Pontifícia Universidade Católica
do Rio de Janeiro

Laura Candido de Souza

Essays in Macroeconomics

Tese de Doutorado

Thesis presented to the Postgraduate Program in Economics of the Departamento de Economia, PUC–Rio as partial fulfillment of the requirements for the degree of Doutor em Economia

Advisor : Prof. Carlos Viana de Carvalho
Co–Advisor: Prof. Eduardo Zilberman

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Abstract

Souza, Laura Candido; Viana de Carvalho, Carlos (Advisor); Zilberman, Eduardo (Co-advisor). **Essays in Macroeconomics**. Rio de Janeiro, 2015. 102p. PhD Thesis — Departamento de Economia, Pontifícia Universidade Católica do Rio de Janeiro.

This dissertation is composed of three articles in macroeconomics. The first article explores the macroeconomics effects of the credit deepening processes observed in Peru and Mexico using a standard New Keynesian dynamic general equilibrium model with financial frictions. From the perspective of the model, the effects on consumption, GDP and investment are small. Hence, our results suggest only a modest contribution of credit expansion to the above-trend growth experienced by Peruvian and Mexican economies during our sample period. In the second article, we documented that the association between consumption growth and credit expansion is stronger in countries with higher income inequality. We use an incomplete-markets model with heterogeneous households, idiosyncratic risk and borrowing constraints to corroborate this empirical finding. A loosening of credit constraints mitigates precautionary motives, inducing households to reduce savings along the transition path to the new steady-state. Therefore, consumption grows more rapidly in the short-run. This consumption boom is amplified in economies with more constrained households. We consider two sources of income inequality in our model: the variance of the idiosyncratic risk and the households' fixed level of human capital. They have different implications for the extent to which households are credit constrained in equilibrium. We show that when the source of income inequality comes from households' lowest fixed level of human capital, our model can rationalize the empirical evidence. In the other cases, the opposite occurs. The third article tests the effects of a major program of interventions in foreign exchange markets announced by the Central Bank of Brazil to fight excess volatility and exchange rate overshooting. We use a synthetic control approach to determine whether or not the intervention program was successful. Our results suggest that the first foreign exchange intervention program mitigated the depreciation of the real against the dollar. A second announcement made later in the year that the program was going to continue on a smaller basis had a smaller effect, which was not significant. This result is corroborated by a standard event study methodology. We also document that both program did not have an impact on the volatility of the exchange rate.

Keywords

credit deepening; financial frictions; crédito por convénio; crédito de nómina; payroll lending; income inequality; credit expansion; consumption growth; incomplete markets; FX interventions; synthetic control.

Resumo

Souza, Laura Candido; Viana de Carvalho, Carlos (Orientador); Zilberman, Eduardo (Co-orientador). **Ensaio em Macroeconomia**. Rio de Janeiro, 2015. 102p. Tese de Doutorado — Departamento de Economia, Pontifícia Universidade Católica do Rio de Janeiro.

Esta tese é composta por três artigos relacionados à macroeconomia. O primeiro artigo analisa os efeitos macroeconômicos dos processos de aprofundamento de crédito observados no Peru e no México através de um modelo padrão Novo Keynesiano dinâmico de equilíbrio geral com fricções financeiras. Do ponto de vista do modelo, os efeitos sobre o consumo, o PIB e o investimento são pequenos. Assim, nossos resultados sugerem apenas uma contribuição modesta da expansão do crédito para o crescimento acima do potencial das economias peruana e mexicana durante o período considerado. No segundo artigo, documentamos que a associação entre o crescimento do consumo e expansão do crédito é maior para países com maior desigualdade de renda. Nós usamos um modelo de mercados incompletos com agentes heterogêneos, risco idiossincrático e restrições ao crédito para verificar em que medida este arcabouço teórico é capaz de racionalizar a evidencia empírica. Em nosso modelo, consideramos duas fontes de desigualdade de renda: a variância do risco idiossincrático e o nível fixo de capital humano dos agentes. Mostramos que, quando a fonte de desigualdade de renda vem da menor nível fixo das famílias do capital humano, o nosso modelo pode racionalizar a evidência empírica. Nos outros casos, o resultado oposto ocorre. O terceiro artigo testa os efeitos de um grande programa de intervenções no mercado cambial anunciado pelo Banco Central do Brasil afim de combater o excesso de volatilidade e overshooting da taxa de câmbio. Nós usamos uma abordagem de controle sintético para determinar se o programa de intervenção foi bem sucedido ou não. Nossos resultados sugerem que o primeiro programa de intervenção cambial mitigou a depreciação do real frente ao dólar. Todavia, um segundo anúncio feito no final do ano que o programa ia continuar com uma intensidade menor teve um efeito menor e não significativo. Esse resultado é corroborado por uma metodologia de estudo de evento padrão. Nós também documentamos que o programa e a continuação do mesmo não tiveram impacto sobre a volatilidade da taxa de câmbio.

Palavras–chave

aprofundamento de crédito; fricções financeiras; crédito por convénio; crédito de nómina; crédito consignado; desigualdade de renda; expansão de crédito; crescimento do consumo; mercados incompletos; intervenções cambiais; controle sintético.

Contents

[1]This is a joint work with Nilda Pasca.
[2]This is a joint work with Nilda Pasca.

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1
Macroeconomic Effects of Credit Deepening in Latin America: Peru and Mexico[1]

1.1
Introduction

In the last decade, Latin America (LA) countries experienced significant growth in different types of credits, which remained sustained despite of the fact that these economies were affected by the last financial crisis.[2] As shown in Figure 1.1, the domestic credit as a percentage of GDP in many countries of LA has shown an increasing trend, specially, since 2005.

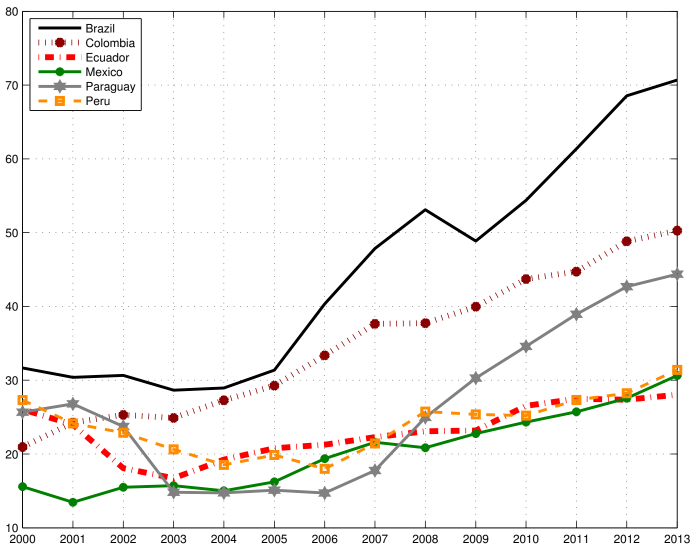


Figure 1.1: Domestic credit to private sector over GDP. Domestic credit to private sector refers to financial resources provided to the private sector, such as through loans, purchases of nonequity securities, and trade credits and other accounts receivable that establish a claim for repayment. For some countries these claims include credit to public enterprises. Source: World Bank, available at data.worldbank.org.

[1]This is a joint work with Nilda Pasca.

[2]After the crisis, the monetary authorities have adopted various measures to mitigate the effect of the crisis on their economies.

During this period, there were also several structural changes in LA, which allowed these economies to maintain macroeconomic stability and higher economic dynamism. For this reason, the expansion of credit is often cited as one of the factors that contributed to their growth. In this paper, we follow Carvalho et al. (2014) to quantify the macroeconomic effects of the credit deepening in LA, in particular, in Peru and Mexico.

To achieve our goal, we calibrate a New Keynesian dynamic general equilibrium model with financial frictions for Mexican and Peruvian economies. We consider an economy populated by patient households that engage with impatient households and entrepreneurs in the credit market through a stylized banking sector. Both impatient households and entrepreneurs face credit constraints.

Concerning the impatient households, the amount they can borrow is constrained by their next period's labor income and by the value of their next period's stock of housing. We do so motivated by the fact that in many countries of LA, as in Peru and Mexico, a sizable part of the increase of credit directed to households was due to a type of credit that is not associated with any type of collateral, known as "*payroll loan.*" In particular, this type of credit is part of the personal credit and the borrowers are charged directly from their payroll. Note that this type of credit has different names in different countries of LA. For example, in Peru, this type of credit is called "*crédito por convénio,*" in Brazil, "*crédito consignado*," in Colombia, "*libranza,*" and in Mexico, "*crédito de nómina.*"

We also consider that the amount that impatient households can borrow is tied to the value of their next period's stock of housing. Our motivation is due to the intense growth of housing credit experienced by several countries of LA in recent years. For example, in Peru, the expansion of this type of credit was accompanied by a boom in the construction sector, mainly real estate.[3] Moreover, in Mexico, recently, the allocation of new loans for housing credit tripled, without considering the subsidized housing credit by the government.

In addition, we consider that entrepreneurs are credit constrained by the value of their collateral asset. In their case, this asset is capital. Hence, we incorporate credit to corporations into our model. As depicted in Figure 1.2, for Peru, this type of credit has increased significantly, growing at a higher rate than credit to households, whereas for Mexico, the opposite occurs.

We calibrate the model to replicate the credit expansion experienced by Peru between 2007 and 2012, and by Mexico between 2006 and 2013. In

[3] In 2008, the construction sector grew 16.5% being one of the sectors that led the Peruvian GDP to grow 9.84% in this year.

particular, we calibrate the credit trajectories of the model to match the credit expansion in data, both credit to corporations and households - including personal and housing credit (see Figure 1.3). We exogenously perturb three time-varying parameters that dictate the tightness of the credit constraints for impatient households and entrepreneurs to match their counterparts in data. The purpose of this exercise is consistent with the idea that a large fraction of these credit expansions was due to exogenous policies that fueled these credit deepening processes. In Peru, there were reforms to stimulate credit to households since 2008. In 2010, a law was promulgated to allow *crédito por convénio*. Since then, this type of credit has shown a strong growth rate, between 2010 and 2012, its participation in personal credit rose from 14.7% to 21.1%. For Mexico, we were not able to find a law related to *crédito de nómina*, but through data available, we found that this type of credit started in 2011.

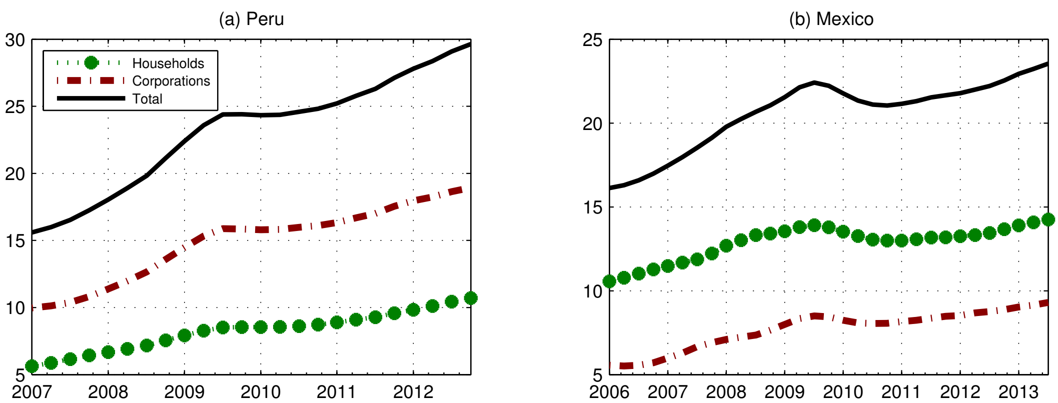


Figure 1.2: Nonearmarked credit outstanding to GDP ratio in Peru and Mexico, by borrower type. Source: Central Reserve Bank of Peru, available at www.bcrp.gob.pe; Bank of Mexico, available at www.banxico.org.mx.

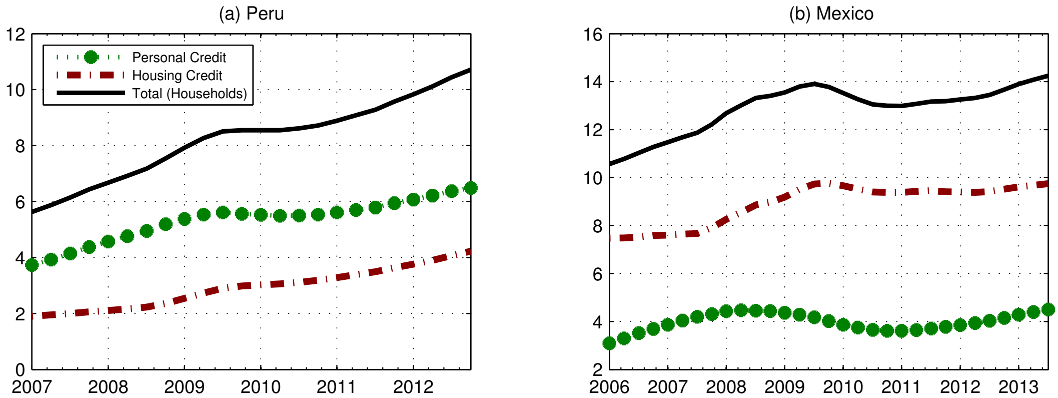


Figure 1.3: Ratio of households nonearmarked credit outstanding to GDP in Peru and Mexico, by type. Central Reserve Bank of Peru, available at www.bcrp.gob.pe; Bank of Mexico, available at www.banxico.org.mx.

From the perspective of our calibrated model, the aggregate effects of the credit deepening process witnessed in Peru and in Mexico are modest. In the case of Peruvian economy, GDP increases only 1.5 percent, consumption barely grows 0.7 percent and investment increases 2.5 percent during the period under review. For Mexican economy, credit deepening increases GDP by almost 0.7% between first quarter of 2006 and third quarter of 2013. During the same period, consumption barely increases and investment grows only 1.3 percent. However, our model does not consider trend growth. For this reason, we need to compare the results generated by our model with above-trend growth experienced by Peru and Mexico during the period of analysis to quantify the macroeconomics effects of these credit deepening processes. If one assumes a trend growth of 5.0 percent per year for Peru, the credit expansion process accounts for 15.6, 6.8 and 3.7 percent of above-trend growth in GDP, consumption and investment, respectively. For Mexico, if we consider a trend growth of 1.5 percent per year, credit deepening accounts for 9.6, 3.5 and 8.4 percent of above-trend growth in GDP, consumption and investment, respectively. Hence, our results suggest that the macroeconomic implications of these credit expansions are modest and accounted for a mild part of above trend growth of these economies.

The remainder of the paper is organized as follows. Section 1.2 presents the related literature. Section 1.3 outlines the baseline model. Section 1.4 describes the quantitative analysis, including the calibration procedure, results and sensitivity analysis. Finally, section 1.5 concludes.

1.2 Related Literature

Our work belongs to a burgeoning strand of literature that incorporates financial frictions into New Keynesian models. This literature builds on the seminal works by Bernanke et al. (1999) and Kiyotaki and Moore (1997). For recent surveys, see Gertler and Kiyotaki (2010), Liu et al. (2013), Iacovello (2014) and Justiniano et al. (2014).

There is a limited literature for Peru and Mexico that uses financial frictions. Recently, some papers use New Keynesian (DSGE) models with financial frictions to address questions related to these countries – e.g., Castillo et al. (2014); Mendoza (2010). However, none of these studies is concerned with the main question of this paper.

Our paper is closely related to Carvalho et al. (2014). Our analyses differ in that, in addition to what they did, we entertain an open-economy version of the model, and experiment with different assumptions about the extent to which the credit deepening process was anticipated. In addition, we contribute to the literature on Latin American economies by providing, to our knowledge, the first assessment of the macroeconomic effects of credit expansions observed in Mexico and Peru.

Similar to Carvalho et al. (2014), we consider the following financial frictions. First, à la Kiyotaki and Moore (1997) and Iacovello (2005), we bind the amount that the entrepreneurs can borrow to the value of their collateral asset, in this case, capital. Note that, through relaxing this borrowing constraint, we can replicate the credit expansion we observe for corporations. Second, we tie the capacity of the impatient household to borrow to his collateral given by the future labor income and, different from Carvalho et al. (2014), we consider housing as well. This financial friction is in line with Peruvian and Mexican experiences, where these types of collateral played a prominent role. Therefore, by relaxing this financial friction, we can emulate consumer credit (personal and housing credit) expansion we observe in practice.

At last, we follow Curdia and Woodford (2010) for modelling the banking sector in which we can generate an endogenous spread between borrowing and lending rates.

1.3 The Baseline Model

This section presents the quantitative model used to analyze the macroeconomics effects of credit expansion in Peru and Mexico, which follows the one in Carvalho et al. (2014). It features eight types of agents: patient households

(p), impatient households (i), entrepreneurs (e), housing and capital producers, retails and final good producing firms. In addition, there is a central bank that defines the basic interest rate by following a simple Taylor-rule.

We assume that impatient households and entrepreneurs have lower discount factor than patient households (i.e., $\beta^e, \beta^i < \beta^p$). For this reason, in equilibrium, patient households are the lenders and impatient households and entrepreneurs are the borrowers in the economy. Both borrowers are subject to credit constraint tied to some type of collateral.

1.3.1 Households

The economy is populated by two types of households i and p which differ by their discount factor β^s. Each type of households has mass ψ^i and ψ^p respectively. Households maximize the following utility function

$$\sum_{t=0}^{\infty}(\beta^s)^t \left\{ \log\left([\xi(C_t^s)^\sigma + (1-\xi)(H_t^s)^\sigma]^{\frac{1}{\sigma}} \right) - \frac{(L_t^s)^{1+\varphi}}{1+\varphi} \right\} \tag{1-1}$$

where C_t^s denotes consumption of final good, L_t^s denotes labor hours and H_t^j is the stock of housing.

Patient Households

Patient households are the lenders in this economy. They have a higher propensity to save, i.e., $\beta^p > \max\{\beta^i, \beta^e\}$. As we are interested in studying the transition from a low to a high credit equilibrium, this condition ensures that patients households are always lenders. They maximize their utility $(1-1)$ subject to the budget constraint given by

$$C_t^p + q_t^S H_t^p + D_t^p \leq W_t^p L_t^p + q_t^H (1-\delta_H) H_{t-1}^p + \frac{(1+r_{t-1}^h)}{\pi_t} D_{t-1}^p + T_t,$$

where $\pi_t = P_t/P_{t-1}$ is the gross inflation rate, q_t^H is the relative price of housing (in terms of final good), δ_H is the rate of depreciation of the housing stock, W_t^p is the real wage and r_t^h is the interest rate on deposits. In particular, given W_t^p, q_t^H and r_t^h, patient households choose a stream of consumption of final good (C_t^p), housing (H_t^p), labor services (L_t^p), and bank deposits (D_t^p). We assume that patient households are the owners of all firms and banks in this economy, therefore all profits are channeled to them, which we denote by T_t.

Impatient Households

Impatient households are one of the borrowers in this economy (i.e., $\beta^i < \beta^p$). They maximize (1 − 1) subject to a budget and a borrowing constraint. The first one is given by

$$C_t^i + q_t^H H_t^i + \frac{1 + r_{t-1}^h}{\pi_t} B_{t-1}^i \leq W_t^i L_t^i + q_t^H (1 - \delta_H) H_{t-1}^i + B_t^i,$$

And the second constraint is the following

$$(1 + r_t^h) B_t^i \leq \tau_t^{WL} \pi_{t+1} W_{t+1}^i L_{t+1}^i + \tau_t^H q_{t+1}^H \pi_{t+1} (1 - \delta_H) H_t^i.$$

Then, given W_t^i, q_t^H and r_t^h, impatient households choose a stream of consumption of final good (C_t^i), housing (H_t^i), labor services (L_t^i), and debt (B_t^i).

Similar to Kiyotaki and Moore (1997), Iacoviello (2005) and Gerali et al. (2010), households' ability to borrow is limited to a fraction τ_t^H of the value of next period's stock of housing, plus a fraction τ_t^{WL} of the value of next period's labor income, which is similar to Carvalho et al. (2014).

Adjusting the collateral requirements by changing τ_t^{WL} to replicate the expansion of personal credit (which includes *credito convénio* for Peru and *credito de nómina* for Mexico) and also τ_t^H to study the expansion of collateralized credit, i.e. mortgage credit for households, we can study the macroeconomic effects of such credit expansion over the period of analysis. In addition, we assume that the household credit rates and the deposit rate r_t^h are the same, motivated by the fact that payroll loan and, mainly, consumer credits apply the lowest interest rate.

1.3.2 Entrepreneurs

The economy is also populated by entrepreneurs that have mass ψ^e. They have the following utility function

$$\sum_{t=0}^{\infty} (\beta^e)^t \log(C_t^e), \tag{1-2}$$

where C_t^e is their consumption of final good.

As for impatient households, entrepreneurs are also borrowers agents in the economy (i.e., $\beta^e < \beta^p$). They use capital K_t and labor (L_t^p, L_t^i) as inputs to produce a wholesale good Y_t^e applying the following production function

$$Y_t^e = K_{t-1}^{\alpha} [(\mu^p L_t^p)^{\theta} (\mu^i L_t^i)^{1-\theta}]^{1-\alpha},$$

where Y_t^e denotes the output for each entrepreneur, K_{t-1} denotes the capital input with its share given by the parameter $\alpha \in (0,1)$ and as in Iacoviello and Neri (2010), we assume complementarity across labor types L_t^p and L_t^i, which participation is measured by the parameter $\theta \in (0,1)$.

Entrepreneurs maximize their utility function subject to the budget constraint

$$C_t^e + W_t^p L_t^p + W_t^i L_t^i + \frac{(1+r_{t-1}^e)B_{t-1}^e}{\pi_t} + q_t^K K_t \leq q_t^W Y_t^e + B_t^e + q_t^K(1-\delta_K)K_{t-1},$$

where δ_K is the depreciation rate of capital, q_t^K is the price of capital in terms of the final good, and $q_t^W \equiv P_t^W/P_t$ is the relative price of the wholesale good Y_t^e. In addition, they are borrowers agents subject to a borrowing constraint given by

$$(1+r_t^e)B_t^e \leq \tau_t^K q_{t+1}^K \pi_{t+1}(1-\delta_K)K_t,$$

where r_t^e is the nominal interest rate faced by entrepreneurs.[4] The amount that entrepreneurs can borrow is limited by a fraction τ_t^K of the value of the collateral asset. In their case, this asset is capital.

Finally, we interpret τ_t^K as an exogenous shock on the entrepreneurs' ability to borrow. Thus, imposing an exogenous path to τ_t^K, we can replicate the expansion of corporate credit in Peru and Mexico and study the macroeconomics effects of the expansion credit over the period of analysis.

1.3.3 Firms

In this section, we study four types of firms: retail firms operating in a monopolistic competitive market, final goods producers, producers of housing and capital producers. All firms are owned by patient households.

Retail Firms and Final Goods Producers

Retailers buy the intermediate good from entrepreneurs at the wholesale price P_t^W. It takes one unit of intermediate output to make a unit of retail output. We introduce nominal rigidities through Rotemberg pricing scheme. In particular, we assume that each firm operates in a monopolistic competitive environment and faces a quadratic cost for adjusting the nominal prices. Moreover, each retailer's price is indexed to a combination of past and steady-state inflation with relative price equal to ι and $(1-\iota)$, respectively.

[4] In the section of Banks, we explain how the credit spread, $\omega_t = (1+r_t^e)/(1+r_t^h) - 1$, is determined endogenously.

Final good producers are competitive, and their role is to simply aggregate, using a CES composite, the continuous sequence of differentiated varieties produced by retailers. The final output composite is given by

$$Y_t = \left[\int_0^1 Y_t(f)^{\frac{\varepsilon-1}{\varepsilon}} df\right]^{\frac{\varepsilon}{\varepsilon-1}},$$

where $Y_t(f)$ is the output produced by retailer f, ε is the elasticity of substitution between varieties and P_t is the associated Dixit-Stiglitz price index. This final good is purchased by patient households, impatient households and entrepreneurs for consumption, and by capital goods and housing producers for production.

Retailer pricing problem is to choose the optimal price $P_t(f)$ that solves the maximization of its profit

$$\sum_{t=0}^{\infty} \Delta_t \left[P_t(f)Y_t(f) - P_t^W Y_t(f) - \frac{\kappa_P}{2}\left(\frac{P_t(f)}{P_t(f-1)} - \pi_{t-1}^{\iota}\bar{\pi}^{1-\iota}\right)^2 P_t Y_t\right],$$

subject to a demand schedule deriving from the cost-minimization problem of final goods producers

$$Y_t(f) = \left(\frac{P_t(f)}{P_t}\right)^{-\varepsilon} Y_t.$$

where the parameter κ_P rules the degree of price stickiness, Rotemberg type, and $\bar{\pi}$ denotes steady-state inflation. Finally, all profits generated in this sector are transferred to patient households.

