv. Gamma	$\mathbb{R}^+$	1.0000	4.0000	Inv. Gamma	$\mathbb{R}^+$	1.0000	4.0000
σ_{y^*}	Inv. Gamma	$\mathbb{R}^+$	1.5000	4.0000	Inv. Gamma	$\mathbb{R}^+$	1.0000	4.0000
σ_{π^*}	Inv. Gamma	$\mathbb{R}^+$	0.5500	4.0000	Inv. Gamma	$\mathbb{R}^+$	1.0000	4.0000

Notes: 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.

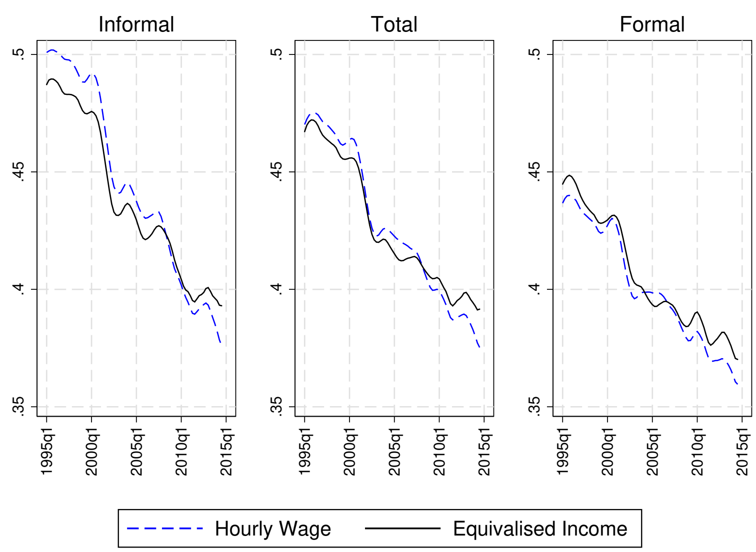
6.2 Figures

Figure 1 – Mexico 1995 – 2014: Evolution of labour-income for the median household (2010 Pesos)



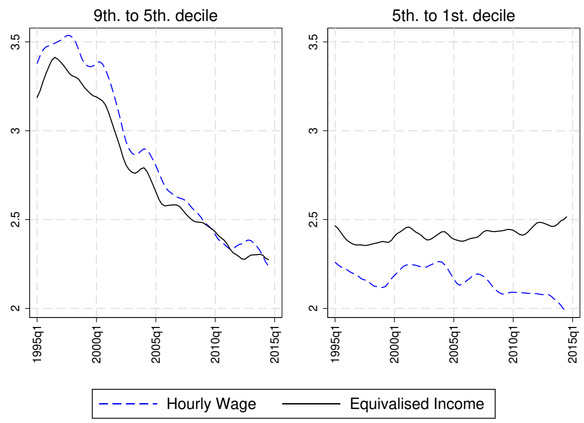
Note: In order to improve the readability of the figure, the series shown were smoothed using a non–parametric locally weighted regression with bandwidth equal to 0.15.

Figure 2 – Mexico 1995 – 2014: Evolution of Gini coefficient for labour–income

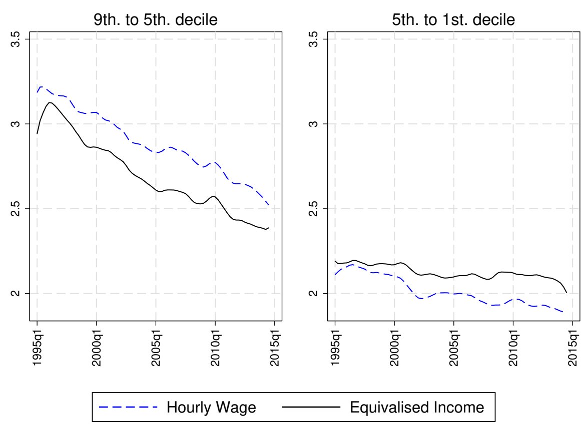


Note: In order to improve the readability of the figure, the series shown were smoothed using a non–parametric locally weighted regression with bandwidth equal to 0.15.

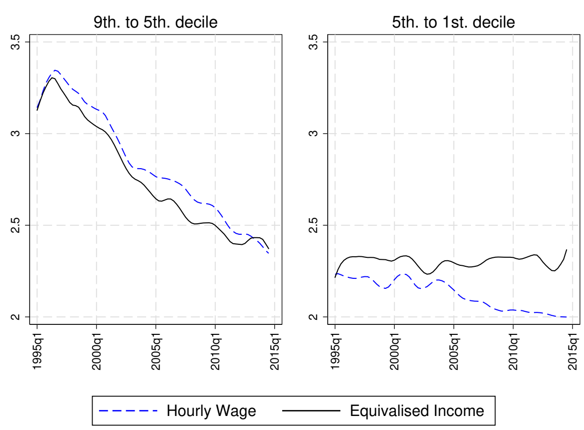
Figure 3 – Mexico 1995 – 2012: Evolution of labour–income decile ratios