Producers of Housing

At the beginning of each period t, the production of new housing is performed by competitive firms. They buy an amount of final good I_t^H from final goods firms and the stock of undepreciated housing $(1-\delta_H)H_{t-1}$ at relative price q_t^H from both patient and impatient households. The stock of undepreciated housing is transformed one-to-one into new housing, while the transformation of final goods into new housing is subject to a quadratic adjustment cost. Then, the level of production is chosen to maximize

$$\sum_{t=0}^{\infty} \Delta_t[q_t^H(H_t - (1-\delta_H)H_{t-1}) - I_t^H],$$

subject to the following law of motion of housing

$$H_t = (1-\delta_H)H_{t-1} + \left[1 - \frac{\kappa_H}{2}\left(\frac{I_t^H}{I_{t-1}^H} - 1\right)^2\right] I_t^H,$$

where $\kappa_H > 0$ is the adjustment cost parameter and, Δ_t is the stochastic discount factor of patient households. The new capital H_t is sold at relative price q_t^H to both patient and impatient households.

Capital Producers

At the beginning of each period t, competitive capital producing firms buy a stock of undepreciated capital $(1-\delta_K)K_{t-1}$ from entrepreneurs at a nominal price q_t^K and an amount of the final good I_t^K from final goods firms. The undepreciated capital is transformed one-to-one into new capital, while the transformation of final goods into new capital is subject to a quadratic adjustment cost. Then, the level of production is chosen to maximize

$$\sum_{t=0}^{\infty} \Delta_t [q_t^K (K_t - (1-\delta_K)K_{t-1}) - I_t^K],$$

subject to the following law of motion of capital

$$K_t = (1-\delta_K)K_{t-1} + \left[1 - \frac{\kappa_K}{2}\left(\frac{I_t^K}{I_{t-1}^K} - 1\right)^2\right] I_t^K,$$

where $\kappa_K > 0$ is the adjustment cost parameter and, Δ_t is the stochastic discount factor of patient households. The new capital K_t is sold back to entrepreneurs at relative price q_t^K.

1.3.4 Banks

We assume competition among banks, both in the loan market and deposit market. Thus, r_t^h and r_t^e are taken as given. Each bank collects deposits from patient households D_t, which are used as resources to lend to impatient households and entrepreneurs. Following Cúrdia and Woodford (2010), we consider that there is a real resource cost in loans to entrepreneurs. Such cost is given by $\eta(B_t^e)^\gamma$, where $\eta > 0$ and $\gamma > 1$, and it can be interpreted as agency and operational costs that are not considered in our model.

All excess funds of the banks are transferred to patient households, $D_t - B_t^e - B_t^i - \eta(B_t^e)^\gamma$. Credit spread ω_t is defined implicitly by $(1+r_t^e) = (1+\omega_t)(1+r_t^h)$. Since assets must be equal to liabilities $D_t = (1+\omega_t)B_t^e + B_t^i$, it follows that

$$\omega_t B_t^e - \eta(B_t^e)^\gamma \tag{1-3}$$

Thus, banks maximize $(1-3)$ to choose B_t^e. The first order condition for optimal credit supply is given by

$$B_t^e = (\eta\gamma/\omega_t)^{1/(1-\gamma)}$$

Note that, since $\gamma > 1$, there is a positive correlation between the credit spread ω_t and loans to entrepreneurs B_t^e.

1.3.5 Monetary Policy

We assume that monetary policy is conducted by means of a Taylor-type interest rate rule of the form

$$(1 + r_t^h) = (1 + r)^{1-\rho}(1 + r_{t-1}^h)^{\rho} \left(\frac{\pi_t}{\pi}\right)^{\phi_\pi(1-\rho)} \left(\frac{y_t}{y_{t-1}}\right)^{\phi_y(1-\rho)},$$

where ρ is the parameter that measures the smoothing of the interest rate, ϕ_π and ϕ_y are the weights assigned to inflation and output stabilization, respectively, and π and r are the steady-state levels of inflation and the policy rate, respectively.

1.3.6 Market Clearing

The market clearing condition for the wholesale good is given by:

$$\int_0^1 Y_t(m)dm = \mu^e Y_t^e.$$

Equilibrium in the final good market is expressed by the resource constraint:

$$Y_t = C_t + I_t^H + I_t^K + \eta(B_t^e)^\gamma + \text{ all adjustment costs.}$$

where $C_t = \mu^p C_t^p + \mu^i C_t^i + \mu^e C_t^e$.

We consider that both types of labor markets are competitive.

1.4 Quantitative Analysis

We calibrate our baseline model, then we use it to quantify the macroeconomic effects of the credit expansion observed in Peru and Mexico by solving for the time-varying paths of τ_t^{WL}, τ_t^H and τ_t^K that generate paths for personal credit, housing credit to households, and credit to corporations that correspond their counterparts in the data (see Figures 1.2 and 1.3). Due to a better data availability for the calibration of the model, we focus in Peruvian economy in the main body of our paper and we present the quantitative analysis for Mexico in the appendix. In addition, we perform some sensitivity analysis to test the robustness of our results.

1.4.1 Calibration

Table 1.1 lists the choice of parameter values for our baseline model that matches with the statistics for Peruvian economy. We consider its average between 2007 and 2012. Time is in quarters. Steady state inflation is 3.37% to match the average inflation rate for the period. We use $\beta^p = 0.9984$ to generate an average interbank nominal interest rate (proxy of monetary policy rate) of 4.03%.

We set the discount factor of impatient and entrepreneurs agents as $\beta^i = \beta^e = 0.91$, implying an annual time-discount rate of 52 percent. We calibrate this extreme value motivated, mainly, to maintain the borrowing constraints active during all the transition period and because with a greater degree of impatience, i.e $\beta^{i,e} < \beta^p$, the model increases its capacity to produce significant aggregate effects.

We also pick the inverse of the Frisch elasticity of labor supply equal to $\varphi = 2$. This value is standard in the literature for Peruvian economy. Following Fernandez-Villaverde and Krueger (2004), we calibrate the parameters associated with preferences for final goods and housing in the utility function. In the absence of estimates for σ, we set it to zero. Then, the aggregator function takes a Cobb-Douglas form $(C_t^j)^\xi (H_t^j)^{1-\xi}$, $j = i, p$. We define the share of consumption of final goods and housing in the Cobb-Douglas aggregator of the utility with ξ equals to 0.8.

For the capital share in the entrepreneurs' production function, we choose $\alpha = 0.26$ following Castillo et al. (2009). Since information on patient and impatient labor income shares in Peru is not available, we set $\theta = 0.7$ following Carvalho et al. (2014). The depreciation rates for capital and housing are set, respectively, to $\delta_K = \delta_H = 0.025$. The adjustment cost parameter for capital and housing is set $\kappa_K = \kappa_H = 2.53$ following Carvalho et al. (2014).

The parameter governing price stickiness (Rotemberg adjustment cost) in the retail sector κ_P is set at 58 which is equivalent to 0.75 in the Calvo model. As usual, it is possible to map these two types of price stickiness since this entails the same first order dynamics of the two models in the case of zero steady state inflation. The elasticity of substitution between varieties is $\varepsilon = 6$, which yields a steady state mark-up of 20 percent. Finally, ι, which governs indexation, is set to 0.158, as in Gerali et al.(2010).

For the monetary policy rule, we use estimates from Castillo et al. (2009). In particular, we choose $\phi_y = 0.16$, $\phi_\pi = 1.5$ and $\rho = 0.79$.

Concerning the calibration of the banking sector parameters, we set $\gamma = 2$ and $\eta = 0.01$ to generate a spread of roughly 0.7 percent per year. This value corresponds to the average difference between the corporate prime interest rate, which is the lending interest rate that banks charge their best corporate clients and the reference interest rate, which is the interest rate that BCRP fixes in order to establish a level of reference interest rate for interbank transactions. Loans to these firms embed lower default risk than loans to other firms. Hence, the targeted value of 0.7 percent per year underestimates the average spread in Peruvian economy. As we show below, in the sensibility section, the calibration of γ and η helps to produce more significant aggregate effects in response to the credit deepening process. Finally, the masses of different agents in the economy ψ^p, ψ^i and ψ^e is set to one.

Parameter	**Description**	**Value**
β^p	Discount Factor - Patients	0.9984
β^i, β^e	Discount Factor - Impatients and Entrepreneurs	0.91
ψ^p, ψ^i, ψ^e	Mass - Patients, Impatients and Entrepreneurs	1
φ	Inverse of the Frisch Elasticity	2
σ	Elasticity Between Final Good and Housing	0
ξ	Weight of the Final Good on the Utility Function	0.8
δ_K, δ_S	Depreciation - Capital Goods and Housing	0.025
κ_K, κ_S	Adjustment Cost - Capital Goods and Housing	2.53
α	Capital Share in the Production Function	0.26
θ	Share of Patient Households in the Production Function	0.7
κ_P	Price Adjustment Cost - Final Good	58
ι	Steady State Inflation Weight - Indexation	0.158
ε	Elasticity of Substitution - Final Good	6
ρ	Smoothing Parameter of the Taylor Rule	0.7
ϕ_y	Output Weight of Taylor Rule	0.1
ϕ_π	Inflation Weight of Taylor Rule	1.5
η	Spread	0.008
γ	Spread	2

Table 1.1: Calibration: Peruvian Economy.

1.4.2 Results

In this section, we study the macroeconomics effects of a credit deepening using the calibrated model for Peruvian economy.[5] For such aim, we solve for the time-varying paths of τ_t^{WL}, τ_t^{H} and τ_t^{K} that generate paths for personal credit, housing credit to impatient households and corporate credit for entrepreneurs. We follow Justiano et al. (2014) and assume that the evolution of τ_t^{WL}, τ_t^{H} and τ_t^{K} are perfectly foreseen after the start of the credit expansion process in 2007, which is an initial unforeseen shock. After 2012, we set τ_t^{WL}, τ_t^{H} and τ_t^{K} constant. In addition, we smooth the paths of τ_t^{WL}, τ_t^{H} and τ_t^{K} using a third degree polynomial.

Figure 1.4 shows their calibrated paths and Figure 1.5 compares the credit expansion generated by our model and their counterparts in the data. Note that our model replicates well enough the trajectories of considered types of credit during the period of analysis.[6]

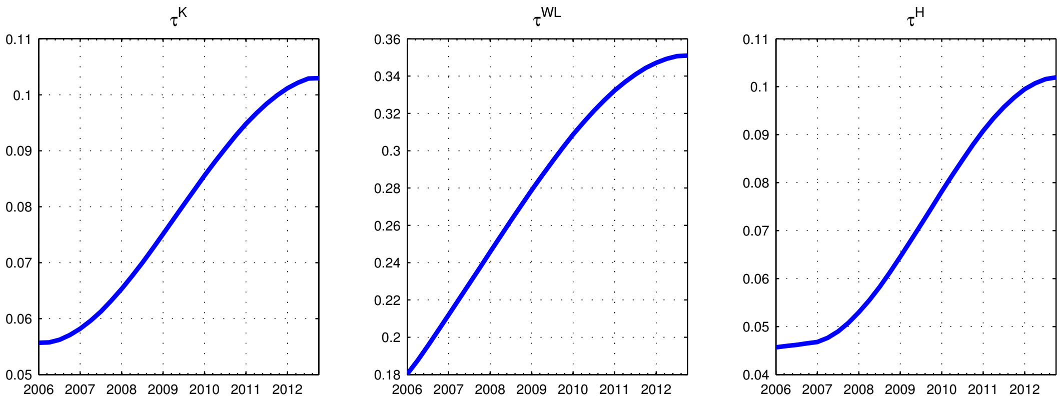


Figure 1.4: Credit deepening experiment for Peru: evolution of τ_t^{K}, τ_t^{WL} and τ_t^{H}.

[5]We used the shooting algorithm in Dynare to solve our model.

[6]Carvalho et al. (2014) report, through a robustness test, that their results do not change once they considered non-smooth paths.

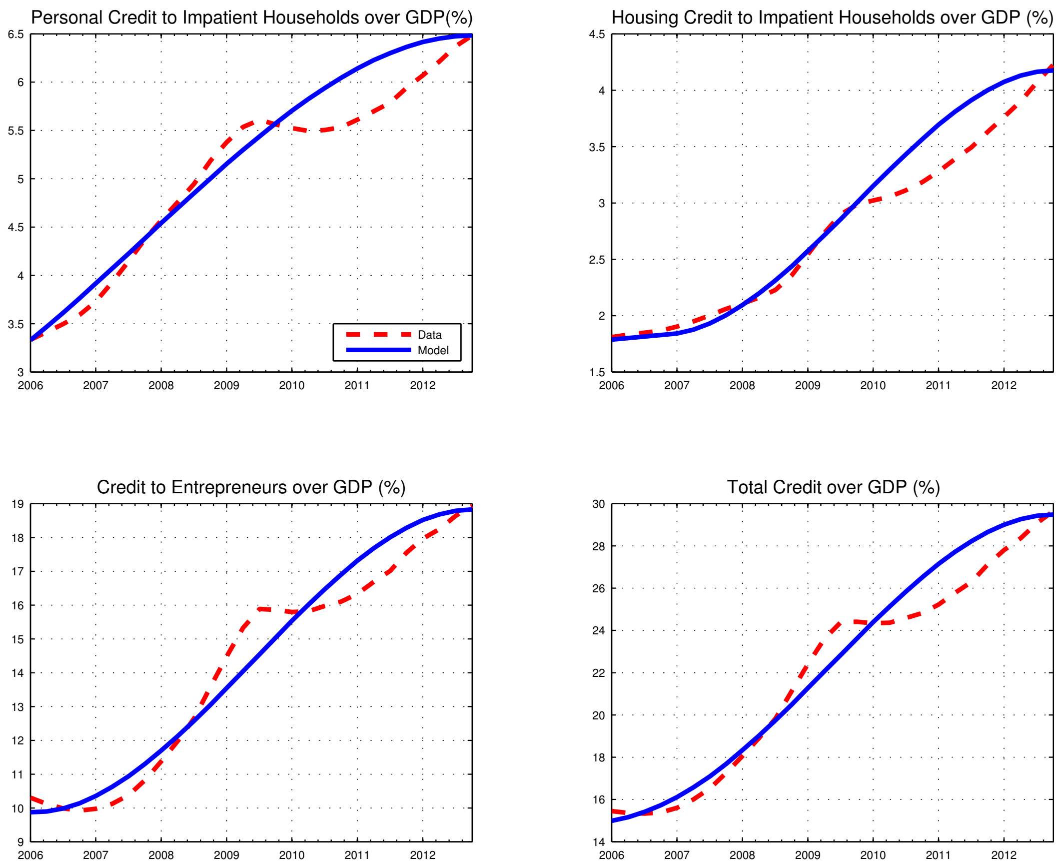


Figure 1.5: Credit deepening experiment for Peru: credit variables (model and data).

The macroeconomic effects of a credit expansion in our model are depicted in Figure 1.6, through the trajectories of GDP, consumption, investment and inflation. The aggregate implications are small in absolute terms. Among the main key variables, investment increases 2.5 percent, which is the highest macroeconomic effect followed by GDP that increases 1.5 percent. Finally, consumption barely grows 0.75 percent.

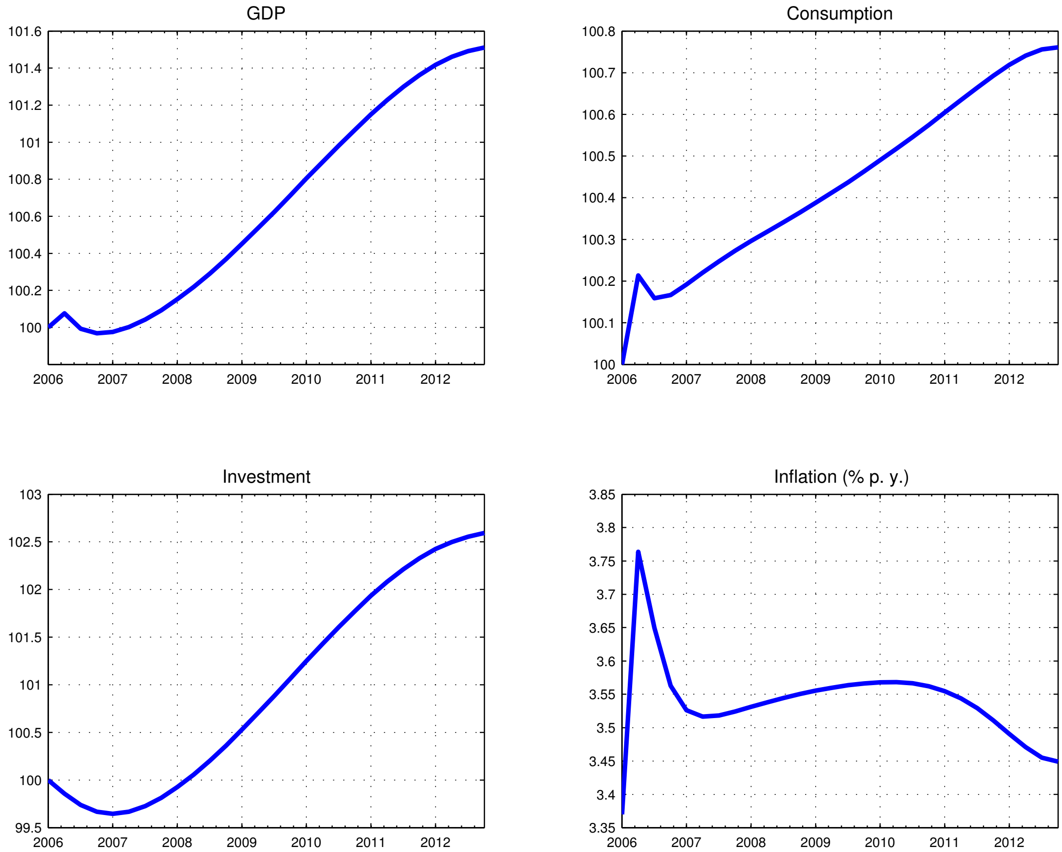


Figure 1.6: Credit deepening experiment for Peru: macro variables (model).

Concerning the effects on disaggregated level, the evolution of investment as well as the stock of housing, capital and consumption of final goods for all agents in the model are presented in Figure 1.7. Note that when collateral constraints are loosened for impatient households and entrepreneurs, their consumption and investment in housing and in capital, respectively, increase. As a consequence of a higher demand, the price of the final good increases (see Figure 1.6) and patient households reduce their consumption of final goods and investment in housing. As the process evolves, consumption and investment of patient households also increase.

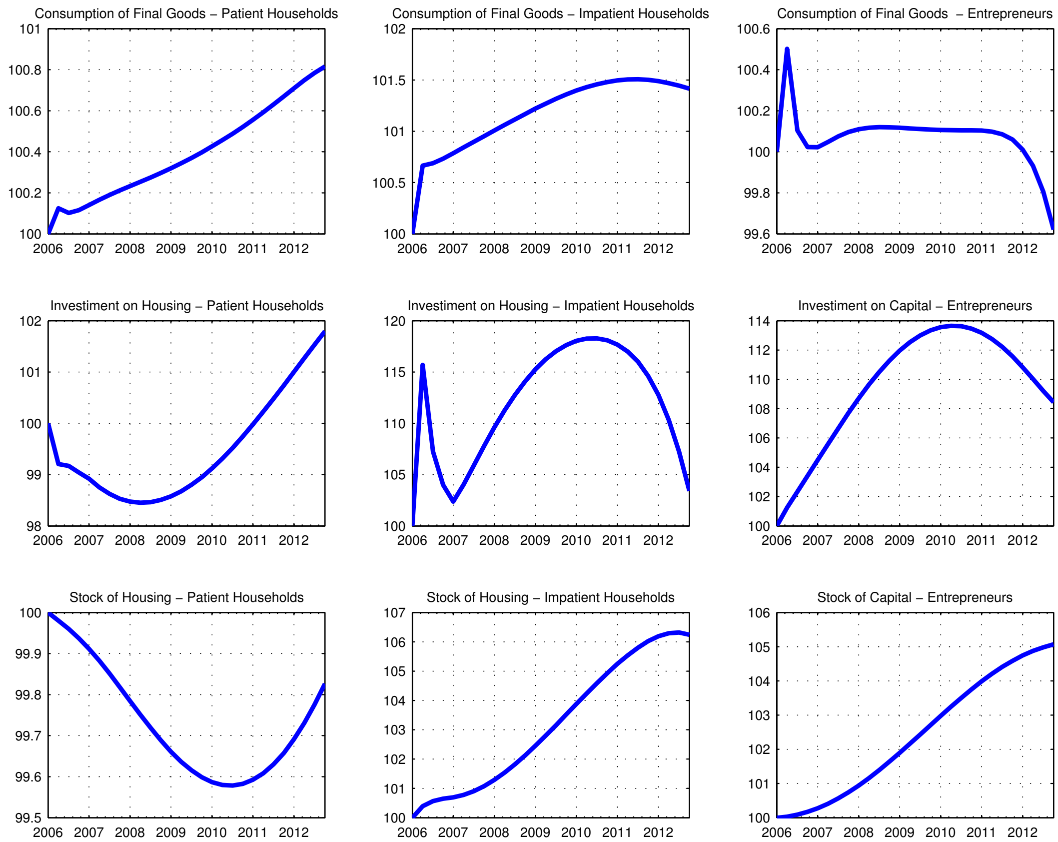


Figure 1.7: Credit deepening experiment for Peru: consumption, investment and stocks.

At the beginning of the credit deepening process, impatient households increase in almost 32 percent their investment in housing and then, it tends to decline, whereas their consumption increases in more than 1 percent and it is kept in this level until the end of the period. Furthermore, initially, the stocks of housing of patient and impatient households move inversely, but in the last periods both increase monotonically. Therefore, in the first years of the credit expansion process, market clearing prices involve that patient households exchange housing for final goods, unlike the impatient ones that consume more housing and final goods. It is noteworthy that housing has a double function in the model. First, it generates utility for impatient and patient households. Second, its value enters as collateral in the credit constraint for the impatient households. Thus, the accumulate stock of housing becomes more valuable for impatient households than for patient ones.

Moreover, investment of entrepreneurs in capital follows an inverse U shaped pattern through the credit expansion process and at the end, it increases 8 percent. This allows them to increase their stock of capital by nearly

5 percent. Also, at the beginning, their consumption of final goods increases by almost 3 percent, but the cumulative effect at the end of the process is zero.

Overall, our model implies that the effects are larger for impatient households, specially, on their investment in housing that increases almost 3 percent and their consumption that rises 1.4 percent.

Figure 1.8 presents the evolution of labor market outcomes. Once credit expansion started, labor services and wages of patient and impatient households move in the same direction, but in different magnitudes. For impatient households, their labor services increase more than their wages, which reflects an extra motive to supply labor due to the fact that it relaxes their credit constraint. For patient households, their wages increase more than their labor services, which could be explained by forces coming from the side of labor demand. The same interpretations are valid for the cumulative effect at the end of the period.

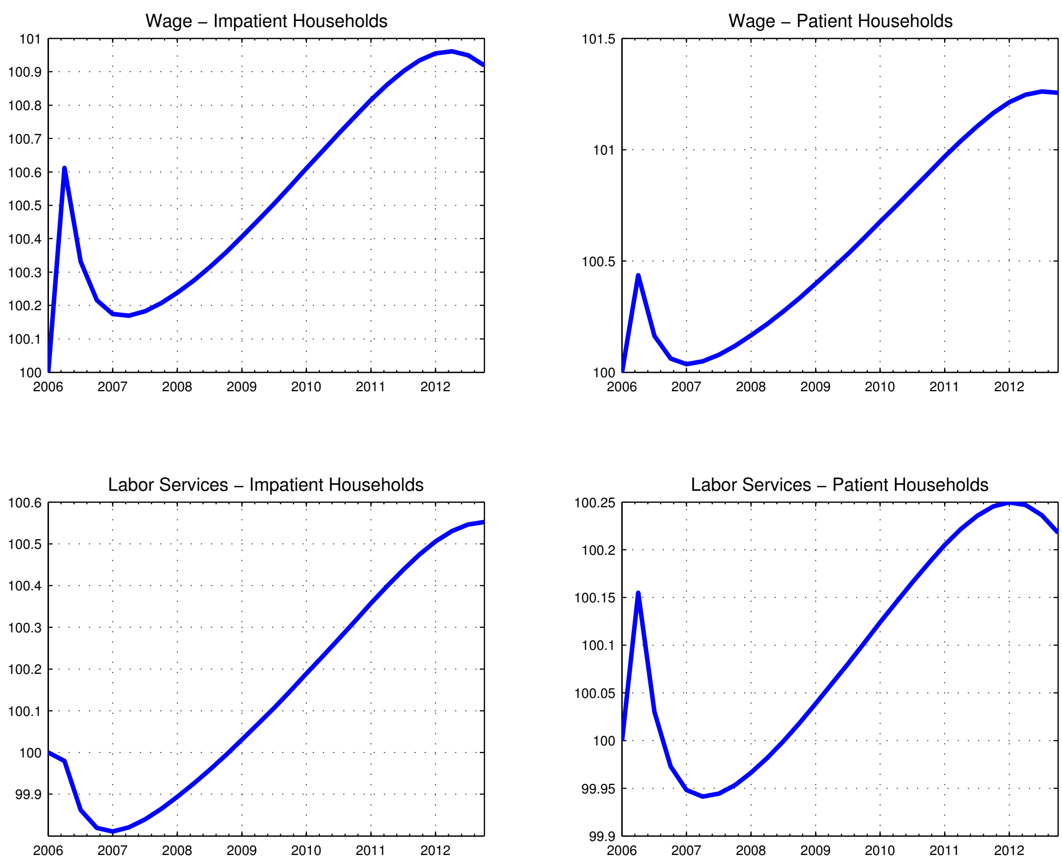


Figure 1.8: Credit deepening experiment for Peru: labor market outcomes.

Figure 1.9 illustrates the evolution for the interest rates and the spread. At the beginning of the period, the deposit interest rate set by the Central Bank increases around 0.4 percentage points and as the credit deepening process unfolds, this interest rate fluctuates between 4.0-4.5 percent. Concerning the interest rate faced by entrepreneurs, at first, it rises substantially and so does the spread. This behavior is due to the fact that credit to entrepreneurs increases, then the intermediation costs associated with these funds also increases, which leads to higher interest rate and spread.

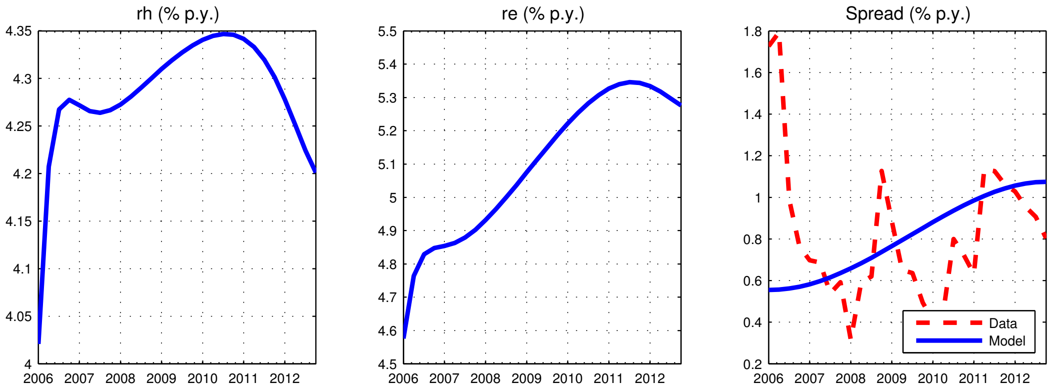


Figure 1.9: Credit deepening experiment for Peru: financial market outcomes. The spread is calculated using the BCRP reference interest rate, which is the interest rate that the BCRP fixed in order to establish a level of reference interest rate for interbank transactions, and the Corporate prime interest rate, which is the lending interest rate that banks charge their best corporate clients. Source: Central Reserve Bank of Peru, available at www.bcrp.gob.pe.

In conclusion, these results suggest that the credit deepening experienced by Peru had relatively modest effects on its main macroeconomic variables. However, our model do not consider trend growth. For this reason, we follow Carvalho et al. (2014) and consider six scenarios for trend growth for Peru, in particular, a range between 3.5 to 6 percent per year, so we can quantify the share of above-trend growth in GDP, consumption and investment between 2007 and 2012 that can be explained by the credit expansion. For each scenario, we divide the cumulative effect in our model for each aggregate variable by the cumulative above-trend growth in the real data.

Among the different scenarios depicted in Table 1.2, we highlight the trend growth of 5.0 percent per year. In this case, the credit expansion process accounts for 15.6, 6.8 and 3.7 percent of above-trend growth in GDP, consumption and investment, respectively. If we consider the worst scenario for trend growth (3.5 percent per year), credit expansion explains 7.7 percent of above-trend GDP growth. On the other hand, under a more positive scenario, i.e., with a trend growth of 6.0 percent, the model accounts for up to 42.2% of the gap for GDP.

	GDP		**Consumption**		**Investment**	
	Growth (data): 46.9% Growth (model): 1.5%		Growth (data): 48.7% Growth (model): 0.75%		Growth (data): 124.6% Growth (model): 2.5%	
Trend growth (% p.y.)	Above trend growth (%)	Model share (%)	Above trend growth (%)	Model share (%)	Above trend growth (%)	Model share (%)
3.5%	19.5%	7.7%	21.0%	3.6%	82.7%	3.0%
4.0%	16.1%	9.3%	17.5%	4.3%	77.5%	3.2%
4.5%	12.8%	11.7%	14.2%	5.3%	72.5%	3.4%
5.0%	9.6%	15.6%	11.0%	6.8%	67.6%	3.7%
5.5%	6.5%	22.9%	7.8%	9.6%	62.9%	4.0%
6.0%	3.6%	42.2%	4.8%	15.5%	58.3%	4.3%

Table 1.2: Credit expansion experiment for Peru: comparison with the real data between 2007-2012. Growth rates of GDP, consumption and investment are obtained from Central Reserve Bank of Peru, available at www.bcrp.gob.pe.

To summarize, these results suggest that credit deepening had moderate implications on key macroeconomic variables on Peruvian economy.

1.4.3 Sensitivity Analysis

In this section, we show that our results on the macroeconomic effects of a credit deepening are robust both to alternative calibrations of key parameters of the model, and to modifications on considered assumptions. As in Carvalho et al. (2014), we show that more extreme calibrations of β^e and β^i (see Figure 1.10), γ (see Figure 1.11) and η (see Figure 1.12) do not generate substantial aggregate effects.[7] In addition, we show that the assumption of a closed economy do not change our conclusions. Moreover, we argue that price stickiness is not the driving force behind our results. Finally, we show that the macroeconomic effects are slightly smaller if we drop the assumption that the credit deepening process is perfectly foreseen.

[7] Also as in Carvalho et al. (2014), our results are robust to variations in the impatient labor share θ and capital share α. Results are available upon request.

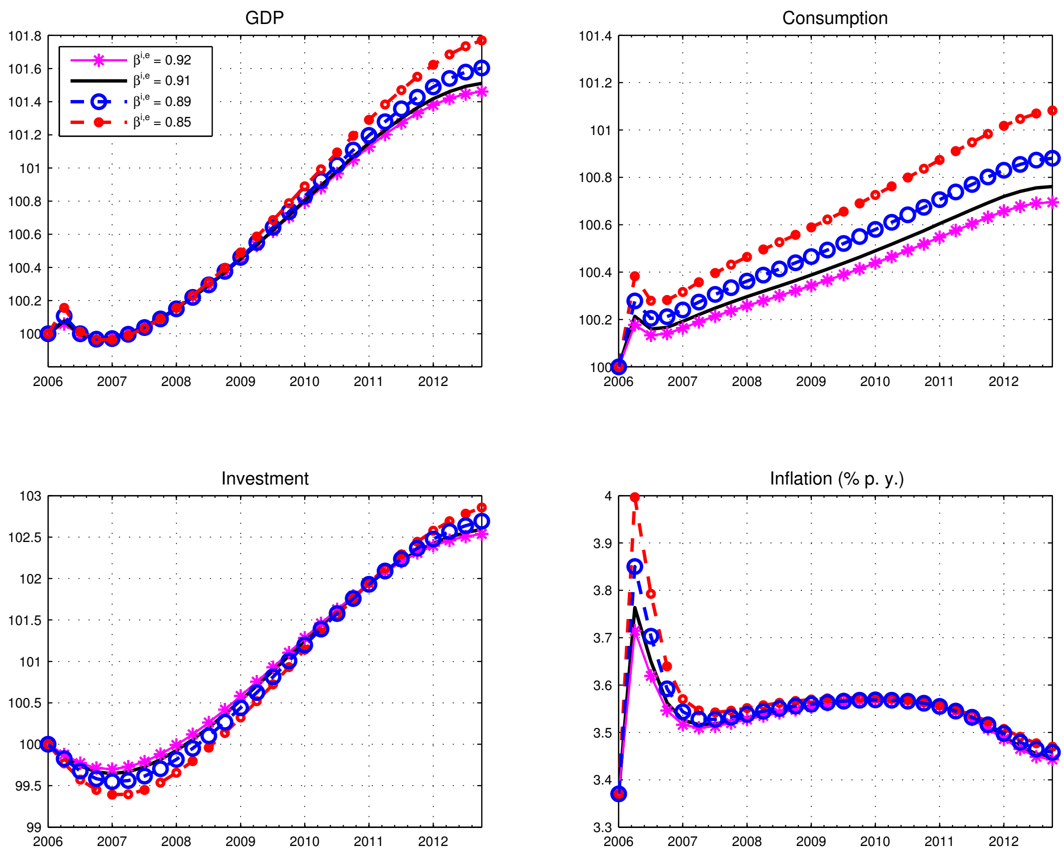


Figure 1.10: Sensitivity analysis - Peru: β^e and β^i.

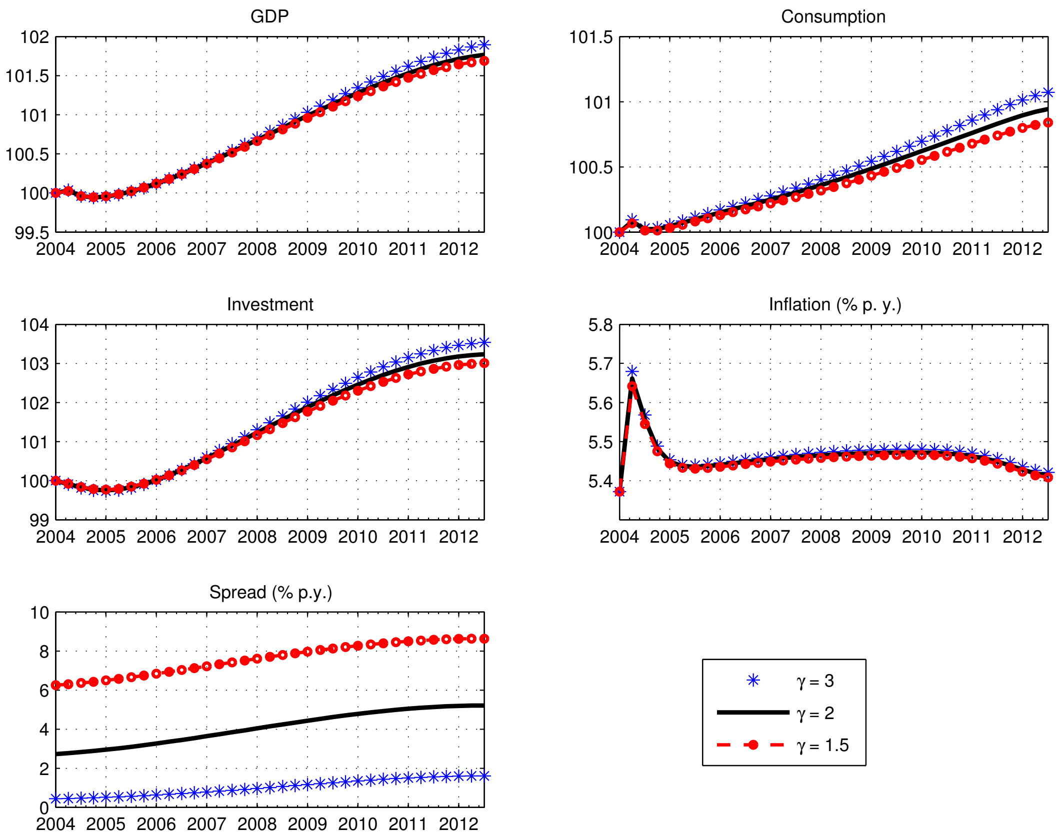


Figure 1.11: Sensitivity analysis - Peru: γ.

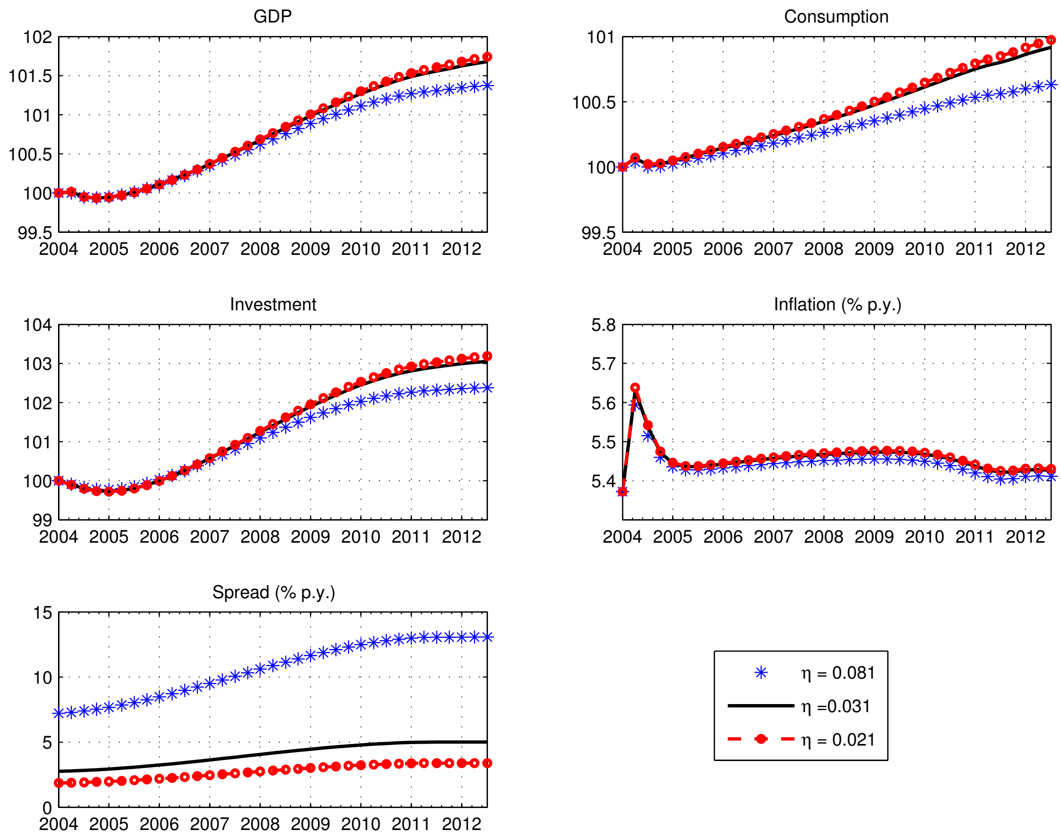


Figure 1.12: Sensitivity analysis - Peru: η.

A small open economy version of our model As we mentioned before, patient and impatient households behave in opposite ways in response to a credit deepening, which might mitigate the aggregate effects. This occurs because borrowers and lenders are the only agents engaged in the credit market in our closed economy. Hence, we consider an extreme opposite case of a small open economy (SOE) without lenders as in Justiniano et al. (2014).

In our SOE version of the baseline model, the economy is populated by impatient households and entrepreneurs with discount factors β^i and β^e, respectively. Both types of agents borrow from abroad. We consider a constant real world interest rate r^*, which is applied in loans for impatient households. There is a credit spread in loans for entrepreneurs that is the same as in the baseline. Moreover, the final goods market is competitive (i.e., there are no retailers) and prices are flexible. Then, market clearing condition of final good is given by

$$Y_t + CAD_t = C_t + I_t^H + I_t^K + \eta(B_t^e)^\gamma + \text{all adjustment costs},$$

where r^e_t is the real interest rate faced by entrepreneurs and $CAD_t = (B^i_t - (1+r^*)B^i_{t-1}) + (B^e_t - (1+r^e_t)B^e_{t-1})$ is the current account deficit. The specifications for producers of housing and capital goods are the same as in the closed economy version.

We calibrate the constant real world interest rate to 1% p.y., which is the steady state real interest rate in our closed economy model. The paths of τ^{WL}_t, τ^H_t and τ^K_t are recalculated so the model can replicate the same smooth credit expansion as in the baseline. Finally, we set $\eta = 0.006$ to generate the same average spread as in the closed economy. The rest of the parameters remains unchanged.

Figure 1.13 shows these results. Consumption grows more at the beginning than the closed economy, but the cumulative effect is smaller. This occurs because, in the baseline, consumption of patient households also grows as the credit deepening process evolves (see Figure 7). Investment expands 8 percent, which is higher than the closed economy. GDP decreases at the beginning due to the effect on the current account deficit of interest payments on a higher stock of debt. Also the cumulative effect is lower. These results suggest that the modest aggregate effects to credit expansion is not due to the closed economy assumption.

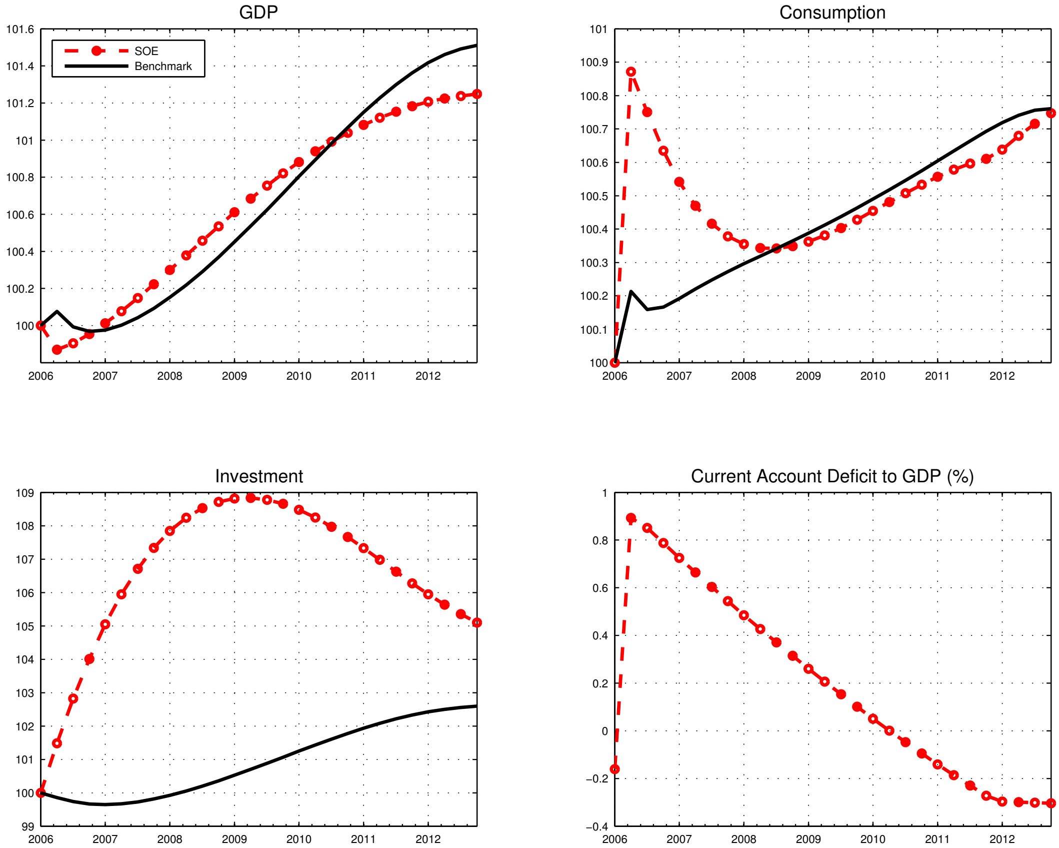


Figure 1.13: Sensitivity analysis - Peru: Small open economy-macro variables (model).

Flexible prices One may wonder about the relevance of price stickiness for our results. To analyze this issue, we set the parameter that determines the degree of price stickiness, κ_P, equal to zero – thus eliminating price rigidities from the model. Results in Figure 1.14 show that, except for small differences in the first few observations, the trajectories of output, investment, consumption, and inflation overlap almost perfectly with those produced by the baseline calibration. We can conclude that price rigidities are not the driving force behind our results.

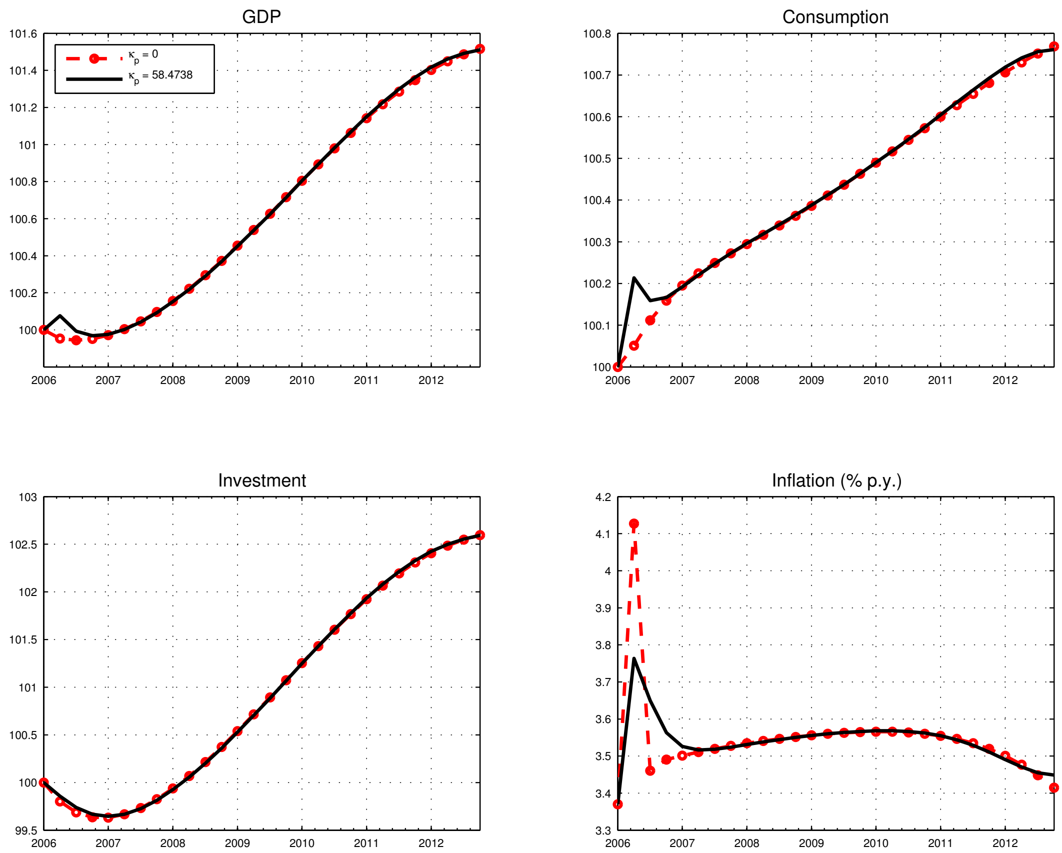


Figure 1.14: Sensitivity analysis - Peru: κ_P.

Unanticipated shocks The assumption that agents perfectly foresee the intensity of the credit deepening process over such a long horizon might be unrealistic. Hence, as a last robustness exercise, we solve the model under an assumption on the other extreme of the "foresight spectrum". Namely, we assume that the credit deepening process takes the form of a sequence of unanticipated shocks to the parameters that govern the credit constraints. Reality should arguably be somewhere in between these two extremes assumptions about agents' foresight. In each period, agents are surprised by new values of τ_t^{WL}, τ_t^{H}, and τ_t^{K}, chosen to fit the trajectories of the credit variables (see Figure 1.15). Figure 1.16 reports the trajectories of the macroeconomic variables that result from this alternative version of the model. The macroeconomic effects of credit deepening are slightly smaller in this case. Output and consumption, for instance, increase by 1.1 and 0.5 percent, respectively, rather than 1.5 and 0.75 percent under the perfect foresight assumption.