(a) Informal sector



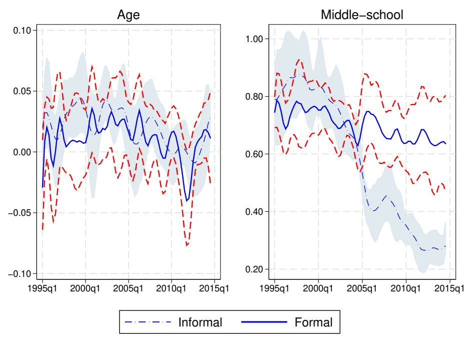
(b) Formal sector



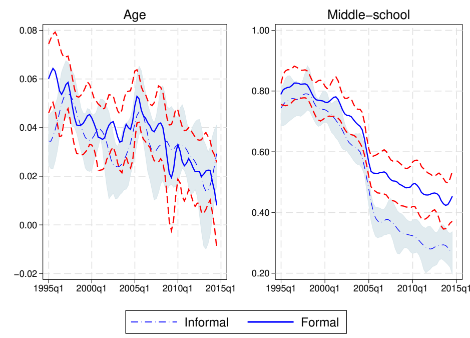
(c) Full sample

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.

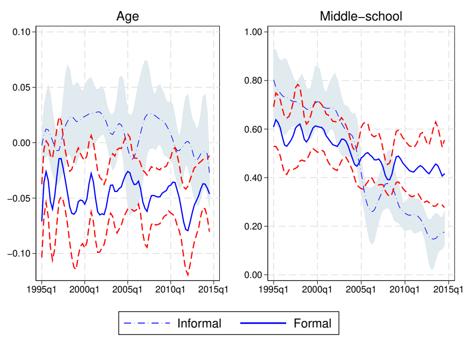
Figure 4 – Regression coefficients



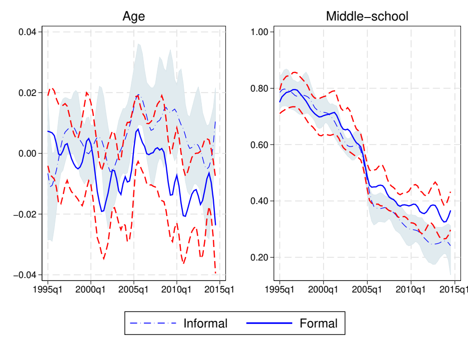
(a) Women: Hourly wages



(b) Men: Hourly wages



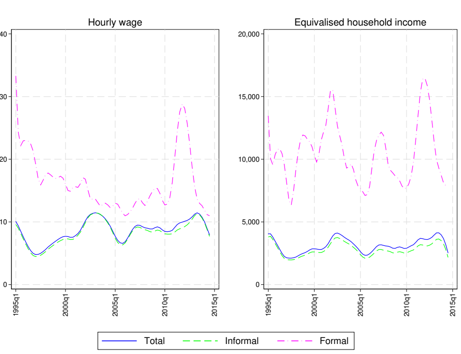
(c) Women: Equivalised income



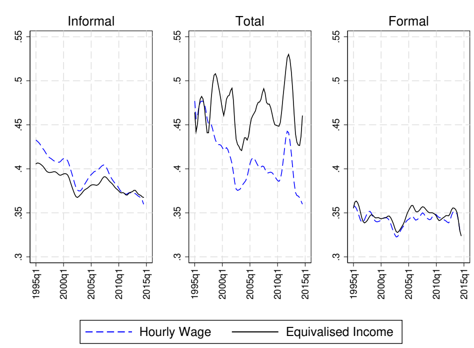
(d) Men: Equivalised income

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 5 – Mexico 1995 – 2014: Evolution of residual labour-income and inequality



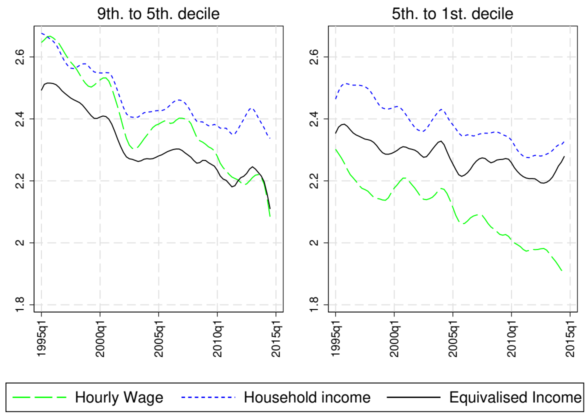
(a) Median household labour–income



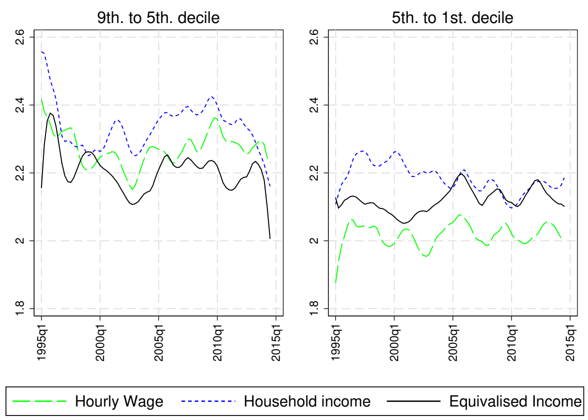
(b) Gini coefficient

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.