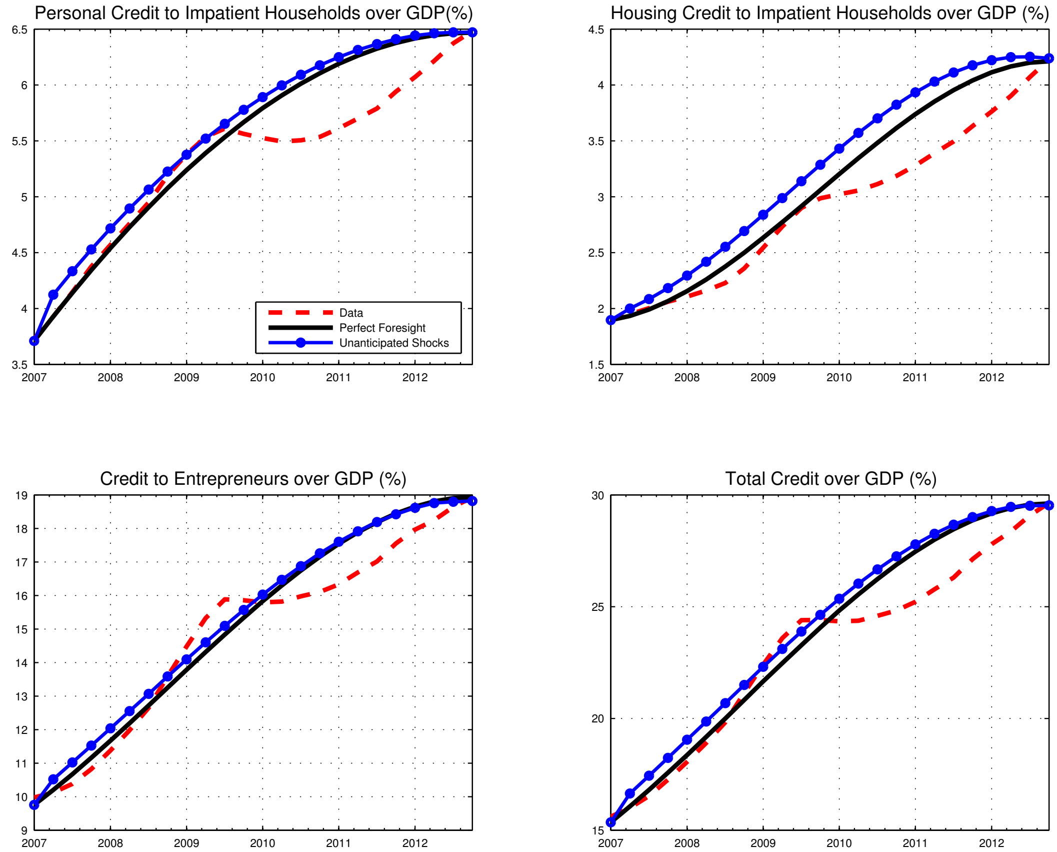


Figure 1.15: Credit deepening experiment for Peru (non-smooth): credit variables (data and model).

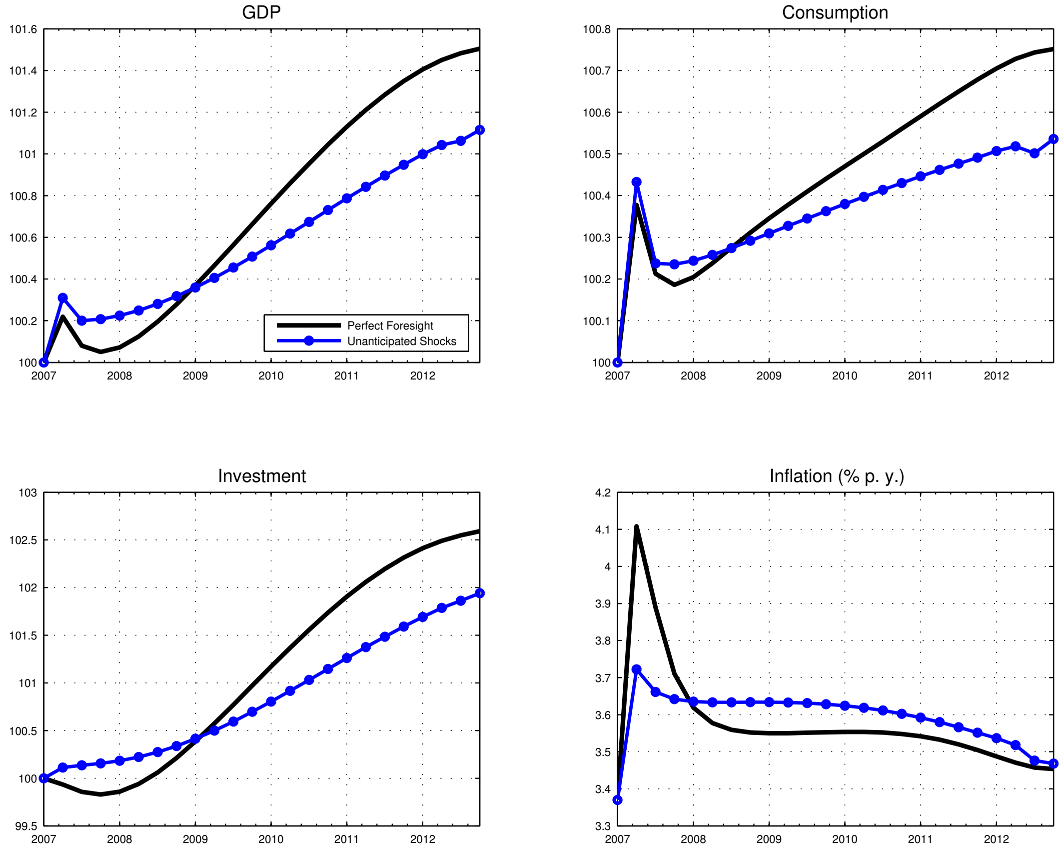


Figure 1.16: Credit deepening experiment for Peru (non-smooth): macro variables (model).

1.5 Conclusion

Recently, Mexico and Peru experienced an intense credit expansion and Peru, in particular, experienced strong economic growth. In order to quantify the macroeconomic effects of these credit deepening processes, we calibrate a standard New Keynesian dynamic general equilibrium model with financial frictions for Peruvian and Mexican economies. Our results suggest that the macroeconomic effects of these credit expansions are modest and accounted for only a small part of above-trend growth in these economies. In addition, we show that these results are robust both to alternative calibrations of key parameters of the model, and to different modeling assumptions. Even when we consider a small open economy version of the model, the effects are still relatively small. One possible reason of these results is that these credit deepening processes coincided with surges in commodity prices, which improved substantially the terms of trade of most of Latin American countries, specially, for Peru. These surges might be one of the leading driving forces behind recent growth. This is a channel that we do not consider in the model, and which we believe should be analyzed in future research.

2
Consumption Boom and Credit Deepening: The Role of Inequality[1]

2.1
Introduction

Does income inequality play a role in the association between the variation of credit and consumption growth? It is not clear a priori how the relationship between income inequality, credit and consumption growth works. In this paper, we address this question through two approaches. First, we document, through cross-country and panel estimations, that the association between consumption growth and credit expansion is stronger in countries with higher income inequality. Second, we use a Aiyagari model to check to which extent this theoretical framework can rationalize the empirical finding. We choose this model because it is the workhorse macroeconomic model used to study quantitatively the interactions between inequality and macroeconomic outcomes. Our interest lies on the role of income inequality on the response of consumption to a credit deepening.

We consider an incomplete markets model with heterogeneous households, idiosyncratic risk and borrowing constraints. The mechanism behind our theoretical model is the following: a loosening of credit constraints mitigates precautionary motives, inducing households to reduce savings along the transition path to the new steady-state. Hence, consumption grows more rapidly in the short-run. This consumption boom is amplified in economies with more households close to the borrowing constraint.

Since we are interested in the role of income inequality on the response of consumption to a credit deepening, we consider two sources of income inequality in our model: the variance of the idiosyncratic risk and the households' fixed level of human capital. They have different implications for the extent to which households are credit constrained in equilibrium.

Our quantitative experiment consists of analyzing, for each source of income inequality, the consumption response to an exogenous credit deepening

[1]This is a joint work with Nilda Pasca.

in economies with different levels of income inequality. We consider the Unites States (US) as the benchmark economy and calibrate our model using standard parameters in the literature. We also consider four economies: two with lower income inequality and the other two with higher income inequality than the benchmark. In each case, we change only one source of income inequality and we set the others in the same level as in the benchmark economy.

Table 2.1 summarizes our main results. The consumption growth at the peak in the benchmark economy is 1.79 percent. We show that when the source of income inequality comes from households' lowest fixed level of human capital, our model can rationalize the empirical evidence. In this case, the economy with the highest income inequality grows at the peak 2.06 percent and the one with the lowest inequality, 1.69 percent. Once other sources of income inequality are considered, the opposite occurs.

Consumption per capita peak growth Benchmark economy: 1.79%		
Source of Inequality	Highest Inequality	Lowest Inequality
Households' lowest fixed level of human capital	2.06%	1.69%
Households' highest fixed level of human capital	1.24%	2.21%
Variance of the idiosyncratic risk	0.13%	5.45%

Table 2.1: Results: Summary of main results.

In addition, we perform an exercise to understand the mechanisms behind our model by calculating the contribution of assets distribution and policy functions on the consumption growth at the peak. Our results suggest that each source of income inequality has different implications for the extent to which households are credit constrained in equilibrium. While households become more close to the borrowing constraint when income inequality comes from households' lowest fixed level of human capital, the opposite occurs when the source of income inequality is the variance of the idiosyncratic risk. Also, the reduction of the precautionary savings due to a credit deepening is higher in economies with income inequality driven by households' lowest fixed level of human capital. For this reason, consumption per capita peak growth is higher. However, the opposite occurs when the source of income inequality is the variance of the idiosyncratic risk.

This paper is organized as follows. Section 2.2 presents the empirical evidence, including the related literature and data description. In section 2.3, we describe our model and include the description of our experiment. In section

2.4, we present our calibration strategy and our main results. Finally, section 2.5 concludes.

2.2 Empirical Evidence

2.2.1 Related literature

Our work is related to a vast literature that has been studying the relation between initial inequality, typically measured by the Gini index, and subsequent economic growth. Alesina and Rodrik (1994) and Persson and Tabellini (1994) found a negative association through a cross-country estimation. Whereas Forbes (2000), estimating a panel with fixed effects, found a positive relationship between initial inequality and subsequent growth in the short run and negative in the long run.

Moreover, there is a literature that has focused on the role of poverty on economic growth. Ravallion (2012) evaluates the relationship between growth (measured by consumption) and initial poverty, using a new database for a set of developing countries. His results suggest that there is an adverse effect of high initial poverty on growth and high initial poverty dulls the impact of growth on poverty. Ravallion also shows that high initial inequality matters to growth and poverty reduction if it entails a high incidence of poverty relative to the mean.

Another strand of the literature that has incorporated credit-market failures suggests that high inequality reduces economy's aggregate efficiency and, therefore, it reduces growth. For example, Banerjee and Newman (1993) and Aghion and Bolton (1997) found that inequality lead to lower economic growth due to credit-market imperfections. They argued that in the short run the relationship might be positive, but in the long run, more income inequality hampered economic growth. Galor and Zeira (1993) show, using a theoretical model, that in the presence of credit-market failures and indivisibilities in investment in human capital, the initial distribution of wealth affects aggregate output and investment, both in the short and in the long run, as there are multiple steady states. Finally, Ravallion (2001, 2007) argues intuitively that poverty retards growth when there are credit-market failures.

There has been considerable interest among economists in the idea that, for various reasons, inequality could actually reduce growth. Recently, this line of inquiry was brought into focus by the latest financial crisis, which has led to a vigorous debate if widening inequality was one of the causes of the crisis. For example, Pegurini et al. (2013), using a panel estimation, finds a statistically significant and positive relationship betweeen income concentration and private sector indebtedness, controlling for conventional credit determinants. This result calls attention to the importance of the distribution of income to macroeconomics outcomes.

Unlike the literature mentioned, the purpose of this paper is to analyze if income inequality plays a role in the association between consumption growth and credit expansion. For such purpose, we use cross-country and panel estimations. Our results show that the association between consumption growth and credit expansion is stronger in countries with higher income inequality.

2.2.2
Data

We use annual data. Our dependent variable is the average growth of household consumption per capita, which is available in the World Bank database. Household final consumption expenditure per capita (private consumption per capita) is calculated using private consumption in constant 2005 prices and World Bank population estimates.

We use the Gini index to measure income or consumption inequality in each country, it measures the extent to which the distribution of income or consumption expenditure among households within an economy deviates from a perfectly equal distribution.[2] We choose to use two datasets of Gini indexes.[3] The first one is constructed by Ravallion (2012). The autor calculates income inequality using directly primary data from household surveys for 92 developing and transition countries. This method eliminates some of the inconsistencies and comparability problems found in existing data compilations from secondary sources. The other dataset is a compilation made by the United Nations University - World Institute for Development Economics Research (UNU-WIDER) that ranks available Gini indexes by the quality of their sources and methods.

At last, for credit data, we use two datasets. The first one is from the Financial Access Survey (FAS) developed by the International Monetary Fund (IMF). This survey collects annual data on access to and the use of financial services around the world, from 2004 to 2011. We consider only household

[2]A Gini index of 0 represents perfect equality, while an index of 100 implies perfect inequality.

[3]Unlike national accounts data which are in principle comparable across countries, there is no agreed basis of definition for the construction of distribution data. Sources and methods might vary, especially across but also within countries. This may be the case even if the data comes from the same source.

outstanding loans with commercial banks. The other database is from the World Bank, which is the domestic credit to private sector, and is bigger than the IMF database. For this reason, we use the domestic credit to private sector to construct the panel data.

2.2.3 Results

In this section we provide evidence on the link between the average consumption per capita growth and the interaction between initial inequality and the variation of credit. We estimate the following regression:

$$Cgrowth_{t:T} = \alpha + \beta_1 \ln(c_t) + \beta_2 gini_t + \beta_3 credit_t + \beta_4 \Delta_{T-t} credit + \beta_5 gini \Delta_{T-t} credit + \varepsilon_t \tag{2-1}$$

The dependent variable $Cgrowth_{t:T} = \frac{100[\ln(c_T) - \ln(c_t)]}{T-t}$ is the average consumption per capita growth rate between years t and T. We follow an extensive literature of economic growth to choose our controls, especially, Alesina and Rodrik (1994) and Person and Tabellini (1994): $gini_t$ is the Gini index for year t, $ln(c_t)$ is the consumption per capita in year t and $credit_t$ is the initial credit to GDP ratio. The main difference from other studies is that we consider the variation of credit between years t and T ($\Delta_{T-t}credit$) and the interaction between $\Delta_{T-t}credit$ and $gini_t$, because we are interested to show that inequality plays a roll in the relation between consumption growth and the variation of credit.

Table 2.2 shows our results. In the first two columns, we consider the Gini index constructed by Ravalion (2012) and the credit to households from the IMF. In this case, our sample consists of thirty one developing countries and the initial and final years are 2004 and 2007, respectively. In column 1, we do not include the interaction $gini\Delta_{T-t}credit$, hence the coefficient of the variation of credit (β_4) is positive and significant at 1 percent. On the other hand, in column 2, we introduce the interaction and its coefficient (β_5) is positive and significant at 5 percent, but the coefficient of the variation of credit becomes negative and insignificant. This result suggests that initial income inequality of countries in 2004 intensifies the association between the variation of credit and the average consumption per capita growth between 2004 and 2007.[4] The average Gini index in this sample is 39.25, thus the correlation estimated between credit expansion and consumption growth is approximately 0.57, all other variables remain constant. This is our main evidence as we use the best source of Gini index and household debt data.

[4]Note that we are not interested in the size of the estimated coefficients. Our interest lies only on their signals.

In addition, we repeat the same regressions in columns 1 and 2 using different databases and present a panel specification. In columns 3 and 4, we use the Gini index calculated by UNU-WIDER and the credit to households from the IMF. Our sample consists of fifty countries and we consider the same years as before (2004 and 2007). Note that the results persist: the coefficient of the interaction between $\Delta_{T-t}credit$ and $gini_t$ is positive and significant at 5 percent and the coefficient of the variation of credit becomes negative and insignificant once the interaction is considered.

In columns 5 and 6, we also consider the Gini index from UNU-WIDER but we use the domestic credit to the private sector from the World Bank, which is a good proxy for household credit, the correlation coefficient between these series is greater than 0.7. This occurs because household credit is in the private sector credit. In this case, we have 93 countries and the initial and final years are 2000 and 2007, respectively. Most of the results are maintained, but the coefficient of the variation of credit is negative and significant at 10 percent once the interaction $gini\Delta_{T-t}credit$ is considered.

Furthermore, we show that the evidence presented above persists once we use a panel estimation with fixed-effects for years and countries. These results are in columns 7 and 8. Our sample consists of thirty one countries and we consider four windows of five years between 1990 and 2009. In this specification, we also use the Gini index calculated by UNU-WIDER and the domestic credit to private sector from the World Bank. Note that the coefficient of the interaction between $\Delta_{T-t}credit$ and $gini_t$ is positive and significant at 10 percent, but it is smaller than the cross-section estimates.

In conclusion, we find evidences that the association between average consumption per capita growth and the variation of credit is stronger in countries with higher income inequality. In the following section, we provide a theoretical framework and we test if this model can corroborate the empirical finding.

	Ravallion (2012)		UNU-WIDER					
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
VARIABLES	04-07	04-07	04-07	04-07	00-07	00-07	90-09	90-09
ginivarcreditHH		**0.0350****		**0.0180****				
		(0.0151)		**(0.00794)**				
ginivarcreditDO						**0.00280*****		**0.00157***
						(0.00106)		**(0.000828)**
consupinitial	-0.812	-0.524	-1.347***	-1.154**	-0.666**	-0.588**	-8.849**	-8.976**
	(0.532)	(0.531)	(0.487)	(0.506)	(0.294)	(0.295)	(3.589)	(3.452)
giniinitial	0.0533	-0.0823	0.0741	0.0203	-0.0137	-0.0420	0.0339	0.0196
	(0.0639)	(0.0685)	(0.0597)	(0.0554)	(0.0244)	(0.0256)	(0.0492)	(0.0448)
credGDPinitialHH	-0.0392	-0.0519	-0.0429	-0.0424				
	(0.0718)	(0.0690)	(0.0328)	(0.0341)				
varcreditHH	0.548***	-0.804	0.328***	-0.271				
	(0.0861)	(0.619)	(0.111)	(0.213)				
credGDPinitialDO					-0.0120*	-0.0128*	-0.0159**	-0.0119*
					(0.00679)	(0.00681)	(0.00749)	(0.00700)
varcreditDO					0.0376***	-0.0621*	0.00955	-0.0516
					(0.0130)	(0.0345)	(0.0103)	(0.0310)
Constant	6.937	10.80**	11.23**	11.54**	9.129***	9.796***	72.59**	73.83**
	(4.188)	(4.606)	(4.661)	(4.379)	(2.720)	(2.824)	(29.90)	(28.70)
Observations	31	31	50	50	93	93	124	124
R-squared	0.493	0.562	0.290	0.319	0.242	0.285	0.294	0.317
F test model	11.80	24.84	7.458	8.625	5.563	4.989	3.276	3.127
Number of countries							31	31

Robust standard errors in parentheses: *** $p<0.01$, ** $p<0.05$, * $p<0.1$

Table 2.2: Empirical Evidence

2.3 Model

2.3.1 The baseline model

Our model is based on Aiyagari (1994) and it is related to papers that consider incomplete markets with heterogeneous agents, idiosyncratic risk and borrowing constraint, such as Aiyagari (1994) and Bewley (1977). Our aim is to study the consumption response to a credit deepening in economies with different levels of income inequality.

Recently, Guerrieri and Lorenzoni (2010), using a model à la Bewley-Huggett, evaluate the effects of a credit tightening on consumer spending and aggregate output. Their results show that overly indebted agents reduce their spending and increase their labor supply. But the precautionary motive is strong enough that the reduction in consumer spending dominates. As a result, there is a decrease in the aggregate output that causes recession.

Endowments

There is a continuum with unit mass of households which live infinitely and faces idiosyncratic income uncertainty. Households differ in their labor endowments. There are two channels that contribute to determinate the total efficiency units supplied by the households to the labor market. First, there is a stochastic component s_t that is the efficiency units of labor whose log follows a finite state Markov chain, which is approximated by a stationary AR(1) process $\log(s_t) = \rho \log(s_{t-1}) + \varepsilon_t$ with $\varepsilon_t \sim N(0, \sigma_\varepsilon^2)$. Second, there is a fixed component $\theta \in \{\theta_1, \theta_2\}$, that is the household level of human capital drawn by nature from a distribution in which each θ has mass μ_θ and $\sum_\theta \mu_\theta = 1$.

The total efficiency units that a household is endowed is the product of these two components, and labor earnings are given by $w\theta s_t$, where w is the wage per efficiency unit.

Preferences

Preferences are described by:

$$E_0 \sum_{t=0}^{\infty} \beta^t U(c_t) \tag{2-2}$$

The future is discounted at rate $\beta \in (0,1)$. We assume that $U(c_t) = \ln(c_t)$.

Production technology

There is a representative firm that produces consumption goods with a Cobb-Douglas function:

$$f(k_t, n_t) = k_t^\alpha n_t^{1-\alpha} \tag{2-3}$$

where $\alpha \in (0,1)$, k_t and n_t are aggregate capital and efficient labor units, respectively.

At each period, capital depreciates at the exogenous rate δ and the firm hires capital and labor from competitive markets.

The problem of the representative firm is given by:

$$\max_{n_t, k_t} f(k_t, n_t) - w_t n_t - (r_t + \delta) k_t$$

The first order conditions of the firm provide the expressions for the net real return of capital r_t and the wage rate per efficiency unit w_t:

$$f_k(k_t, n_t) - \delta = r_t \tag{2-4}$$

$$f_n(k_t, n_t) = w_t \tag{2-5}$$

Other market arrangements

There are no insurance markets for the idiosyncratic shock, but households can self-insure by saving through the risk-free asset. Moreover, there is an exogenous borrowing limit, $\underline{b} > 0$, which is tighter than the natural limit. Hence, in our model, markets are incomplete in the sense that households can smooth consumption by borrowing and lending only through one type of asset.

We also consider a small open economy in the financial market and a closed one in the labor market. Thus, at each period, the interest rate r is exogenously determined in the international capital market, but the wage rate w clears the domestic labor market.

2.3.2
Stationary equilibrium

Our focus is on the properties of a stationary competitive equilibrium in which the measure of agents, defined over an appropriate family of subsets of the individual state space, remains invariant over time.

Household problem

The value function of a household with current asset holdings a, current labor endowment s and fixed level of human capital θ is denoted by $V(a, s; \theta)$. The household problem in its recursive form can be represented as follow:

$$\begin{aligned} V(a, s; \theta) &= \max_{c,a'} \{log(c) + \beta \sum_{s'} V(a', s'; \theta)\Pi(s', s)\} \\ &\text{subject to} \\ c + a' &= (1 + r)a + w\theta s \\ c \geqslant 0, &\qquad a' \geqslant -\underline{b} \end{aligned}$$

The first restriction is the budget constraint. The last set of restrictions implies that consumption is feasible and households are borrowing constrained. As denoted before, s is an uninsured idiosyncratic risk that follows a finite state Markov chain with transition probabilities $\Pi(s', s) = Prob(s_{t+1} = s'|s_t = s)$, where $s, s' \in \{s_1, s_2, ..., s_n\}$.

We obtain the decision rules for consumption $c(a, s; \theta)$ and asset holdings $a(a, s; \theta)$ by solving the problem above.

Recursive Equilibrium

The definition of stationary competitive equilibrium is given by $[V(a,s;\theta),\ c(a,s;\theta),\ a(a,s;\theta),\ r,\ w,\ k,\ n,\ T,\ \tau]$ and a distribution $\lambda(a,s;\theta)$ such that:

(i) Given prices (r,w), $[V(a,s;\theta), c(a,s;\theta), a(a,s;\theta)]$ are the solutions of the agent's problem.[5]

(ii) The wage per efficient unit and the price of capital are given by:

$$f_n\left(k,n\right) = w \tag{2-6}$$

$$\delta = f_k\left(k,n\right) - r \tag{2-7}$$

(iii) $\lambda(a,s;\theta)$ is a stationary distribution associated with the transition function implied by the decision of $a(a,s;\theta)$ and the law of motion s:

$$\lambda(a,s;\theta) = \int_{\mathbf{A}\times\mathbf{S}} P(a,s,\mathbf{a},\mathbf{s};\theta)d\lambda \tag{2-8}$$

for all $\mathbf{a}\times\mathbf{s}\subseteq\mathbf{A}\times\mathbf{S}$. The transition function P is the probability that a household with state (a,s) will have a state belonging to $\mathbf{a}\times\mathbf{s}$ next period.

(iv) Labor market clears:

$$n = \sum_{\theta}\mu_\theta \int_{\mathbf{A}\times\mathbf{S}} \theta s\lambda(a,s;\theta)dads \tag{2-9}$$

2.3.3
Description of the experiment

We consider a credit deepening as an unexpected shock that increases permanently the borrowing limit $\underline{b}$ to $\underline{b}'$. Note that there are two sources of income inequality in our model: variance of the idiosyncratic risk σ_ε^2 and households' fixed level of human capital θ.[6] We are interested in analyzing the consumption response to a credit deepening in economies with different levels of income inequality measured by the Gini index. Therefore, we consider a benchmark economy and four alternative economies: two with higher income inequality and the other two with lower inequality than the benchmark. We

[5]We use the endogenous grid point method proposed by Carroll (2006) to solve the household problem.

[6]We tried to incorporate this sources of income inequality on the empirical exercise in spite of using only the Gini index, but we were not able to find data that correspond to each inequality source in our model.

consider extreme values of income inequality in order to check if our results are monotonic. In our experiment, we analyze the consumption growth at the peak.

Since we have two different sources of income inequality, we change only one source at each experiment and we set the other source at the benchmark level. For the variance of the idiosyncratic risk (σ_ε^2), we consider different values of σ_ε^2 to generate economies with lower and higher income inequality than the benchmark. Thus, we modify the grid of possible values of s_t to generate a mean preserving spread in the distribution of labor efficient units, so the aggregate labor units (n) are not changed.

In addition, we perform another exercise to understand the mechanisms behind our model. When we generate economies with different levels of income inequality, there are two effects. First, the policy functions (consumption and assets) are changed and second, the asset distributions that come from these policies also change. Then we calculate the contribution of the policy functions and the assets distribution in the consumption growth driven by a credit deepening. In order to calculate the contribution of the policy function, we keep assets distribution as in the benchmark economy and calculate the consumption peak growth under this scenario. Concerning the contribution of the assets distribution, we keep the same policy functions as in the benchmark economy and calculate the consumption growth at the peak when the assets distribution is changed from the benchmark to another distribution, which corresponds to a different level of income inequality.

2.4 Quantitative analysis

We analyze the consumption growth driven by a credit deepening in economies that have different levels of income inequality. There are two sources of income inequality in our model:[7] the variance of the idiosyncratic risk σ_ε^2 and the households' fixed level of human capital θ. In this section, we present the results for these two cases.

[7]Note that ρ is another source of inequality in the model, but changes in this parameter modify the transition probabilities. We want to compare economies with different levels of inequality and with the same transition probabilities.

2.4.1 Calibration

Each period corresponds to one year. We calibrate our model for the US economy,[8] considering standard parameters in the literature therefore our results would not be biased by the calibration.

We solve the stationary recursive equilibrium of the model and the transitional dynamics numerically using the algorithm of Ríos-Rull (1999). We consider that the log of the efficiency units of labor s_t follows a finite state Markov chain, which is approximated by a stationary AR(1) process $\log(s_t) = \rho \log(s_{t-1}) + \varepsilon_t$ with $\varepsilon_t \sim N(0, \sigma_\varepsilon^2)$. We use Rouwenhorst (1995)'s algorithm with 3 states to approximate this AR(1) process using a Markov chain.[9]

We choose the parameters ρ and variance σ_ε^2 in line with the evidence found in Floden and Lindé (2001) that use yearly panel data from PSID to calculate these parameters for the US economy. We set the persistence parameter ρ to 0.96 and the variance of the idiosyncratic risk σ_ε^2 is set to 0.0441. Thus, our calibrated model has an income distribution with a Gini index equals to 42.

We consider a Cobb Douglas production function. The share of capital α is equal to 0.36, which is a common value for the US economy (e.g. Ayagari (1994)). We calibrate the interest rate $r = 0.03$, which is in line with US data. Furthermore, we follow Aiyagari (1994) to calibrate the discount factor $\beta = 0.96$.

We follow Erosa and Ventura (2002) to approximate households' fixed level of human capital θ. These parameters are calculated using data from the US Census Bureau. The population is divided in two groups according to education levels and their labor earnings are computed. The parameters θ_1 and θ_2 are approximated by the ratio of labor earnings between the high labor earnings group with respect to low labor earnings. Thus, $\theta_1 = 1$ and $\theta_2 = 1.84$ with share of households in each group $\mu_{\theta_1} = 0.69$ and $\mu_{\theta_2} = 0.31$, respectively.

Finally, we set the borrowing limit $\underline{b}$ to the wage per efficiency unit $w = 1.18$ in our calibrated economy. This value corresponds to an initial credit-to-GDP ratio of 8 percent, which we compute using the sum of households debt divided by the product in our model.

Table 2.3 summarizes this information.

[8] In our model, we consider an open economy. One may argue that this hypothesis is not valid for the US, but the mechanisms behind our results would be the same in a closed economy.

[9] We choose this method because it is superior to the commonly used Tauchen (1986) procedure as it perfectly matches persistence of the process even for low number of states.

Parameter	Description	Value	Sources
α	Capital Share	0.36	Aiyagari (1994)
β	Rate of time preference	0.96	Aiyagari (1994)
σ_{ε}^2,ρ	Variance and Persistence	$0.0441; 0.96$	Floden and Lindé (2001)
$\underline{b}$	Borrowing Limit	$w = 1.18$	-
r	Interest rate	0.03	U.S data
θ_1, θ_2	Households' fixed level of human capital	$\{1; 1.84\}$	Erosa and Ventura (2002)

Table 2.3: Benchmark Economy: US.

2.4.2
Credit deepening

We explore the response of our economy to a credit deepening. We consider an economy that at $t = 0$ is in steady state with the borrowing limit $\underline{b} = w = 1.18$. We then look at the effects of an unexpected shock at $t = 1$ that permanently rises the borrowing limit to $\underline{b}' = 2w = 2.36$.

Figure 2.1 shows the optimal values of consumption and assets as a function of households' assets holdings (a) with $\theta_1 = 1$ and efficiency units of labor equals to $s = s_2$ in two steady states: before and after the credit shock.[10]

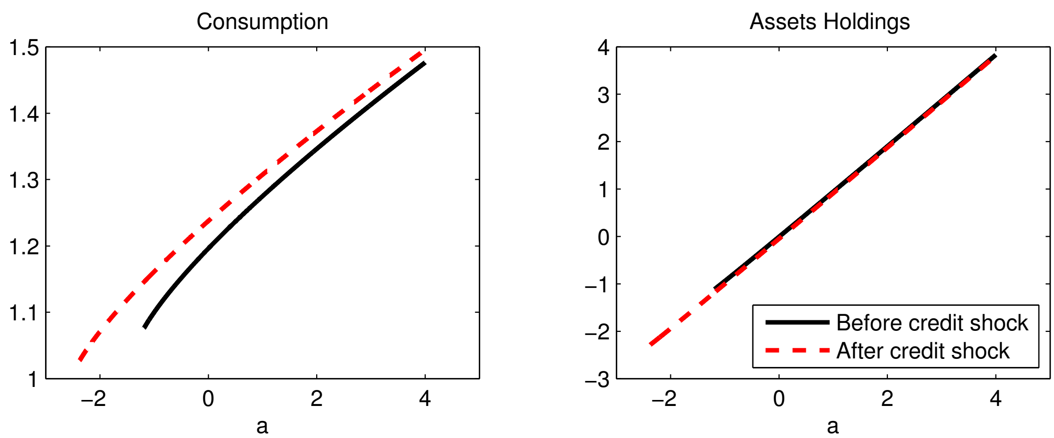


Figure 2.1: Optimal consumption and assets holdings at $s = s_2$ and $\theta_1 = 1$ - Benchmark Economy.

At high levels of a, household behavior is close to the permanent income hypothesis and the consumption function is almost linear in a. For lower levels of assets holdings, the consumption function is concave, as it is common in

[10]We choose this type of household as an example, but the interpretation in this section is valid for all other types and levels of efficiency units.

precautionary savings models.[11] Also, the optimal values of assets holdings increase as a increases.

After the credit shock, consumption is higher for all levels of assets holdings. Besides, their savings are reduced for all levels of a because an increase in the borrowing limit reduces the precautionary motive in the economy.

Note that consumption response differs with the level of assets holdings. For households close to the borrowing constraint, precautionary motive is higher. Then, their consumption response to the credit shock is higher than households with high level of assets holdings.

2.4.3 Households' fixed level of human capital θ

One source of income inequality in our model is the households' fixed level of human capital, $\theta \in \{\theta_1, \theta_2\}$. In this section, we analyze the consumption response to a credit deepening in economies with different levels of income inequality driven by each type of θ.[12]

Households' lowest fixed level of human capital θ_1

We change θ_1 in order to generate four economies with different levels of income inequality relative to the benchmark (Gini index of 42). Therefore, a higher (lower) value of θ_1 reduces (increases) income inequality. Particularly, we set θ_1 equals to 0.7 and 0.5 to generate economies with a Gini index of 47 and 52, respectively. On the other hand, to generate economies with less income inequality than the benchmark, we set θ_1 equals to 1.5 and 2 which correspond to a Gini index of 32 and 37, respectively.

Figure 2.2 shows the results of this experiment which are in line with the empirical evidence found: the response of consumption per capita to a credit deepening is higher in economies with higher income inequality. Note that in the economy with the highest income inequality (the lowest θ_1) than the benchmark, consumption per capita grows at the peak 2.06 percent. However, in the other one with the lowest income inequality (the highest θ_1), consumption per capita at the peak grows 1.37 percent.

[11]See Carroll and Kimball (1996).

[12]In the appendix, we show a mean preserving spread exercise for this source of income inequality.

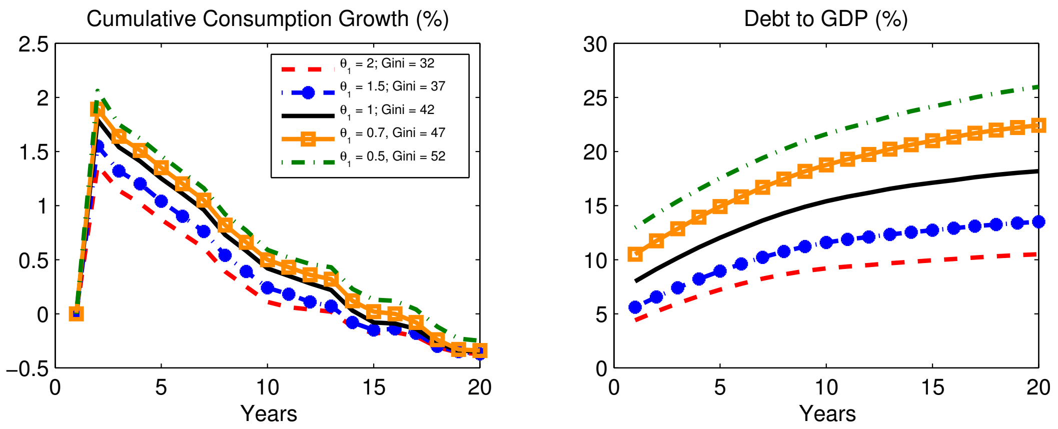


Figure 2.2: θ_1: Cumulative consumption growth ad Debt to GDP by income inequality.

Table 2.4 presents the contributions of the policy functions and the assets distribution on consumption growth at the peak. For the first case, we consider the same assets distribution as in the benchmark and then calculate consumption growth at the peak. Our results suggest that consumption growth at the peak for economies with higher inequality than the benchmark is higher, even when we consider the same assets distribution as in the benchmark. For example, in the economy with Gini index of 47 and 52, consumption growth at the peak is 1.85 and 1.93 percent, respectively, which is higher than the benchmark growth of 1.79 percent. The opposite occurs when we consider the economies with lower income inequality than the benchmark. Hence, the reduction of the precautionary motive due to a credit deepening is higher in economies with higher income inequality driven by lower θ_1, as a result the consumption growth is also higher.

Consumption growth at peak (%)				
θ_1	**Gini Index**	**Total**	**Policy**	**Distribution**
2	32	1.37	1.76	-0.20
1.5	37	1.55	1.69	-0.32
1	**42**	**1.79**	**-**	**-**
0.7	47	1.89	1.85	0.03
0.5	52	2.06	1.93	0.14

Table 2.4: θ_1: Policy functions and Assets distribution.

For the second case, we analyze the contribution of the assets distribution. For such objective, we consider the same policy functions as in the benchmark economy. Thus, we calculate the consumption growth at the peak when the assets distribution is changed from the benchmark to another distribution that corresponds to the alternative levels of income inequality. Our results shows that the contribution of assets distribution for economies with Gini index of 47 and 52 is 0.03 and 0.14 percent, respectively. On the other hand, there is a negative contribution of assets distribution for economies with lower income inequality than the benchmark. This means that a decrease (increase) on the fixed level of human capital θ_1 makes this group of households poorer (richer) than in the benchmark. For this reason, their assets policy, in the steady state, lower (higher) than the same group in the benchmark, which is shown in Figure 2.3 for the economies with the highest and the lowest income inequality in our exercise and for a household with $s_t = s_2$ and $\theta = \theta_1$. Consequently, there are more (less) households close to the borrowing constraint for this group. Figure 2.4 corroborates this result by reporting the assets distribution for the economies according to their income inequality. The higher the income inequality, the less the assets distribution is concentrated in the higher levels of assets.

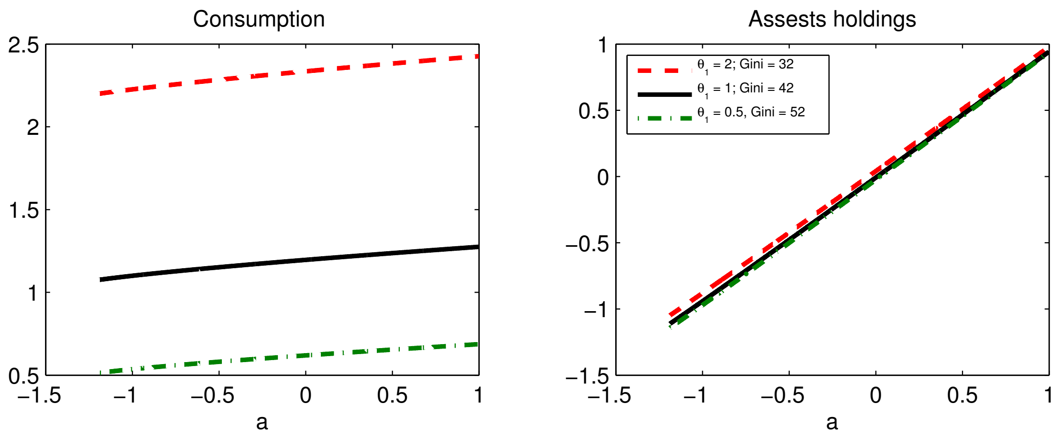


Figure 2.3: θ_1: Optimal consumption and assets holdings at $s = s_2$ by income inequality.

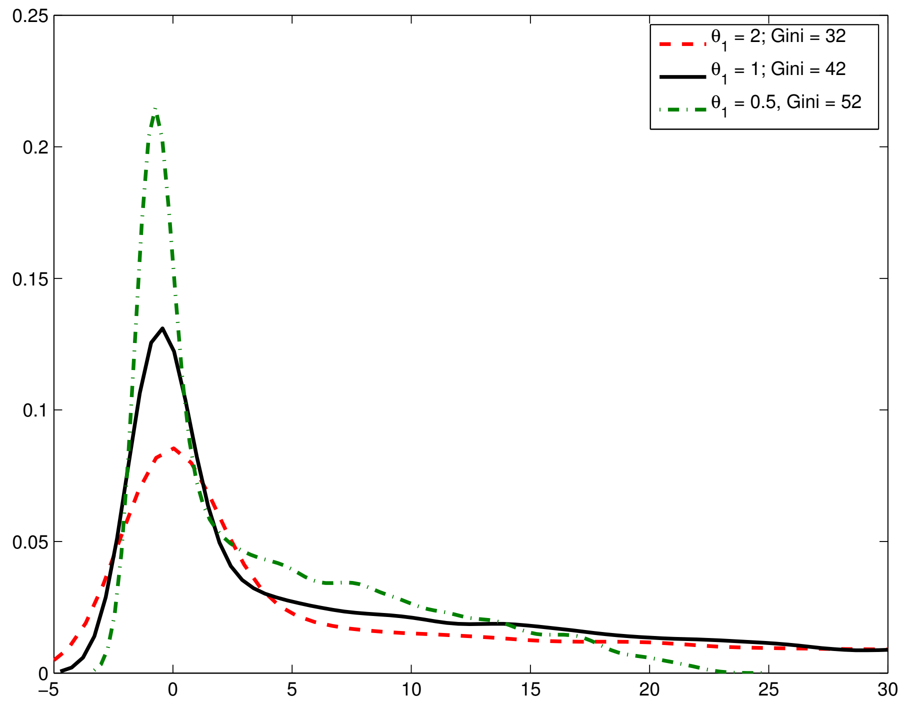


Figure 2.4: θ_1: Marginal density of assets holdings by income inequality.

In conclusion, our theoretical model replicates the empirical association found once we consider that income inequality is driven by θ_1.

Households' highest fixed level of human capital θ_2

In this section, we repeat both performed exercises for θ_1 considering θ_2 as the source of income inequality. Hence, a lower (higher) value of θ_2 reduces (increases) income inequality. At first, we set θ_2 equals to 3 and 4 to generate economies with a Gini index of 47 and 52, respectively. Whereas to generate economies with less income inequality than the benchmark, we set θ_2 equals to 1.2 and 0.9 which correspond to a Gini index of 32 and 37, respectively.

Figure 2.5 shows that the results for this exercise are not in line with the empirical finding. Note that consumption growth at peak is higher in economies with lower income inequality. For example, the economy with Gini index of 32 shows a consumption growth at the peak of 2.21 percent, while the other one with Gini index of 37 is 2.03 percent. On the other hand, consumption growth at peak is 1.47 and 1.24 in the economies with Gini index of 47 and 52, respectively.

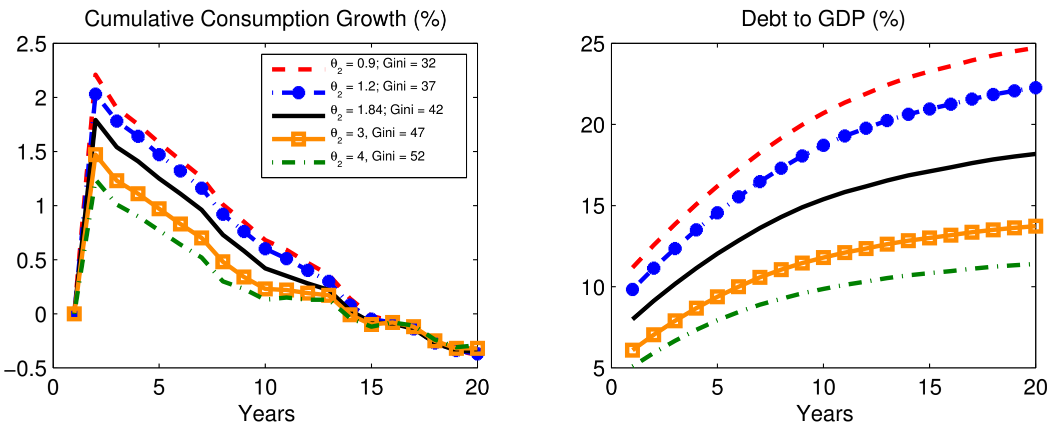


Figure 2.5: θ_2: Cumulative consumption growth ad Debt to GDP by income inequality.

Table 2.5 shows the contributions of the policy functions and the assets distribution on consumption growth at the peak. Concerning the contribution of the policy functions, consumption growth at the peak for economies with higher inequality than the benchmark is lower. For example, in the economy with Gini index of 47 and 52, consumption growth at the peak is 1.63 and 1.49 percent, respectively. The opposite occurs when we consider the economies with lower income inequality than the benchmark. This means that the reduction of the precautionary motive due to a credit deepening is lower in economies with higher income inequality driven by higher θ_2, so the consumption growth is also lower. Concerning the contribution of the assets distribution, there is a positive contribution for economies with lower income inequality and a negative one for economies with higher income inequality than the benchmark. This occurs because the assets policy functions are, in equilibrium, higher (lower) in economies with higher (lower) income inequality than the benchmark, which is shown in Figure 2.6 for economies with the highest and the lowest Gini index in our exercises. Therefore, there are more households close to the borrowing constraint in economies with low income inequality and the opposite happens for the ones with high income inequality driven by θ_2, as shown in Figure 2.7.

Consumption growth at peak (%)				
θ_2	**Gini Index**	**Total**	**Policy**	**Distribution**
0.9	32	2.21	1.93	0.27
1.2	37	2.03	1.89	0.14
1.84	**42**	**1.79**	**-**	**-**
3	47	1.47	1.63	-0.16
4	52	1.24	1.49	-0.25

Table 2.5: θ_2: Policy functions and Assets distribution.

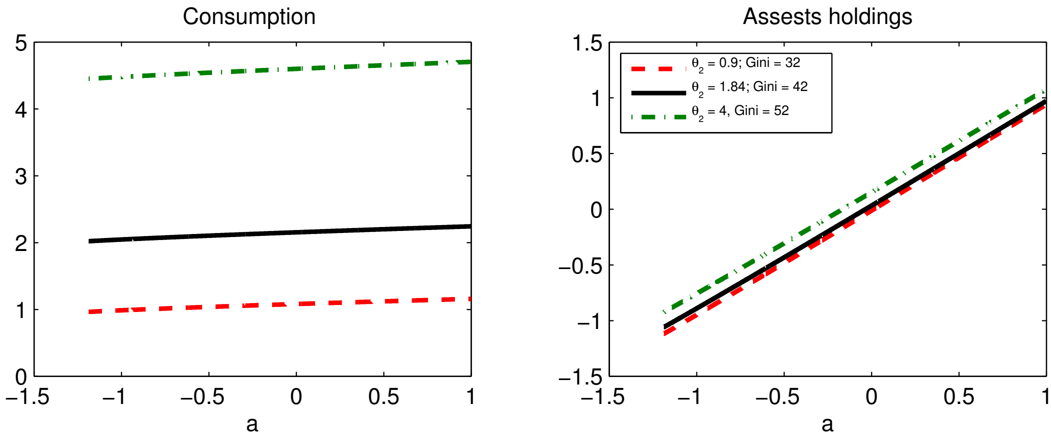


Figure 2.6: θ_2: Optimal consumption and assets holdings at $s = s_2$ by income inequality.

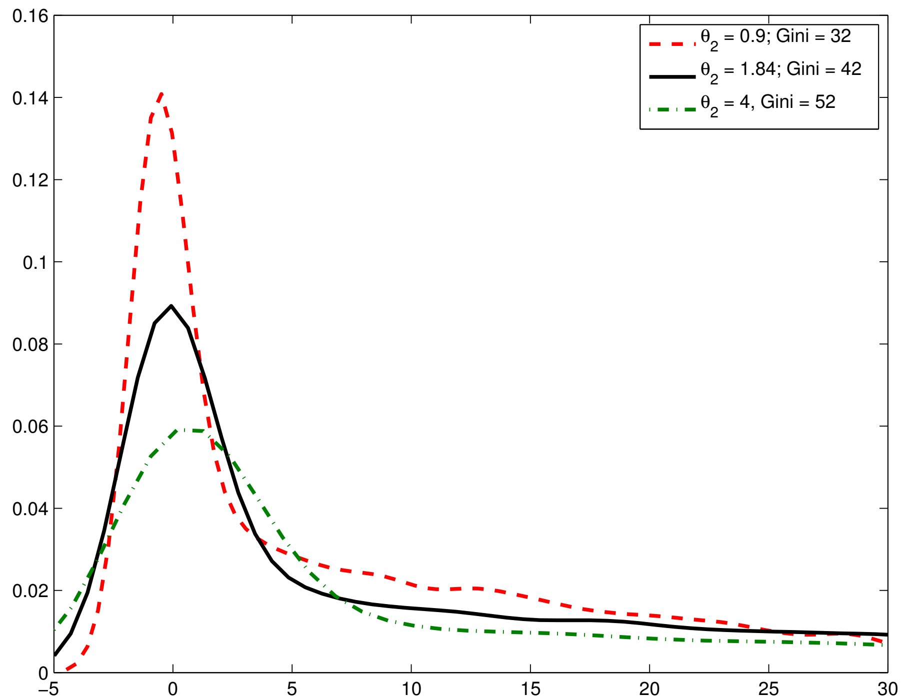


Figure 2.7: θ_2:: Marginal density of assets holdings for households with by income inequality.