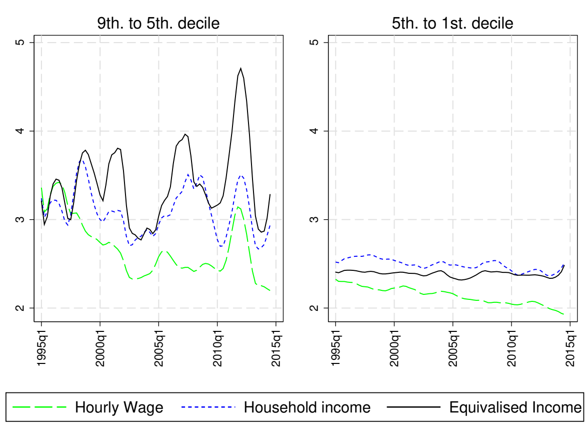
Figure 6 – Mexico 1995 – 2012: Evolution of residual labour–income decile ratios



(a) Informal sector



(b) Formal sector



(c) Full sample

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 7 – Impulse Response Functions to a Monetary Policy Shock (Benchmark DSGE model)

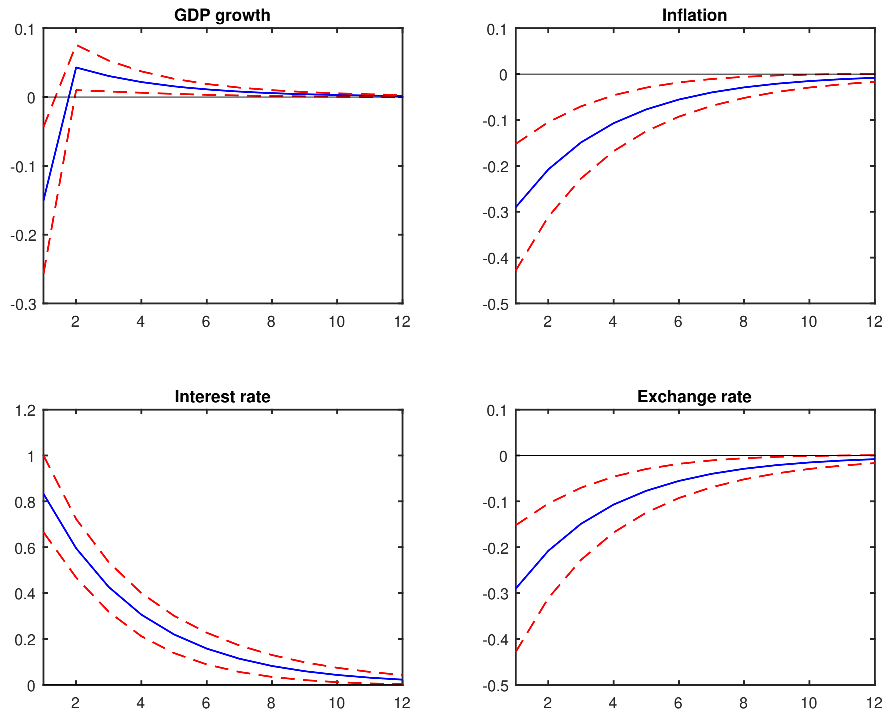


Figure 8 – Smoothed Monetary Policy Shocks (Benchmark DSGE model)