Finally, when we consider economies with income inequality driven by θ_2, our model does not rationalize the empirical finding.

2.4.4 Variance of the idiosyncratic risk σ_ε^2

One of the sources of inequality inherent in our model is given by the variance of the idiosyncratic risk, σ_ε^2. This parameter determines the possible values of efficiency units of labor. The effect of a higher variance of the idiosyncratic risk on income inequality is twofold. First, a higher σ_ε^2 implies that the grid of possible values of s_t has more extreme values,[13] as a result income inequality increases. Second, a higher σ_ε^2 increases the precautionary motive in the economy, then households increase their savings. Mainly, households close to the borrowing constraint increase their savings more than the others as their precautionary motive is higher, consequently, income inequality decreases. We show that the first effect dominates.

We calibrate four different economies with different levels of income inequality relative to the benchmark and, for each case, we consider a mean preserving spread by adjusting the possible values of s_t to generate the same aggregate labor units as in the benchmark economy. Specifically, we set σ_ε^2

[13]In Rouwenhorst (1995)'s algorithm, changes in σ_ε^2 modify the values of the states s_t but the matrix of transition remains the same.

equals to 0.0941 and 0.1410 to generate economies with a Gini index of 47 and 52, respectively. In order to generate economies with less income inequality than the benchmark (42), we set σ_{ε}^2 equals to 0.0141 and 0.0241 that correspond to a Gini index of 32 and 37, respectively. After these economies are calibrated, we analyze the effect of a credit deepening on consumption per capita. Figure 2.8 presents the results for this experiment.

These results show that in economies with higher inequality driven by more uncertainty on households' income, a credit deepening has a lower effect on consumption per capita than in an economy with lower inequality, driven by less uncertainty. In order to understand the mechanisms behind these results, we quantify the contributions of the assets distribution and the policy functions for each case. Table 2.6 presents these exercises.

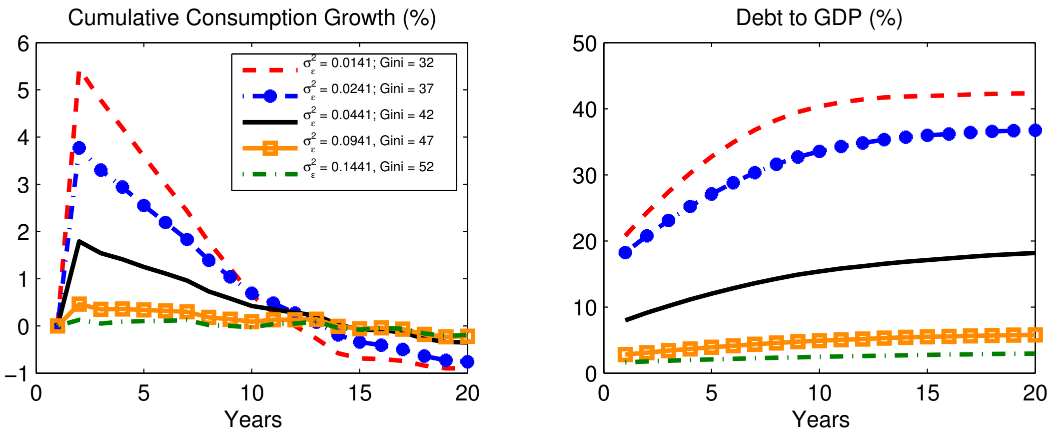


Figure 2.8: σ_{ε}^2: Cumulative consumption growth ad Debt to GDP by income inequality.

Consumption growth at peak (%)				
σ_{ε}^2	**Gini Index**	**Total**	**Policy**	**Distribution**
0.0141	32	5.45	2.19	3.19
0.0241	37	3.77	1.95	1.78
0.0441	**42**	**1.79**	**-**	**-**
0.0941	47	0.46	1.35	-0.88
0.1410	52	0.13	0.91	-0.77

Table 2.6: σ_{ε}^2: Policy functions and Assets distribution.

To calculate the contribution of the assets distribution, we keep the same policy functions as in the benchmark economy. Then, we calculate the consumption growth at the peak when the assets distribution is changed from the benchmark to another distribution that corresponds to the alternative levels of income inequality. The contribution of assets distribution for economies with Gini index of 47 and 52 is -0.88 and -0.77 percent, respectively. The opposite occurs for economies with low income inequality. This happens because, in our model, in an economy with high income uncertainty, households increase their precautionary savings, especially the ones close to the borrowing constraint. Therefore, in equilibrium, there are less households close to the borrowing constraint than in another economy with low income uncertainty. Figure 2.9 corroborates this result by reporting the marginal density of assets holdings for the economies according to their income inequality. The higher the income inequality, the more the assets distribution is concentrated in higher levels of assets.

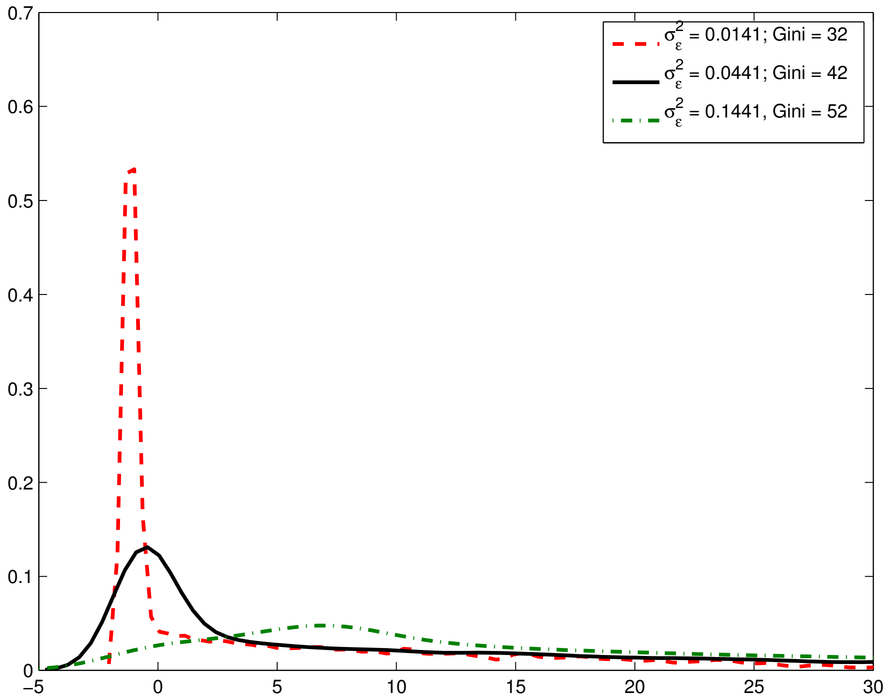


Figure 2.9: σ^2_ε: Marginal density of assets holdings by income inequality.

Note that the pattern of the assets distribution is due to the policies functions. Figure 2.10 shows the consumption and assets policy functions before the credit deepening for the benchmark and for the economies with highest and lowest income inequality in our exercise, considering a household with $s_t = s_2$ and $\theta = \theta_1$. The assets policy function in the economy with higher income uncertainty (i.e., higher income inequality) is above all others and the opposite is valid for the consumption policy function. Therefore, assets accumulation is higher in economies with higher uncertainty of income. This result is in line with Huggett (2003). The author shows, using an incomplete markets framework, that given two economies with the same borrowing limit, but with different earnings processes, the one with the riskier earning process has more assets accumulation.

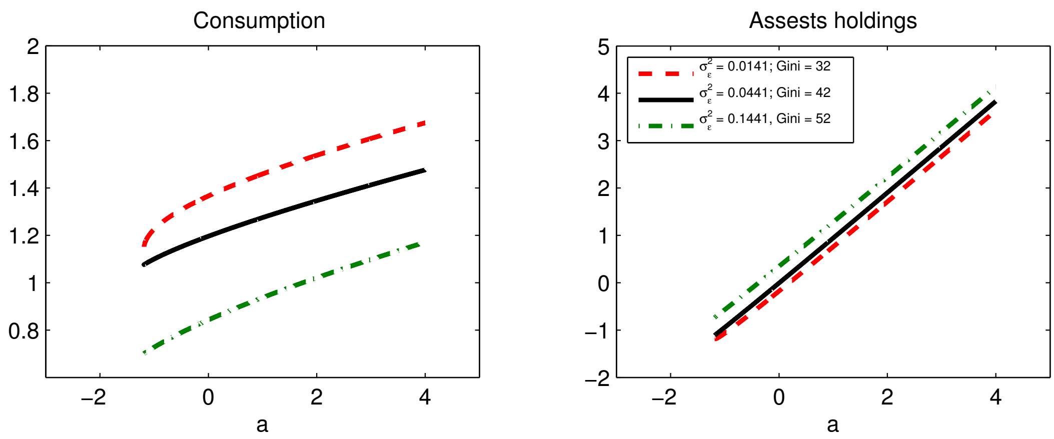


Figure 2.10: σ^2_ε: Optimal consumption and assets holdings at $s = s_2$ and $\theta_1 = 1$ by income inequality.

Concerning the contribution of the policy function, consumption growth at the peak for economies with higher inequality than the benchmark is lower, even when we consider the same assets distribution as in the benchmark. For example, in the economy with Gini index of 47 and 52, consumption growth at the peak is 1.35 and 0.91 percent, respectively, which is lower than the benchmark growth of 1.79 percent. The opposite occurs when we consider the economies with less income inequality than the benchmark. Hence, considering the assets distribution fixed, consumption growth at the peak is higher in economies with higher income inequality driven by income uncertainty. This means that the reduction of the precautionary motive due to a credit deepening is smaller in economies with higher income inequality due to higher income uncertainty, as a result the consumption growth is also smaller. The intuition behind this result is the fact that, in an economy with high income uncertainty, households have more incentives to do not stay close to the borrowing constraint, because their precautionary motive is higher.

Furthermore, these results are robust to changes in $\underline{b}$. Note that, once the borrowing limit increases, our model approximates to a complete market framework. Therefore, the consumption policies in economies with different levels of income inequality become more linear due to the fact that a higher $\underline{b}$ corresponds to a decrease in the precautionary motive. But the incentive to households to stay out of the borrowing remains the same.

All in all, when we consider economies with income inequality driven by income uncertainty, our model does not rationalize the empirical finding.

2.5 Conclusion

In this paper, we study the role of income inequality in the association between credit deepening and consumption growth. We do so through two approaches. First, we use cross-country and panel estimations. Second, we use a workhorse Aiyagari model to check to which extent this theoretical approach can rationalize this empirical finding. Our results in the empirical part show that the association between consumption growth and credit expansion is stronger in countries with higher income inequality.

In the theoretical part, we consider two sources of income inequality in our model: the variance of the idiosyncratic risk and the households' fixed level of human capital. Our results suggest that when the source of income inequality comes from the households' lowest fixed level of human capital, our model can rationalize the empirical evidence. In the other cases, the opposite occurs.

3
FX interventions in Brazil: a synthetical control approach

3.1
Introduction

The Fed's taper announcement on May 2013 led to a major repricing of risk, adding pressure on several emerging market currencies. The Brazilian real (BRL) depreciated about 15 percent during the following three months, despite sizable interventions by the Central Bank of Brazil (BCB in the Portuguese acronym) in the foreign exchange market. On August 22, 2013, the BCB announced a major program of intervention through FX swaps, with the aim of satisfying the excess demand for hedging and providing liquidity to the FX market. The program consisted of daily sales of US$ 500 million worth of currency forwards (US dollar swaps) in the Brazilian markets, that provided investors insurance against a depreciation of the real. These swaps settle in domestic currency and provide investors the very same hedging they would obtain by buying spot dollars and holding them until the maturity of the swap.[1] The program also indicated that on Fridays, the central bank would offer US$ 1 billion on the spot market through repurchase agreements (short term credit lines in USD). The program announcement stated it would last until at least December 31, 2013. On December 18, 2013, the BCB announced that it would extend the program until at least mid-2014, although the daily interventions were reduced to US$ 200 million. On June 24, 2013, that program was extended until at least end-2014, and eventually extended until March 31, 2015.[2]

Figure 3.1 shows the behavior of the BRL exchange rate (an increase in the exchange rate denotes a depreciation of the BRL) and the magnitude of these interventions. The BRL was depreciating at a rapid pace prior to the announcement. That trend is immediately reversed, with the BRL appreciating 10 percent in the month following the announcement. All in all, the announcement implied a cumulative intervention of about US$ 50

[1]Because they settle in real, they involve convertibility risk. For a detailed discussion of these contracts, please refer to Garcia and Volpon (2014).

[2]For a detailed discussion of the program, please refer to Kang and Saborowski (2015).

billion through 2013-end. The program was eventually extended, as discussed above, and the total amount of currency forwards stood at about US$ 110 billion as of the time of writing. This amounts to roughly a third of total FX reserves, making the program one of the largest episodes of reserve deployment in countries with a floating exchange rate regime. Another unique aspect of the program is that intervention took place through swaps, which is a temporary form of intervention since the additional FX liquidity provided is eventually removed once the swap expires. The program and its extensions spanned a year and a half, so much of the maturing swaps were rolled-over. Nevertheless, it still provides an example of large scale temporary intervention (albeit over a long horizon), which stands in contrast to many other country experiences (and studies) where intervention occurs mainly in the direction of accumulating reserves.

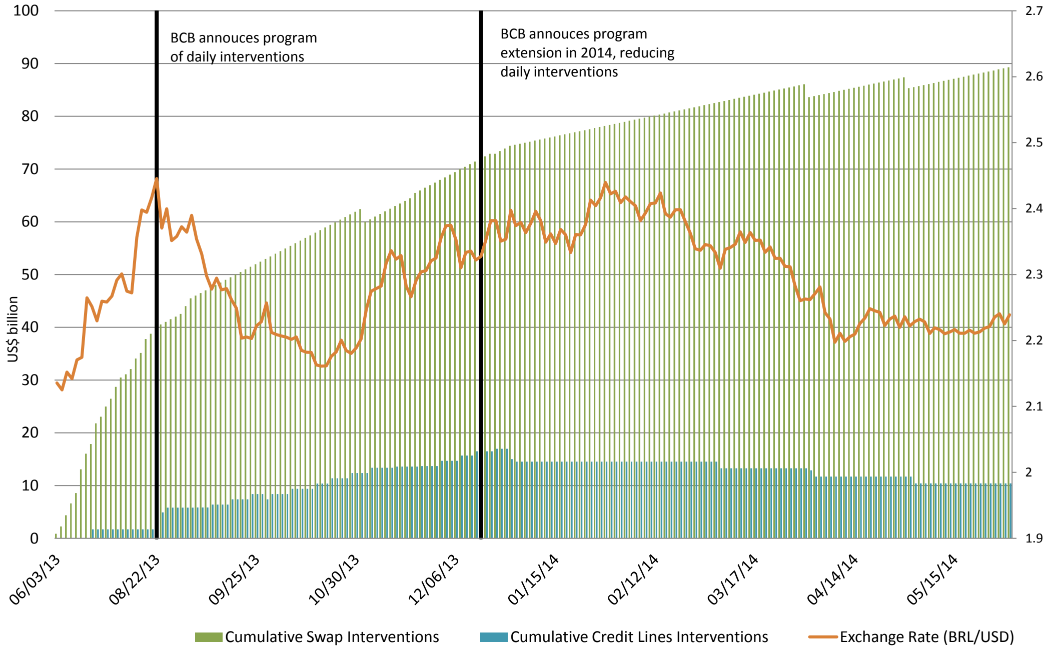


Figure 3.1: Cumulative Swap Interventions, Cumulative Credit Lines Interventions and Exchange Rate (BRL). Source: BCB and AC Pastore.

Most modern open economy models, assume uncovered interest parity holds, which leaves no scope for FX intervention to affect the exchange rate (some noteworthy exceptions include Benes et al. 2012, and Ghosh et al. 2015). Nevertheless, there is a very large empirical literature analyzing the effectiveness of central bank interventions. Sarno and Taylor (2001) survey the early literature, which typically focused on Advanced Economies and generally concluded that sterilized intervention was not very effective (with the possible exception of signaling future monetary policy). That is not surprising, since the amount of FX intervention pursued in advanced economies was a tiny fraction of the size of their bond markets. But in the case of Emerging Markets (EMEs), FX intervention has a non-trivial effect on the relative supply of local currency bonds. For example, in the case of Brazil, the stock of reserves corresponds to about a quarter of the stock of government bonds. So it seems reasonable to expect that a change in the relative supply of assets of that magnitude to have some effect on the exchange rate. A number of more recent papers focusing on emerging markets tends to find more supportive evidence for an effect, but the evidence remains somewhat mixed. Menkhoff (2013) provides an excellent survey of that literature.

In the Brazilian context, a number of papers have shown that FX intervention, including through swaps, can affect the exchange rate. For example, Andrade and Kohlscheen (2014) show that the Brazilian real moved about 0.33 bps following the announcement of a currency swap auction. Barroso (2014) estimates that a purchase or sale of US$ 1 billion lead to a 0.51 percent depreciation or appreciation of the Brazilian real. Werther (2010) found that the effects of sterilized interventions are very small on its magnitude (between 0.10 and 1.14 percent for each US$ 1 billion) and of low duration. More generally, estimates for the effect of a US$ 1 billion dollar intervention on the exchange rate typically range from 0.10 to 0.50 percent.

Studies on FX intervention face a substantial, perhaps insurmountable, endogeneity problem, since a central bank tends to purchase FX when it wants to slow down an appreciation, and vice-versa. That can bias regression estimates (perhaps even to the point of flipping the sign of the effect). Different strategies have been used to address this problem, including VARs, IV strategies, and relying on high-frequency data. All of these strategies have some drawbacks, including the extent to which they truly tackle this endogeneity.

In this paper we use a synthetic control approach to estimate the effects of the Brazilian swap program. To our knowledge, we are the first paper to use this technique to study the effects of FX interventions.[3] We follow Abadie et al. (2010), which in a nutshell, consists of constructing a synthetic control group that provides a counterfactual exchange rate against which we can compare the evolution of the Brazilian real after that announcement. This methodology is not appropriate for studying the effect of frequent interventions, but it is well suited for an event-study setting where a large change in intervention

[3]Jinjarak et al. (2013) use the synthetic control method to analyze the effects of the adoption and removal of capital controls in Brazil on capital flows and the exchange rate. Their results show that capital controls had no effect on capital flows and small effects on the the exchange rate.

policy is announced, as in the case of Brazil. Our counterfactual uses data from other countries, with weights that are based on the pre-announcement co-movement with Brazil. As a result, whatever noise and error is involved in this type of analysis, it will be orthogonal to the endogeneity problem that plagues the literature on FX intervention. Our findings point to an appreciation of the BRL in the first few weeks following the announcement of the program in excess of 10 percentage points. This is consistent with a surprise effect on the market, which by all accounts was not expecting the program. This result is particularly striking, once we take into account that the BCB was already intervening substantially in the market prior to the program, albeit in a discretionary fashion. In fact, the pace of intervention declined after the program (as shown in Figure 3.1). We also construct synthetic control groups using the methodology proposed by Carvalho et al. (2015), which allows us to make inference of the results. That approach points to a similar effect on the BRL (if anything stronger) following the announcement of the FX swap program, and that effect is statistically significant. Our results on the option-implied volatility are more mixed, with some of our estimates pointing to a tangible decline while others do not. A similar analysis of the follow-up announcements (extending the program) point to a more muted effect, which is not surprising since by most accounts the market was expecting the program to be extended in some form (so the surprise element was much smaller than in the previous announcements). Finally, as a robustness check, we also perform a more standard event-study analysis, which confirms a large effect on the exchange rate following the August announcement, but not for the latter announcements.

The remainder of the paper is organized as follows. Section 3.2 outlines the methodologies used, section 3.3 presents data description and section 3.4 shows our results. Finally, section 3.5 concludes.

3.2 Methodology

In this section, we present the synthetic control approach proposed by Abadie et al. (2010) and by Carvalho et al. (2015). Then, we use these methodologies to evaluate the effects of the BCB intervention programs on the Brazilian exchange rate.

3.2.1 Abadie et al. (2010)

Let Y_{it}^I denote the exchange rate in a country i in period t for a country that adopts a policy (e.g. an FX intervention program) at time T_0, and Y_{it}^N denote non-observed exchange rate that would have occurred had the country not adopted the FX interventions program.

We assume that there is no effect of the intervention program in the period preceding the policy change ($t < T_0$), i.e., $Y_{it}^N = Y_{it}^I$. Hence, the effect of the intervention program is given by $\alpha_{it} = Y_{it}^I - Y_{it}^N$ from period T_{0+1} to T. Without loss of generality, suppose the policy change occurred on country $i = 1$ (Brazil in our case). We assume that Y_{it}^N follows a factor model given by:

$$Y_{it}^N = \delta_t + \theta_t Z_i + \lambda_i \mu_t + \varepsilon_{it} \tag{3-1}$$

where λ_i is a factor loadings, Z_i is a vector of observable variables, θ_t is a vector of parameters and μ_t is an unknown common factor that depends on time. At last, ε_{it} is a mean zero iid shock.

In addition, consider $W = (\omega_2, ..., \omega_{j+1})'$ as a vector of weights such that $\omega_i \geq 0$ and $\sum_{i=2}^{j+1} \omega_i = 1$. Suppose that there is an optimal weight vector $\hat{W}$ that can accurately replicate pre-treatment observations in Brazil. Abadie et al. (2010) show that under regular conditions $Y_{it}^N = \sum_{i=2}^{j+1} \hat{\omega}_i Y_{it}$. Thus, we can calculate $\hat{\alpha_{1t}} = Y_{it} - \sum_{i=2}^{j+1} \hat{\omega}_i Y_{it}$ for $t \geq T_0$.

Define X_1 as a vector of pre-treatment characteristics of the Brazilian exchange rate that contains Y and Z, and similarly X_0 for the control countries. Hence, the optimal weight vector $\hat{W}$ is chosen through the minimization of the following equation

$$\sqrt{(X_1 - X_0\hat{W})'V(X_1 - X_0\hat{W})} \tag{3-2}$$

where V is a $k \times k$ symmetric and positive semi-definite matrix (k is the number of explanatory variables). Also V is chosen to minimize the mean square prediction error in the period prior to the policy change. We use the STATA *synth* routine to obtain V.

Finally, we use permutations tests to examine the significance of our results, due to the fact that the usual statistical inference is not available. For each control country in our sample, we assume that it implemented a FX intervention program in T_0. We then produce counterfactual synthetic control for each "placebo control" and calculate the effect α_{it}^P for $t \geq T_0$. Therefore, we can check if the effect found for Brazilian exchange rate is different from the effects on the control currencies.

3.2.2
Carvalho et al. (2015)

Consider n countries for T periods indexed by $i \in \{1, ..., n\}$. As in Abadie et al. (2010), assume that one country implemented a policy change in T_0. Furthermore, consider that we observe q variables for each country i and that they all follow jointly a covariance-stationary process. We can then stack all the n countries in a vector $y_t = (y_{1t}, ..., y_{nt})'$ and use the Wold decomposition to write the following equation for $1 \leq t \leq T$

$$\mathbf{y_t} - \mu_\mathbf{t} = \sum_{j=0}^{\infty} \phi_{t-j}\varepsilon_{t-j} \tag{3-3}$$

where each ϕ_{t-j} is a $(nq \times nq)$ matrix and the constraint $\sum_{j=0}^{\infty} \phi_{t-j}^2 < \infty$ must be satisfied for $1 \leq t \leq T$. Also, ε_t is a nq-dimensional serially uncorrelated white noise with covariance matrix Σ_t.

Moreover, consider that Brazil is indexed by 1 and define the direct effect in our variable of interest y_{1t} as

$$\delta_\mathbf{1t} = \mathbf{y_{1t}} - \mathbf{y_{1t}^*} \tag{3-4}$$

where $\mathbf{y_{1t}^*}$ is our variable of interest without the FX intervention program. But, $\mathbf{y_{1t}^*}$ is not observed, therefore, we have to estimate $\mathbf{y_{1t}^*}$ before estimate $\delta_\mathbf{1t}$. For this reason, we consider the best linear predictor as $(\mathbb{E}(y_{1t}^*|1, y_{-1t}^*))$

$$\mathbf{y_{1t}} = \mathbf{y_{1t}^*} = \mathbf{w_0} + \mathbf{w_1}\mathbf{y_{-1t}} + \mathbf{v_{1t}}, 1 \leq t \leq T_0. \tag{3-5}$$

where $\mathbf{y_{-1t}}$ is a matrix with all q variables for all $n-1$ countries (not including Brazil), $\mathbf{w_1}$ is a $(q \times (n-1)q)$ matrix and $\mathbf{w_0}$ is $(q \times 1)$ vector.

We estimate w by OLS for all the q equations.[4] Note that Abadie et al. (2010) approach consider that the weights should be non-negative and their sum should be equal to one. These restrictions provide a possible interpretation for the weights. However, Carvalho et al. (2015) argues that it is not clear the relevance of the interpretation when all that is needed is a strong correlation. For example, consider an extreme case where there is a perfectly negatively correlated country with Brazil. Under the restrictions adopted by Abadie et al. (2010), this peer would be disregarded despite the fact that using it would result in an almost perfect synthetic counterfactual. The opposite case is also troublesome, consider that all the peers are uncorrelated to Brazil. Due to the restriction to sum to one, the estimator automatically assign weights to countries that have no contribution in explaining the counterfactual trajectory.

[4]As stressed by Carvalho et al. (2015), it is one of the possible ways to estimate equation $(3-4)$.

Differently from Abadie et al. (2010), Carvalho et al. (2015) presents the statistical inference for the average direct effect between period T_{0+1} and T. Hence, we can test if the effect of the intervention programs on the Brazilian exchange rate is statistically significant. In addition, another moments can be tested. In our case, we are also interested to analyze if the FX swap program had an effect on the variance of the exchange rate. We consider the same linear specification as in (3-5) and our dependent and independent variables becomes $\ddot{\mathbf{y}}_{\mathbf{1t}} = (\mathbf{y}_{\mathbf{1t}} - \bar{\mathbf{y}}_{\mathbf{1t}})^2$ and $\ddot{\mathbf{y}}_{\mathbf{-1t}} = (\mathbf{y}_{\mathbf{-1t}} - \bar{\mathbf{y}}_{\mathbf{-1t}})^2$, respectively. Therefore, the average effect is also estimated and all the hypothesis testing can be carried on (see Carvalho et al. (2015) for more details.).

3.3 Data

Our analysis consider three outcome variables of interest: the exchange rate (bilateral exchange rate with respect to the USD), its 3-month option-implied volatility, and risk reversal. The latter measures the difference between the volatility implied by an out-of- the-money put option (25 delta) and an equivalent out-of-the-money call option, which is a measure of the insurance premium investors are willing to pay to insure against a risk-off episode. Figures 3.2 plots the evolution of the option-implied volatility over time. There was a rapid increase in volatility following the "tapering" speech. Volatility declines substantially after the program announcement, eventually settling at a lower level (although still higher than the volatility prior to the tapering speech). Volatility does not respond much in the immediate aftermath of the program extension announcements. Figure 3.3 is analogous to Figure 3.2 but plots the evolution of the option-implied risk reversal. There is a marked reduction following the program and the first extension.

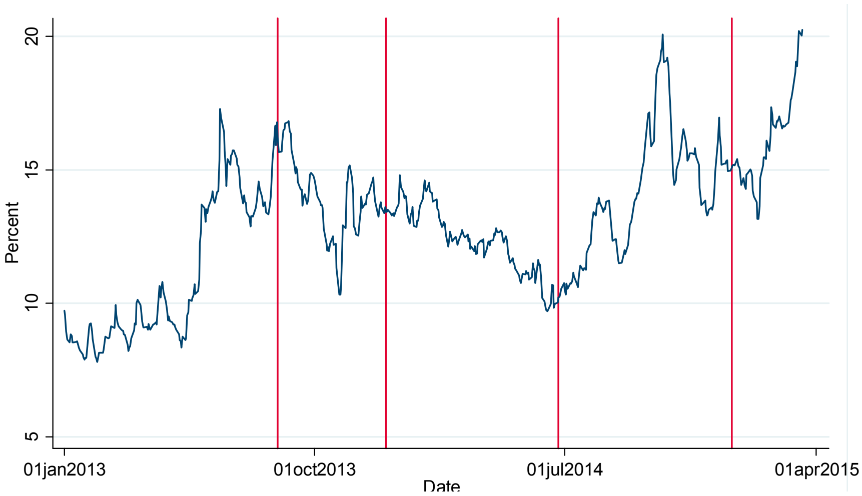


Figure 3.2: Brazilian Real Option-Implied Volatility. Notes: Vertical bars indicate the program announcement and extensions. Source: Bloomberg.

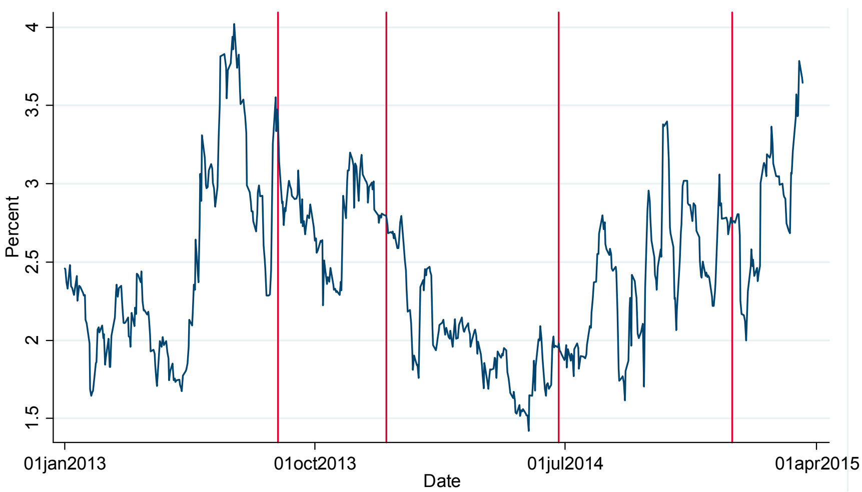


Figure 3.3: Brazilian Real Option-Implied Risk Reversal. Notes: Vertical bars indicate the program announcement and extensions. Risk Reversal measures the difference between implied volatility of out-of-the-money put and out-of-the-money call (25 delta). Source: Bloomberg.

In addition to these outcome variables, explanatory variables include capital flows, and stock and bond market indices. The source of all data is Bloomberg, except for the capital flow series which comes from the Emerging Portfolio Fund Research (EPFR) database. We use weekly data in our synthetic estimates (the highest frequency at which the capital flows series is available). For each event, we consider a window consisting of the 12 weeks prior to the announcement, the week of the announcement, and the 12 weeks afterwards.

We consider a sample of 16 countries when estimating the synthetic for Brazil, which includes: Australia, Brazil, Chile, Colombia, India, Indonesia, Korea, Malaysia, Mexico, New Zealand, Peru, Philippines, Poland, Russia, South Africa, Thailand, and Turkey. We included all the emerging market countries with EPFR data plus Korea, and Australia and New Zealand (the latter two because they are major carry trade currencies).

For the implementation of both methodologies, the series used should be stationary. For this reason, we use the log difference of the exchange rate, equity and bond indices, and the difference of the option-implied volatility and risk-reversal in our analysis. Capital flows to each country are scaled by the 2012 GDP in US dollars for each country.

3.4 Results

In this section, we use the approaches presented on the methodology section to analyze the FX intervention programs in Brazil. In addition, we present an event study to check the robustness of our results.

3.4.1 Program Announcement

Level effect

Figure 3.4 presents our estimates for the effect of the program announcement on the exchange rate. As mentioned above, the estimation uses the log change in the exchange rate as the dependent variable. But in order to more easily illustrate the resulting effect on the level, we accumulate the weekly log differences for the actual and for the synthetic exchange rates, and report the gap between the two. That gap is set to zero on the last observation prior to the announcement (so the level at any date t corresponds to the gap in the accumulated log differences from t to the announcement, and vice-versa). Figure 3.4(a) shows the estimates using Abadie et al. (2010) approach. In addition to the log change in the exchange rate, the explanatory variables considered include capital flows, the change in volatility, and the log change in the equity and bond indices. The thick dark line indicates the gap between the actual BRL and its synthetic (a negative value indicates that the BRL was more appreciated than its synthetic), while light gray lines indicate the gap for the other countries, which is used as a placebo test. The gap for the BRL is slightly negative and broadly stable during most of the pre-announcement period. But the gap declines sharply after the announcement, remaining at a substantially negative level. The bulk of the change takes place in the first week (about 10 percentage points). But the trend persists with the gap peaking at close to

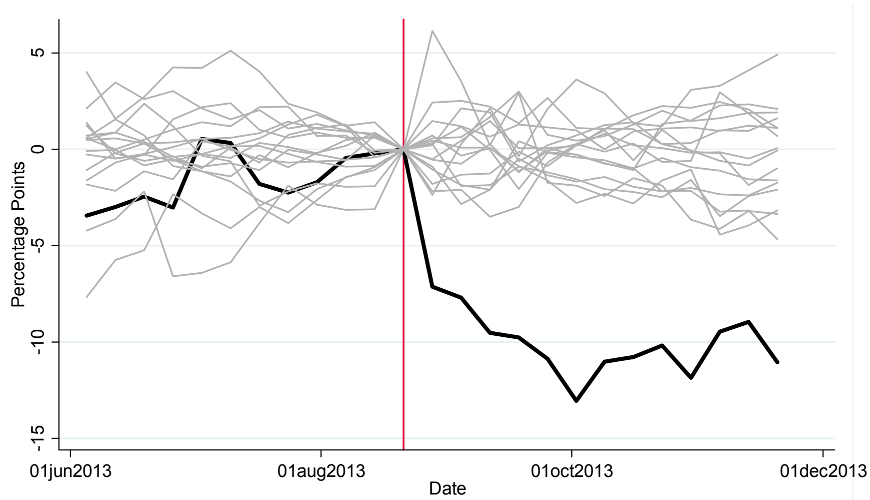
15 percentage points before narrowing slightly. These results imply that the BRL was over 10 percentage points stronger than what its synthetic would suggest weeks after the announcement. Moreover, please note that the gap for the BRL is a major outlier vis-a-vis the placebos in the post-announcement period, with none of the placebos experiencing nearly as large a shift (in the pre-announcement period, both the BRL and placebos should hover around zero by construction). The weights and countries used for the construction of the synthetic control group do not have an economic interpretation, a point that is stressed in the literature (e.g. Abadie et al. 2010).[5,6] The means for Brazil and for its synthetic are reported in Table C.1.

The effect of this program is also estimated using a univariate approach that considers only the exchange rate, following the methodology proposed in Carvalho et al. (2015). Under this approach, we cannot consider all peers and control variables (otherwise there would be more parameters being estimated than the data available). We choose 3 peers that maximize the fit of the exchange rate regression: South Africa, Thailand and Peru. The counterfactual is estimated through a regression of the BRL on the others peers' change in log of exchange rate and a constant.[7] The gap between the actual and synthetic BRL is reported in Figure 3.4(b). The results point to a cumulative effect that is even stronger, peaking at around 20 percentage points. This approach provides a statistical inference for the average effect, which is statistically significant (with a p-value below 2 percent at four lags). The effect is smaller when the counterfactual is estimated without a constant (around five percentage points).

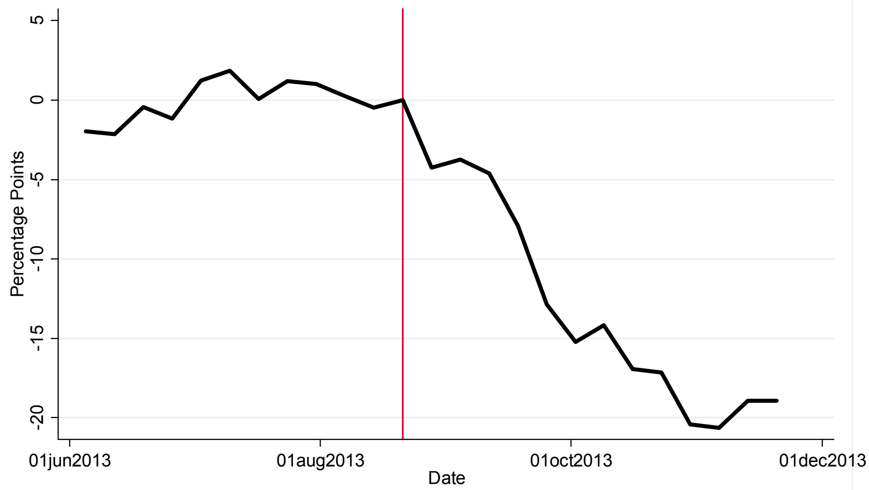
[5] With that caveat in mind, the synthetic draws from India, Indonesia and Malaysia, with weights of 14, 76, and 9 percent, respectively.

[6] Results are similar when we consider only Inflation Targeting countries.

[7] The R^2 of a regression of BRL in these currencies is equal to 0.8.



3.4(a): Gap Between Actual and Synthetic



3.4(b): Gap Between Actual and Univariate Synthetic

Figure 3.4: Effect of the Program Announcement on the Level of the Exchange Rate and Placebo Tests. Notes: Figures plot gap between the cumulative change in the log of the actual exchange rate and that implied by the synthetic estimates. Thick dark line indicates the gap for Brazil, and light gray lines indicate the gap for estimates from other countries (placebos). For ease of illustration, gaps are set to zero on the last observation prior to the announcement, which is indicated by the vertical line. Panel A based on the methodology in Abadie et al. (2010) and Panel B based on Carvalho et al. (2015).

Volatility effect

The approach in Carvalho et al. (2015) allows us to estimate other moments of the exchange rate. We can estimate an effect on volatility by using the squared change in the log of the exchange rate as the dependent variable (and the corresponding variable for other countries as the explanatory variable). The estimates suggest the average effect on the variance is close to zero and not statistically significant.

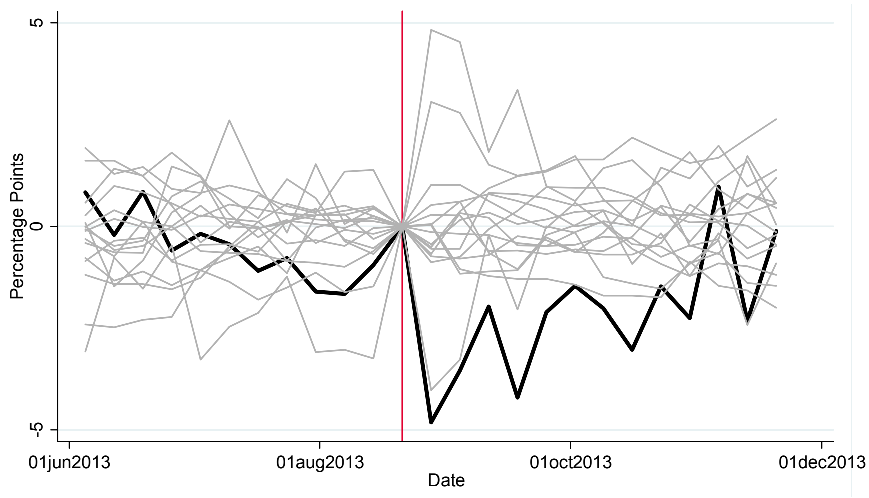
We can also assess the impact of the program on volatility using the option-implied exchange rate volatility. This readily available series provides a forward-looking measure of volatility (since it is based on option prices) that can quickly respond to the program (unlike say, measures of volatility constructed from past exchange rate data). Figure 3.5 reports the results for the change in the volatility. In Figure 3.5(a) we use the changes in the exchange rate, equity and bond indices, and capital flows as explanatory variables. For ease of illustration, we accumulate all the changes so as to report the resulting level of effect (setting the level at the last observation prior to the announcement to zero). Again, the thick dark line corresponds to the BRL while the thin gray lines to the placebo tests. There is a sharp decline in the gap in volatility after the announcement, by 5 percentage points, which is driven mainly by an increase in volatility among the countries in the synthetic control (India in particular) rather than an absolute decline in volatility for Brazil).[8] If we drop India from the pool of potential countries for the synthetic control, the results continue to point to a decline in volatility, but of only 2 percentage points.[9] That would still be a sizable decline (to put magnitudes in perspective, the volatility of the BRL was about 17 percent in the last observation prior to the announcement, so a 2 percentage point decline amounts to over 10 percent of the original volatility). The placebo tests point to the BRL being an outlier after the announcement. But the discrepancy between the BRL and the placebos is much smaller than in Figure 3.5(a).

Figure 3.5(b) reports the results using the univariate approach, drawing on Peru and India. The results are more muted, and not statistically significant.

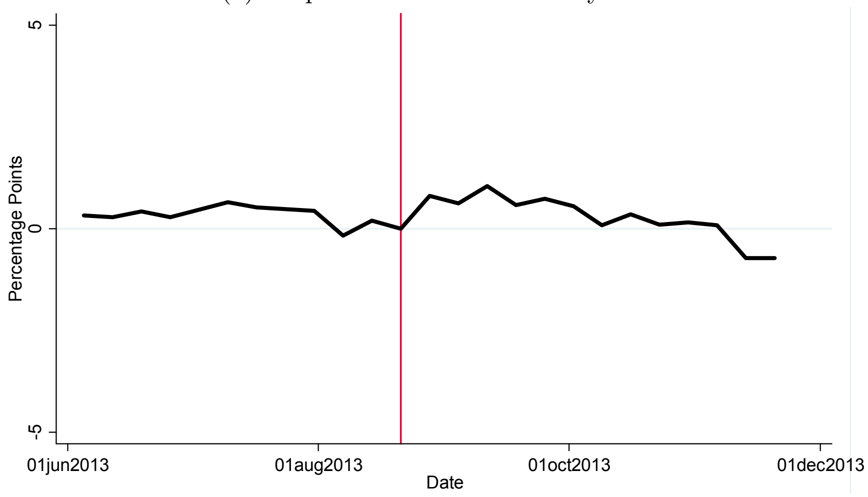
Finally, Figure 3.6 is analogous to Figure 3.5(a) but reports results for the risk-reversal measure. There is a sharp decline following the announcement (driven mainly by a decline in that variable for Brazil, which goes from 3.5 to 2.7 in the two observations before and after the announcement). A comparison with the placebos suggests the behavior of the BRL was an outlier in the two weeks following the announcement, but not afterwards.

[8]The synthetic draws on Australia and India, with weights of 31 and 69 percent, respectively.

[9]The synthetic would draw on Australia and Indonesia, with weights of 64 and 36 percent, respectively.



3.5(a): Gap Between Actual and Synthetic



3.5(b): Gap Between Actual and Univariate Synthetic

Figure 3.5: Effect of the Program Announcement on the Option-Implied Volatility of the Exchange Rate and Placebo Tests. Notes: Figures plot gap between the cumulative change in the option-implied volatility and that implied by the synthetic estimates. Thick dark line indicates the gap for Brazil, and light gray lines indicate the gap for estimates from other countries (placebos). For ease of illustration, gaps are set to zero on the last observation prior to the announcement, which is indicated by the vertical line. Panel A based on the methodology in Abadie et al. (2010) and Panel B based on Carvalho et al. (2015).

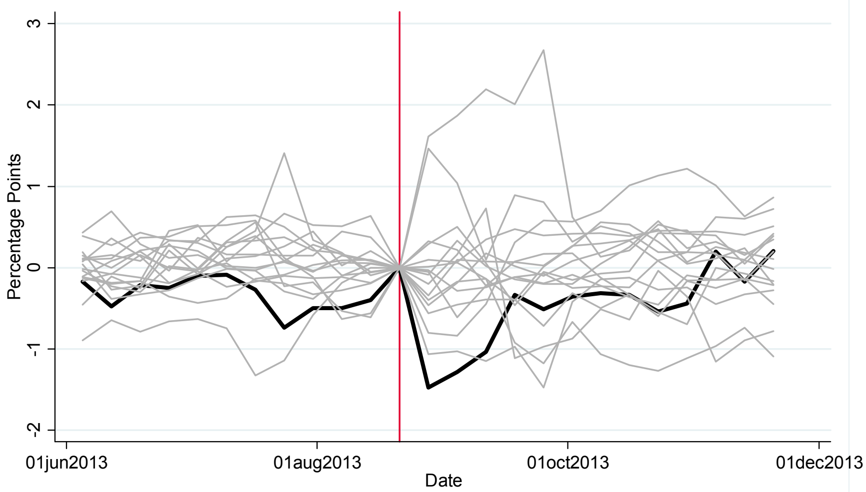


Figure 3.6: Effect of the Program Announcement on the Option-Implied Risk Reversal of the Exchange Rate and Placebo Tests. Notes: Figures plot gap between the cumulative change in the risk reversal and that implied by the synthetic estimates. Thick dark line indicates the gap for Brazil, and light gray lines indicate the gap for estimates from other countries (placebos). For ease of illustration, gaps are set to zero on the last observation prior to the announcement, which is indicated by the vertical line. Based on the methodology in Abadie et al. (2010).

3.4.2
Program Extension Announcement

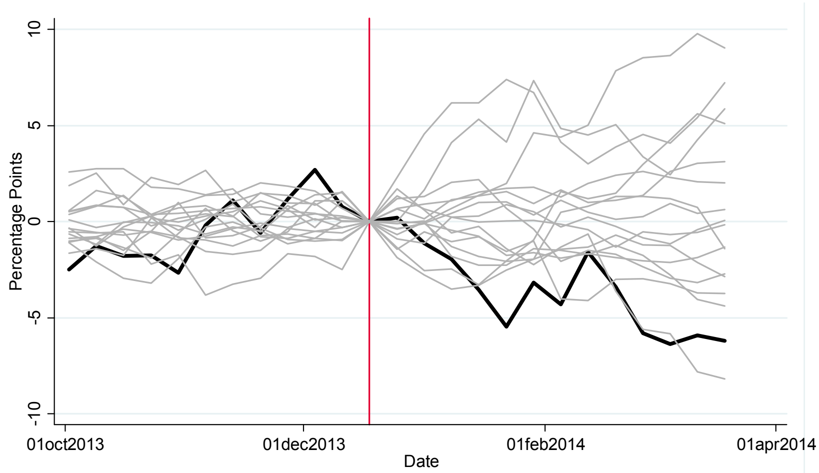
Level effect

On December 18, 2013, the intervention program was extended until mid-2014, but with reduced daily interventions. There were expectations that the swap sales would continue (i.e. the market did not expect it to end abruptly at the end of 2013), but the announcement removed that uncertainty and clarified the scope of the program going forward. Therefore, the announcement could still impact the exchange rate, but that impact should be less dramatic than the one following the first announcement.

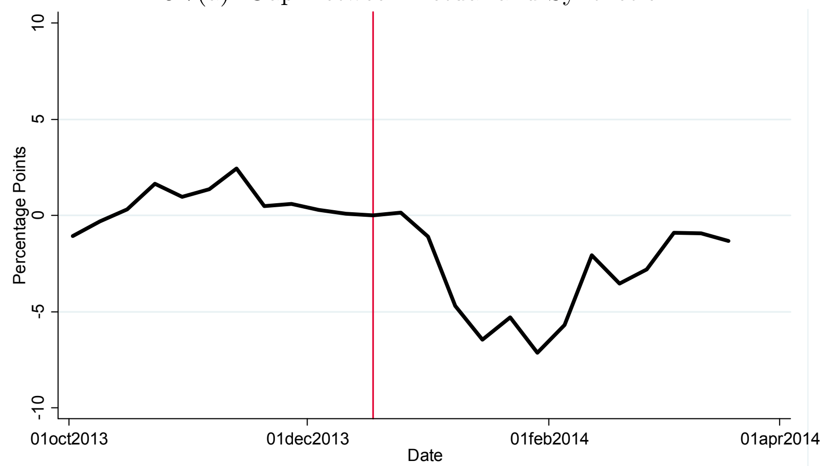
Figure 3.7 is analogous to Figure 3.4, but reports the results for the cumulative changes in the exchange rate around this second announcement. Figure 3.7(a) points to a gradual appreciation of the BRL vis-a-vis its synthetic, with that gap reaching about 5 percentage points, and remaining close to that level. A comparison with the gaps for the placebos suggest that the BRL was clearly on the stronger side, but was not nearly as much of an outlier as in

Figure 3.4(a).[10]

Figure 3.7(b) reports the result under the univariate approach. The results also point to a decline of around 5 percentage points over the first four weeks, but that is gradually reversed over time. The effect is not statistically significant under any lag structure.



3.7(a): Gap Between Actual and Synthetic



3.7(b): Gap Between Actual and Univariate Synthetic

Figure 3.7: Effect of the December 2013 Announcement on the Level of the Exchange Rate and Placebo Tests. Notes: See notes to Figure 3.4.