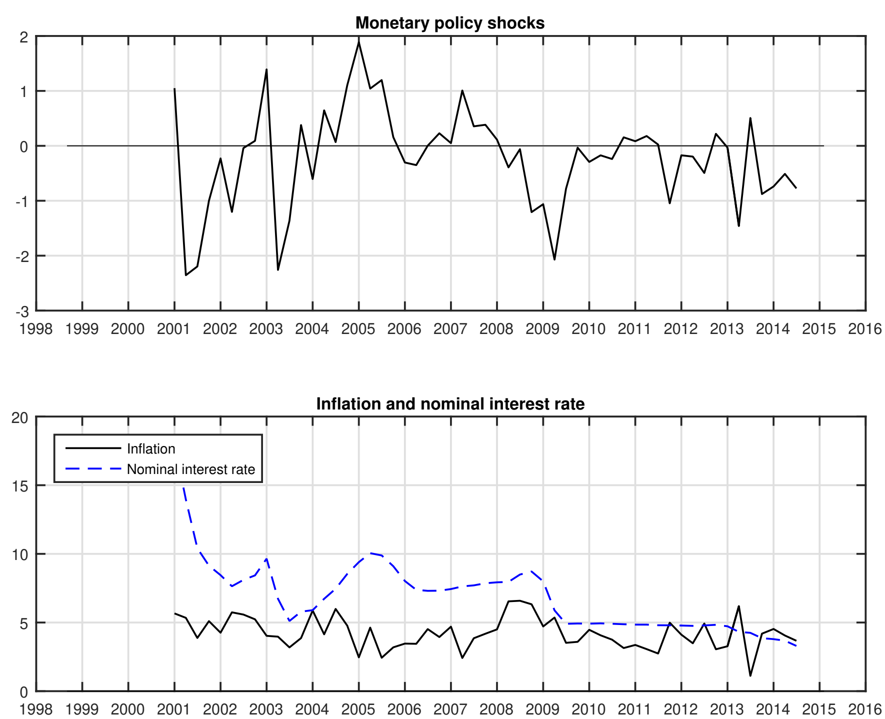
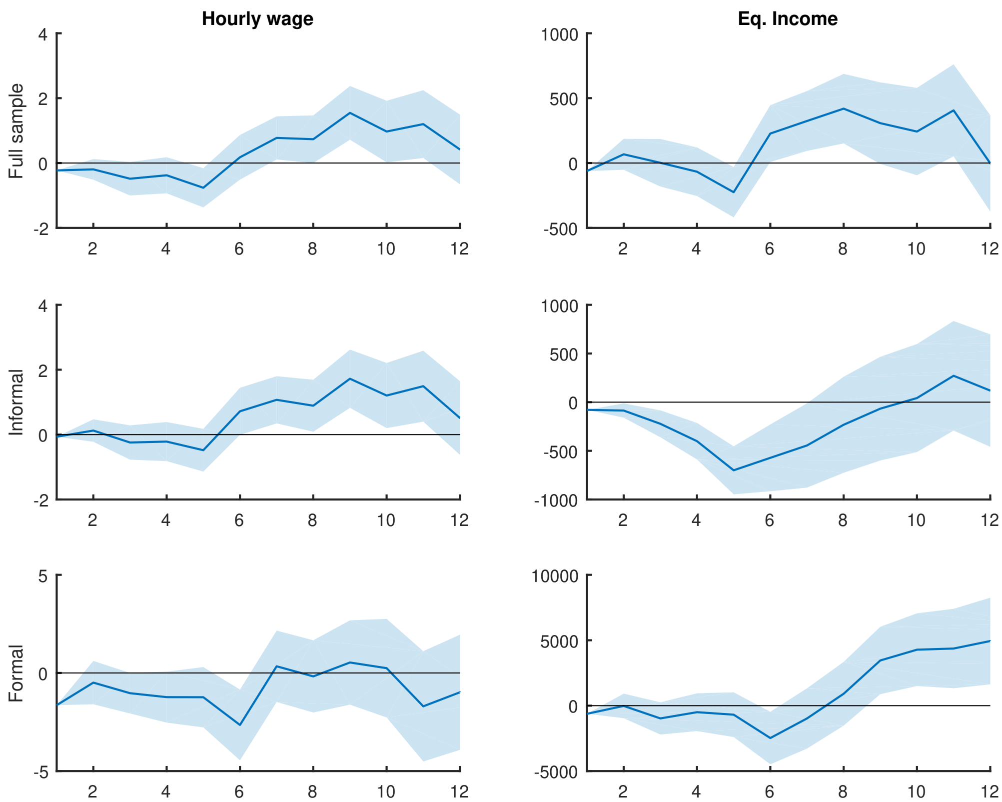
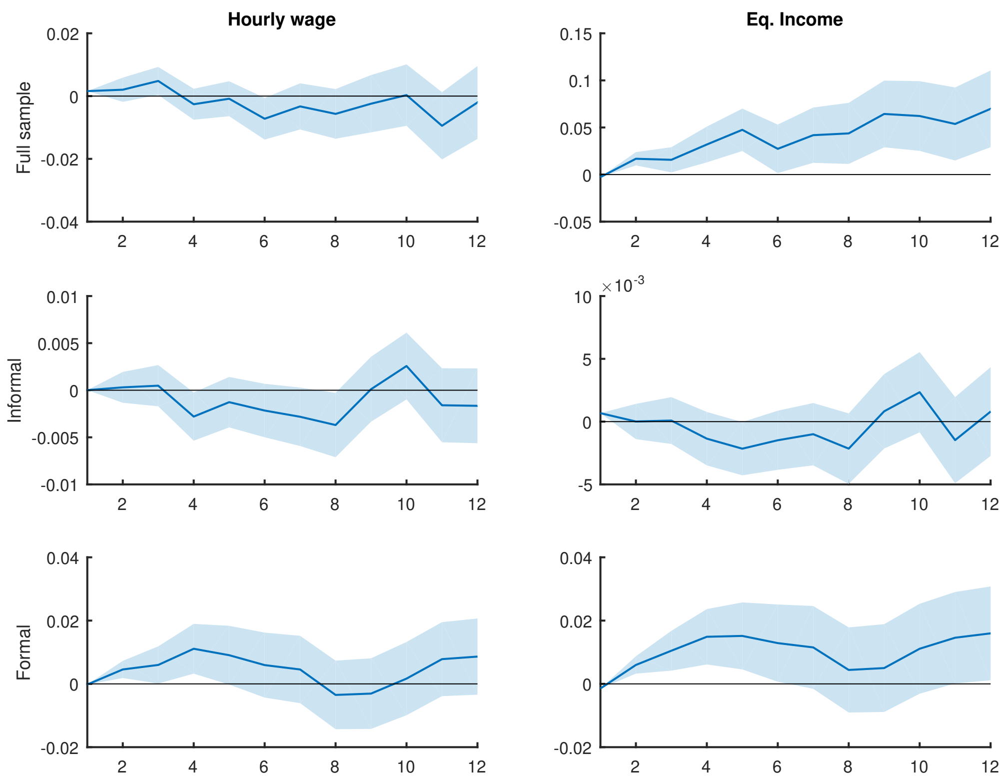


Figure 9 – Impulse Response Functions of Median Income to a Monetary Policy Shock (Benchmark DSGE model – by labour market))



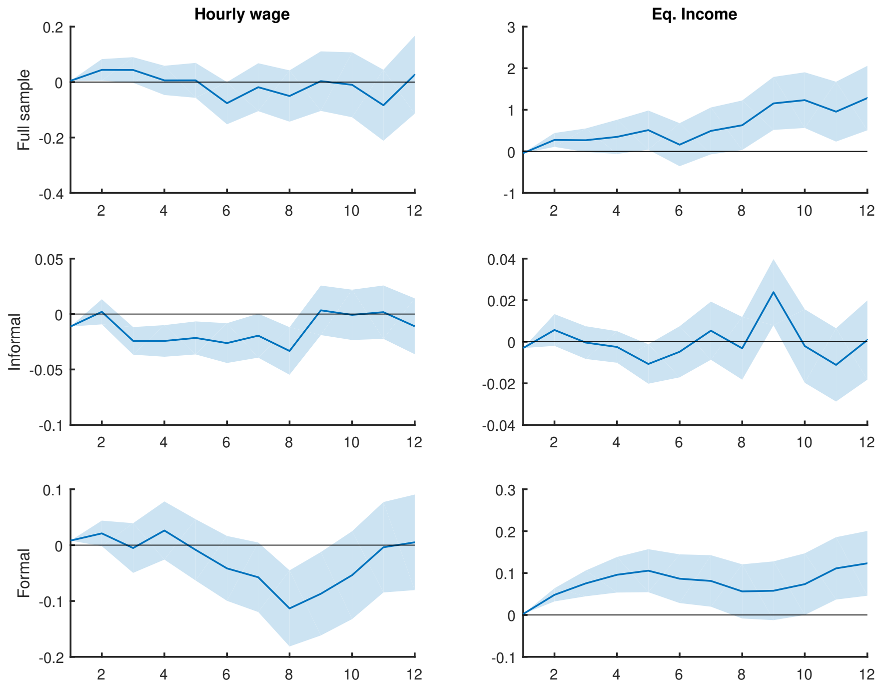
Note: The graphs show the cumulative response to monetary policy shocks. The responses were estimated using local projections over a 12-quarter period. The solid line is the response's point estimate, and the shaded area its confidence interval.

Figure 10 – Impulse Response Functions of Gini coefficient to a Monetary Policy Shock (Benchmark DSGE model – by labour market)

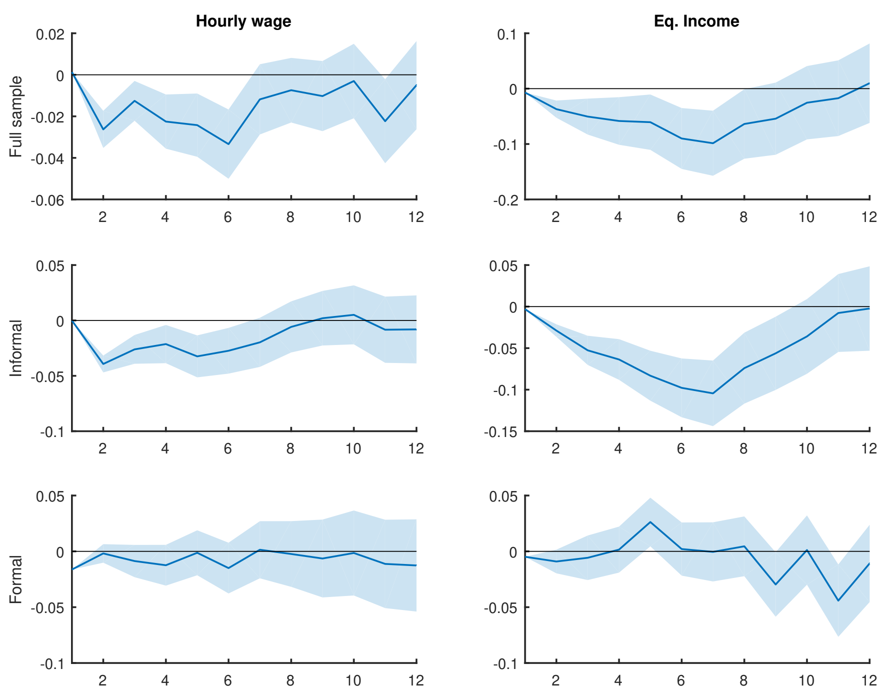


Note: The graphs show the cumulative response to monetary policy shocks. The responses were estimated using local projections over a 12-quarter period. The solid line is the response's point estimate, and the shaded area its confidence interval.

Figure 11 – Impulse response functions of labour–income decile ratios to a monetary policy shock
(Benchmark DSGE model – by labour market)



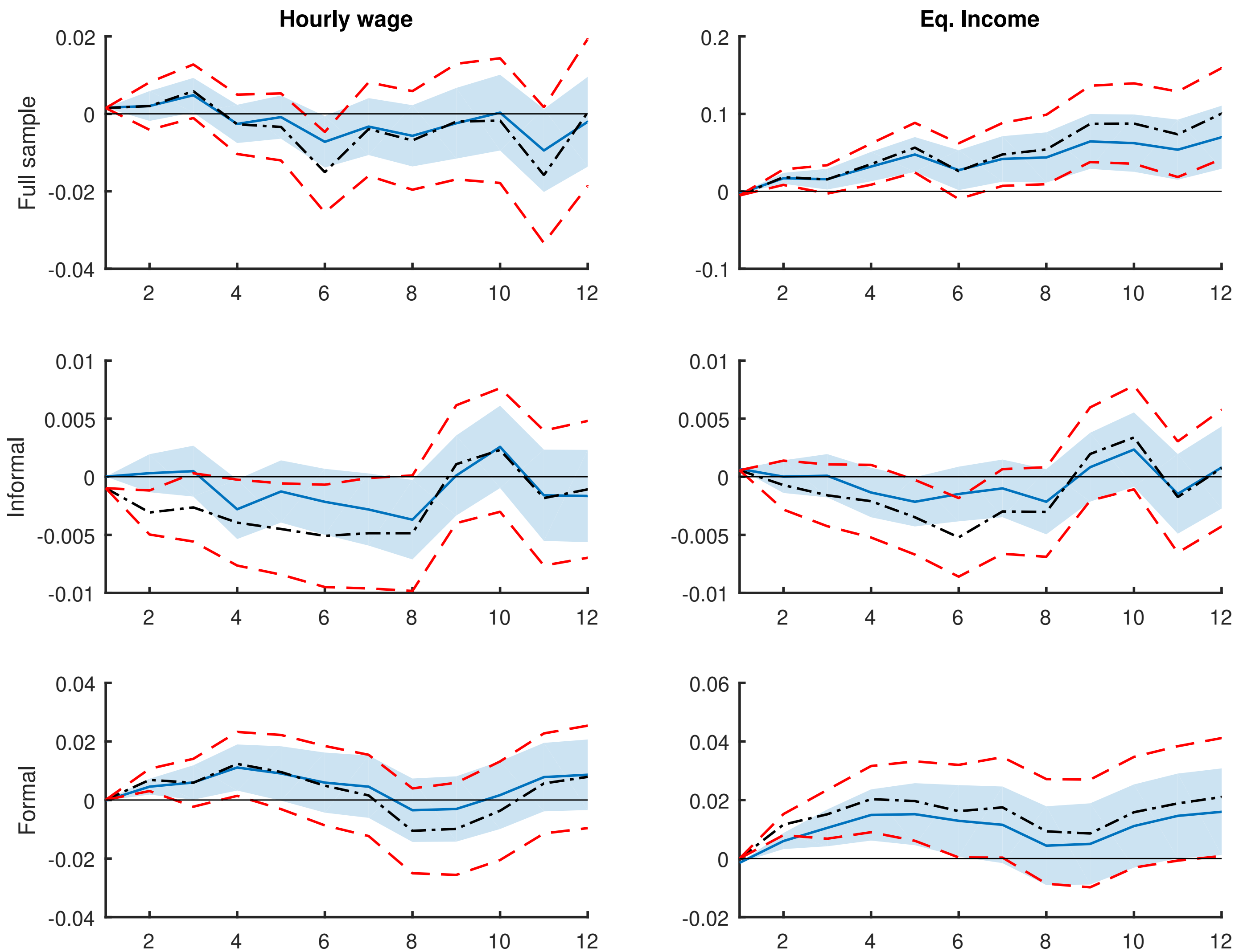
(a) Ratio of 9th. to 5th. income decile



(b) Ratio of 5th. to 1st. income decile

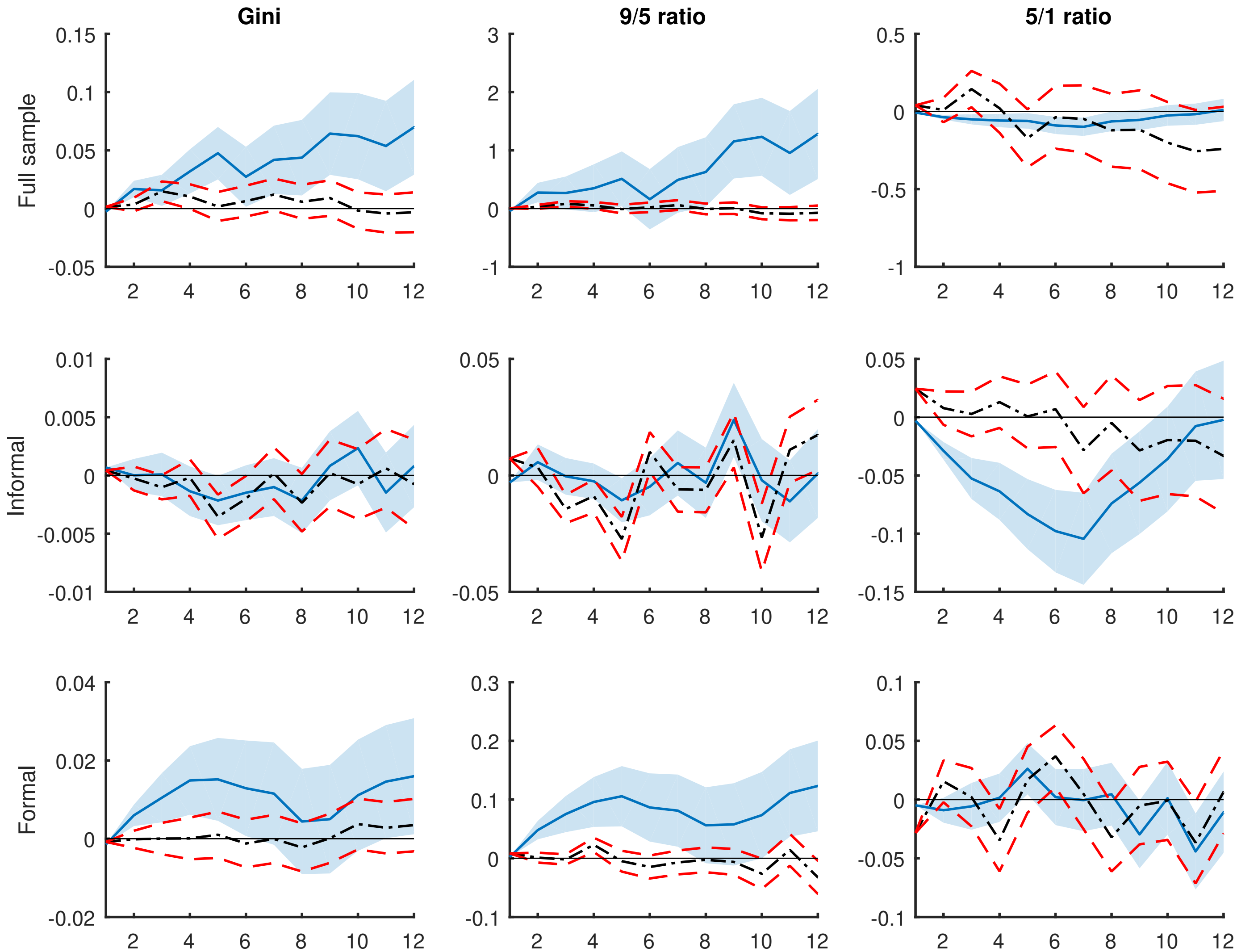
Note: The graphs show the cumulative response to monetary policy shocks. The responses were estimated using local projections over a 12-quarter period. The solid line is the response's point estimate, and the shaded area its confidence interval.

Figure 12 – Robustness of results with respect to the measure of inequality



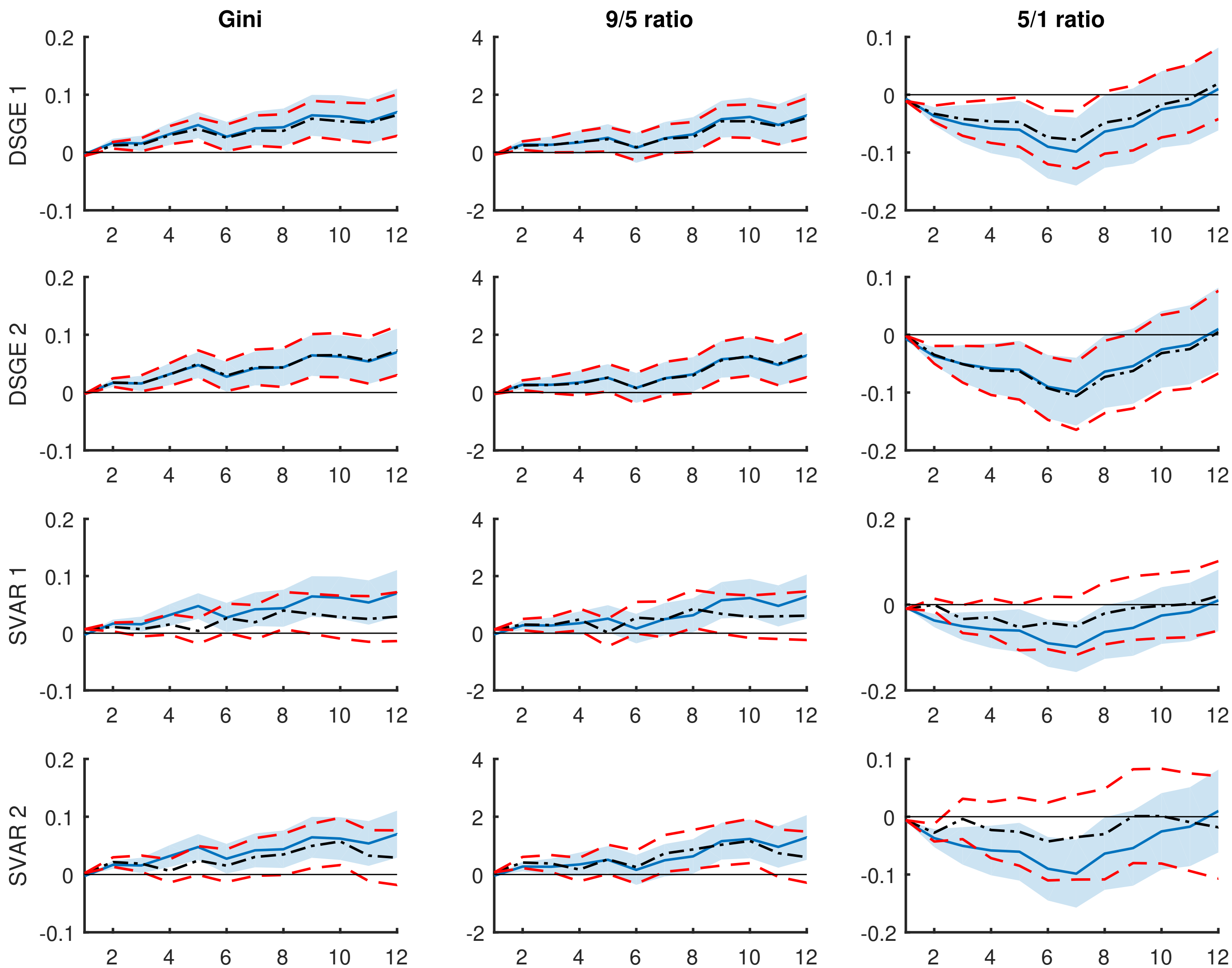
Note: The graphs show the cumulative response to monetary policy shocks. The responses were estimated using local projections over a 12-quarter period. The solid line and shaded area are, respectively, the response and confidence interval under the benchmark specification, while the dot and dashed and dashed lines are, respectively, the response and confidence interval under the alternative specification.

Figure 13 – Robustness of results with respect to the definition of informality



Note: The graphs show the cumulative response to monetary policy shocks. The responses were estimated using local projections over a 12-quarter period. The solid line and shaded area are, respectively, the response and confidence interval under the benchmark specification, while the dot and dashed and dashed lines are, respectively, the response and confidence interval under the alternative specification.

Figure 14 – Robustness of results with respect to alternative shock identification schemes.



Notes: The graphs show the cumulative response to monetary policy shocks. The responses were estimated using local projections over a 12-quarter period. The solid line and shaded area are, respectively, the response and confidence interval under the benchmark specification, while the dot and dashed and dashed lines are, respectively, the response and confidence interval under the alternative specification. DSGE 1 and DSGE 2 denote the response of inequality measures with respect to shocks identified under the alternative prior distributions summarised in table 3; whereas SVAR 1, and SVAR 2 correspond to responses to shocks identified from the VAR restriction schemes discussed in section 4.3.3.

A Shock identification using restrictions on VAR impulse response functions

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:

$$\mathbf{Y}_{t+1} = B(L)\mathbf{Y}_t + u_{t+1} \tag{7}$$

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 impulse responses of system (7) by estimating its coefficients subject to constrains of particular elements of the response matrices A_0 and A_+.

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, 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.