Volatility effect

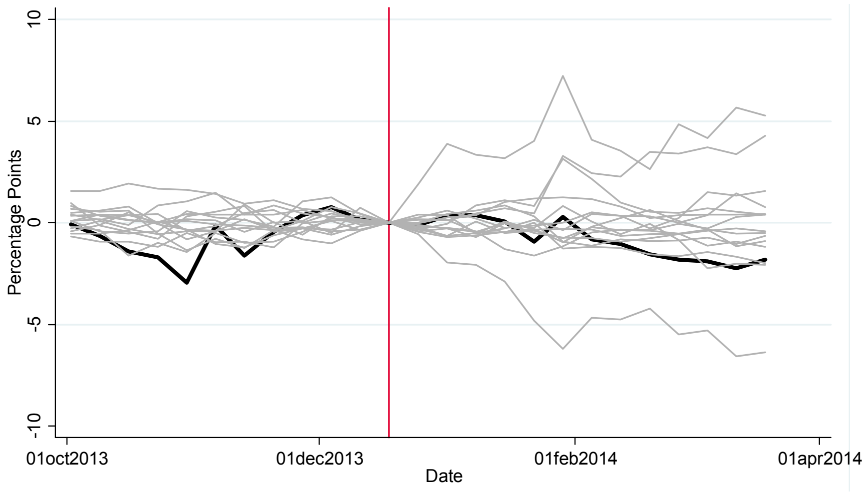
Figure 3.8 is analogous to Figure 3.5, but reports the effect on the option-implied volatility following the second announcement. There is virtually no change in volatility under neither of the methodologies considered. We also do not find any statistically significant effect of the second announcement when we estimate the synthetic for the squared log change in the exchange

[10]The synthetic draws on Australia, Indonesia, Peru and Turkey, with weights of 19, 9, 5 and 67 percent, respectively.

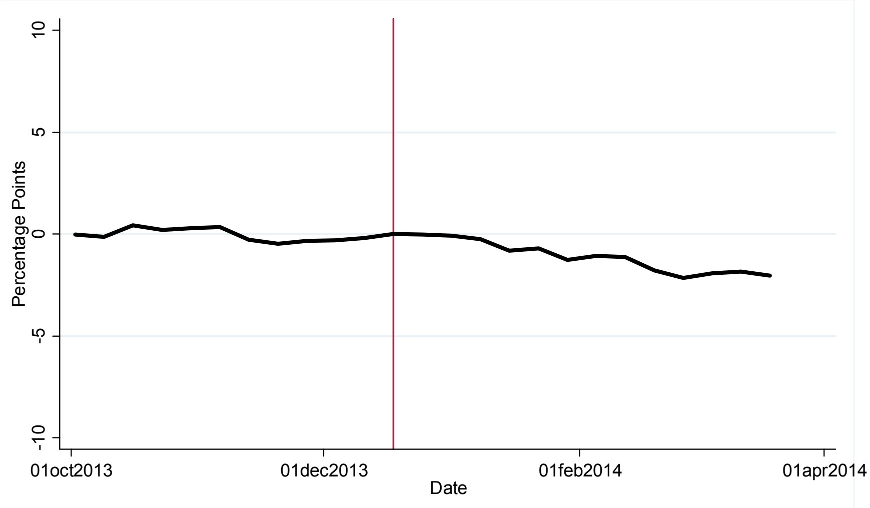
rate, using the univariate approach. There is also virtually no effect on the risk reversal following the second announcement (Figure 3.9). While there is a sharp decline in risk reversal for Brazil following the second announcement, as shown in Figure 3.3, the same was true for its synthetic (which draws heavily from Peru, where a sizable decline also took place around that time).

3.4.3 Addition Program Extensions

There were two additional announcements. One on June 24, 2014 extending the program until at least 2014-end, and a final announcement on December 30, 2014 extending the program until March 31, 2015. Figures 3.10 and 3.11 reports the results for the level of the exchange rate. The estimates suggest virtually no effect on the BRL exchange rate following the June 2014 announcement. The results point to a larger gap following the December 2014 announcement, which peaks at an appreciation of around 5 percent before quickly reversing. But overall, the results for the BRL are broadly in line with the placebos during most of the post-announcement period, suggesting no significant effect. The results for the volatility and risk reversal also point to little or no effect, and are not reported for the sake of conciseness.



3.8(a): Gap Between Actual and Synthetic



3.8(b): Gap Between Actual and Univariate Synthetic

Figure 3.8: Effect of the December 2013 Announcement on the Option-Implied Volatility of the Exchange Rate and Placebo Tests. Notes: See notes to Figure 3.5.

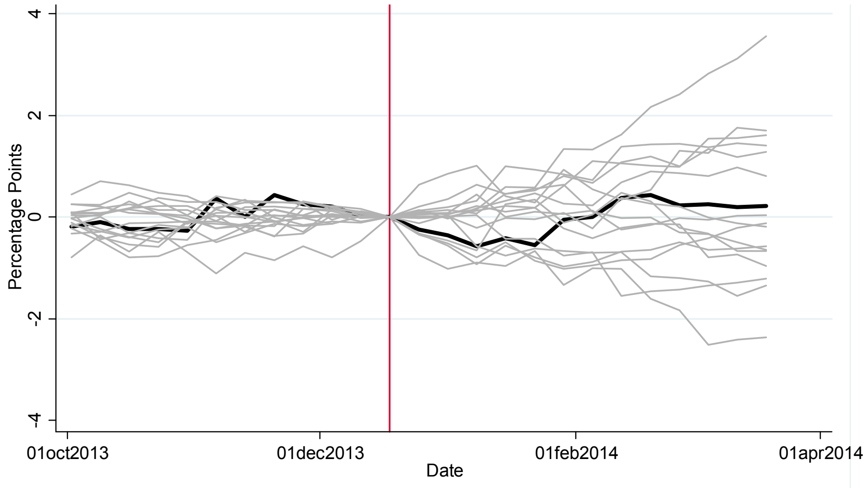


Figure 3.9: Effect of the December 2013 Announcement on the Option-Implied Risk Reversal of the Exchange Rate and Placebo Tests. Notes: See notes to Figure 3.6.

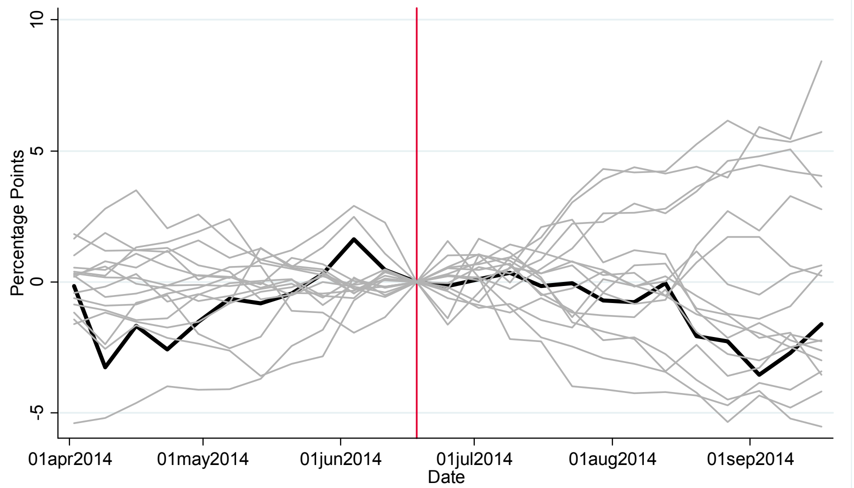


Figure 3.10: Effects of the June 2014 Announcement on the Level of the Exchange Rate and Placebo Tests. Notes: See notes to Figure 3.4.

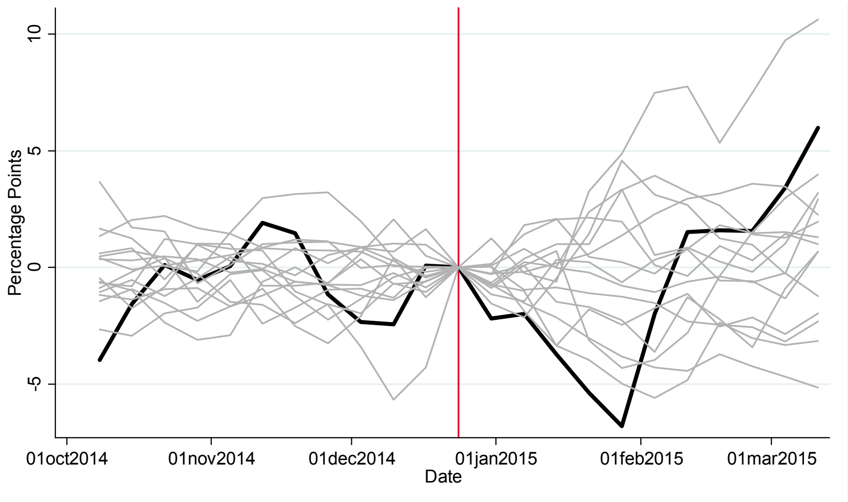


Figure 3.11: Effects of the December 2014 Announcement on the Level of the Exchange Rate and Placebo Tests. Notes: See notes to Figure 3.4.

3.4.4 Event Study

As a robustness check, we complement our analysis with a standard event-study analysis around the announcement of the FX swap program.[11] Using daily data, we estimate:

$$\begin{aligned}\Delta log(e_t) = c + \gamma_1 \Delta(CDI_t - LIBOR_t) + \gamma_2 \Delta log(VIX_t) \\ + \gamma_3 \Delta log(Commodities_t) + \gamma_4 \Delta log(DollarIndex_t) \\ + \gamma_5 \Delta(Dollar - AsiaIndex_t) + \gamma_6 FXInt_t + \varepsilon_t \end{aligned} \quad (3\text{-}6)$$

Where e is the dollar-real bilateral exchange rate, and explanatory variables include the change in the spread between the one-month CDI (Brazil's interbank rate) and the one-month $LIBOR$, the change in the log of the VIX, the change in the log of the CRB commodity price index, the change in the log of an index constructed by the Federal Reserve for the value of the dollar relative to major currencies of advanced economies weighted by US trade shares, the change in the log of the Bloomberg JP Morgan Asia and Latin America currency indices (we recomputed the latter, based on published weights, to exclude the BRL), and the Foreign Exchange Intervention by the central bank (based on announced swaps, netting out maturing ones).[12]

[11]Please refer to Campbell, Lo and MacKinlay (1996) for a description of the event study approach.

[12]The data sources are: Central Bank of Brazil for the exchange rate; Federal Reserve Economic Data for the dollar index, and Bloomberg for the remaining series.

We estimate this regression using data for January 2013 until 20 days prior to the August 22 announcement. We then compute the change in the log of the exchange rate beyond what would have been implied by that fitted model (analogous to the Cumulative Abnormal Returns in a standard finance event study) and the corresponding error bands around that estimate. We consider a +/- 20 working day window around the two announcements. Figure 3.12 reports the results, which point to a statistically significant cumulative appreciation of about 10 percent after the August 22 announcement, in line with our synthetic cohort estimates. In contrast, there is virtually no response following the December 18 announcement.

We estimate a similar regression but using the change in the option-implied volatility and risk reversal as the dependent variables. Figure 3.13(a) reports the results for volatility. While there is a decline following both announcements, it is not statistically significant (the error bands are too wide and span a zero effect). Figure 3.13(b) reports the results for the risk reversal. It declines following both announcements. That cumulative decline is statistically significant in the immediate aftermath for the first announcement, but over time the error bands become wider and that is no longer the case. In the case of the second program, the error bands initially span zero, but that is no longer the case towards the end of the post-announcement window the cumulative effect. The cumulative effect points to a 1.4 percentage point decline, which is sizable (the risk reversal stood at 2.8 prior to the announcement).

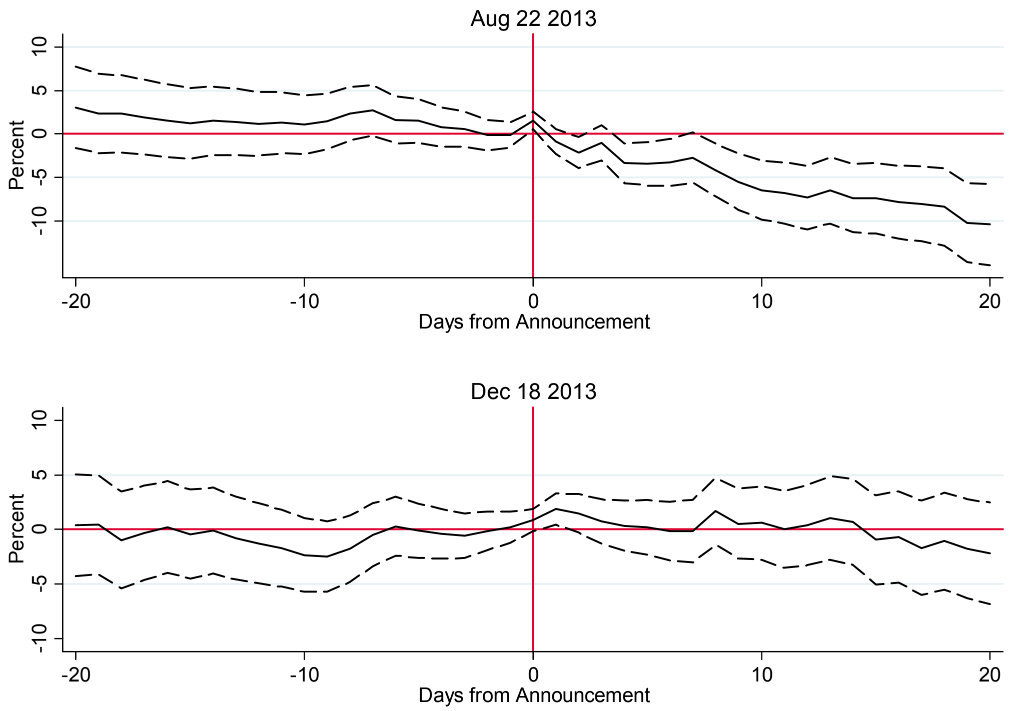
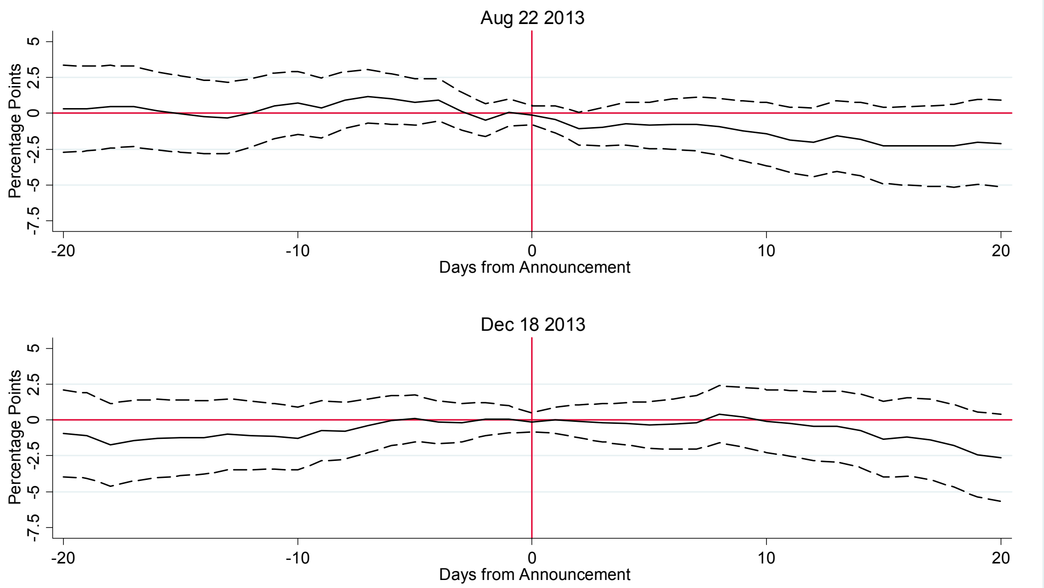
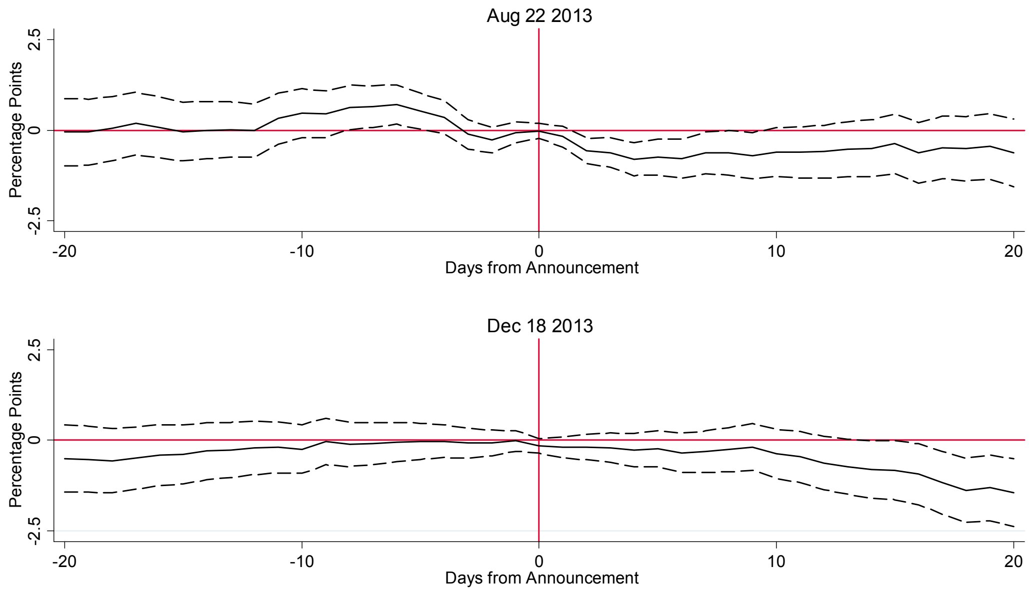


Figure 3.12: Cumulative Changes in the Exchange Rate Around Program Announcement and Extension. Notes: Dashed lines correspond to +/- 2 Standard Deviations. Cumulative changes start at 0 for both and after period.



3.13(a): Option-Implied Volatility



3.13(b): Risk Reversal

Figure 3.13: Cumulative Changes in the Option-Implied Volatility and Risk Reversal of the Exchange Rate Around Program Announcement and Extension. Notes: Dashed lines correspond to +/- 2 Standard Deviations. Cumulative changes start at 0 for both and after period.

3.5 Conclusion

The gyrations in international capital markets have brought renewed interest in tools to manage capital flows, with intervention in FX markets being one of the most commonly used tools. This paper has analyzed the effect of the large scale program of FX swaps that the BCB has embarked following the market's "taper tantrum" of 2013. This program was fairly unique because of its large scale (amounting to about a quarter of reserves) and the fact that the intervention took place through swaps (which makes the intervention temporary in nature, despite the long horizon of the program).

Immediately after announcement of the program, on August 22 2013, the Brazilian real reverted its depreciation trend, and eventually stabilized at a significantly more appreciated level. Our synthetic estimates point to an eventual appreciation relative to the synthetic in the range of 10-19 percentage points. The event-study analysis in the previous section also points to an appreciation of around 10 percent. If we compare this effect with the total volume of intervention mobilized during that program, it would be broadly in line with the point estimates for the effectiveness of FX intervention in Brazil from previous studies. Despite this large effect on the level of the exchange rate following the first announcement, the results on the volatility are more mixed. Some estimates point to a sizable decline, but overall the estimates are less robust than those for the level. Our estimates for the announcement of the extension of the program on December of 2013 had smaller effect on the exchange rate, ranging from no effect to 5 percent, and does not seem to have had an effect on its volatility. This smaller response may be the result of that extension being already expected and priced-in by the market. The third and fourth extensions had a fairly muted effect, likely for the same reason.

Our results are consistent with the view that FX interventions can be effective in deter- ring exchange rate overshooting in times of market turmoil. The large size of the program, and the market surprise following its announcements facilitate the identification of an effect, which would be more challenging in the context of small and frequent interventions that have come to be expected by the market.

4
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A
Quantitative Analysis: Mexico

In this appendix, we calibrate our baseline model for Mexican economy. Then, in order to assess the macroeconomic effects of the credit expansion observed, we solve for the time-varying paths of τ_t^{WL}, τ_t^H and τ_t^K that generate paths for personal credit, housing credit to households, and credit to corporations that resemble their counterparts in the data (see Figures 1.2 and 1.3).

A.1
Calibration

Table A.1 summarizes our calibration of the main parameters for Mexican economy. Our analysis is between the first quarter of 2006 and third quarter of 2013. As in the calibration for Peru, we follow the literature to set the values of some parameters when estimates are not available. For Mexican economy, we can not find the values of some parameters in the literature. In these cases, we use the values as for Peruvian economy. Then, calibration for Mexico deviates from the Peru specification along the following dimensions. We set steady state inflation to 4.09% to match the average inflation in the period. For the discount factor of patient households, we choose $\beta^p = 0.996$ to generate an average bank funding rate (a proxy for monetary policy rate) of 5.74%.

We also pick the inverse of the Frisch elasticity of labor supply equal to $\varphi = 1$. This value is standard in the literature for Mexican economy. For the capital share in the entrepreneurs' production function, we choose $\alpha = 0.33$ following Comin et al. (2010).

The parameter that measures price stickiness (Rotemberg adjustment cost) in the final goods market κ_P is set at 81.97, which is equivalent to 0.78 in the Calvo model.

For the monetary policy rule, we use estimates from Cicco (2010). In particular, $\phi_y = 0.54$, $\phi_\pi = 1.35$ and $\rho = 0.56$.

For the banking sector parameters, we set $\gamma = 2$ and $\eta = 0.06$ to generate a spread of roughly 3.1 percent per year. This value corresponds to the average difference between the implicit interest rate to enterprise, which is the lending interest rate that banks charge their clients and the bank funding rate, which is an interbank rate that banks use to lend to each other in the period of analysis.

Parameter	Description	Value
β^p	Discount Factor - Patients	0.99606
β^i, β^e	Discount Factor - Impatients and Entrepreneurs	0.91
μ^p, μ^i, μ^e	Mass - Patients, Impatients and Entrepreneurs	1
φ	Inverse of the Frisch Elasticity	1
σ	Elasticity Between Final Good and Housing	0
ξ	Weight of the Final Good on the Utility Function	0.8
δ_K, δ_H	Depreciation - Capital and Housing	0.025
κ_K, κ_H	Adjustment Cost - Capital and Housing	2.53
α	Capital Share in the Production Function	0.33
θ	Share of Patient Households in the Production Function	0.7
κ_P	Price Adjustment Cost - Final Good	81
ι	Steady State Inflation Weight - Indexation	0.158
ε	Elasticity of Substitution - Final Good	6
ρ	Smoothing Parameter of the Taylor Rule	0.56
ϕ_y	Output Weight of Taylor Rule	0.54
ϕ_π	Inflation Weight of Taylor Rule	1.35
η	Spread	0.06
γ	Spread	2

Table A.1: Calibration: Mexican economy.

A.2 Results

In this section, we report the macroeconomics effects of a credit deepening using the calibrated model for Mexican economy. For the interpretation of these results, see section 1.4.2.

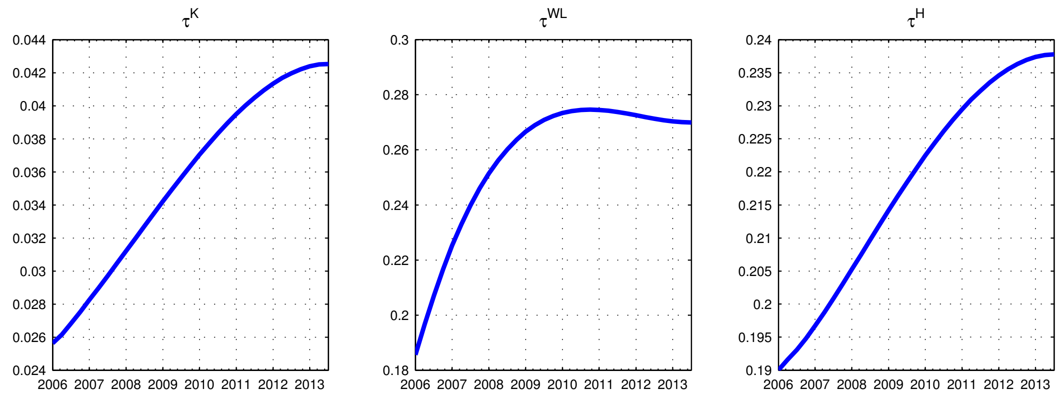

Figure A.1: Credit deepening experiment for Mexico: evolution of τ_t^K, τ_t^{WL} and τ_t^H.

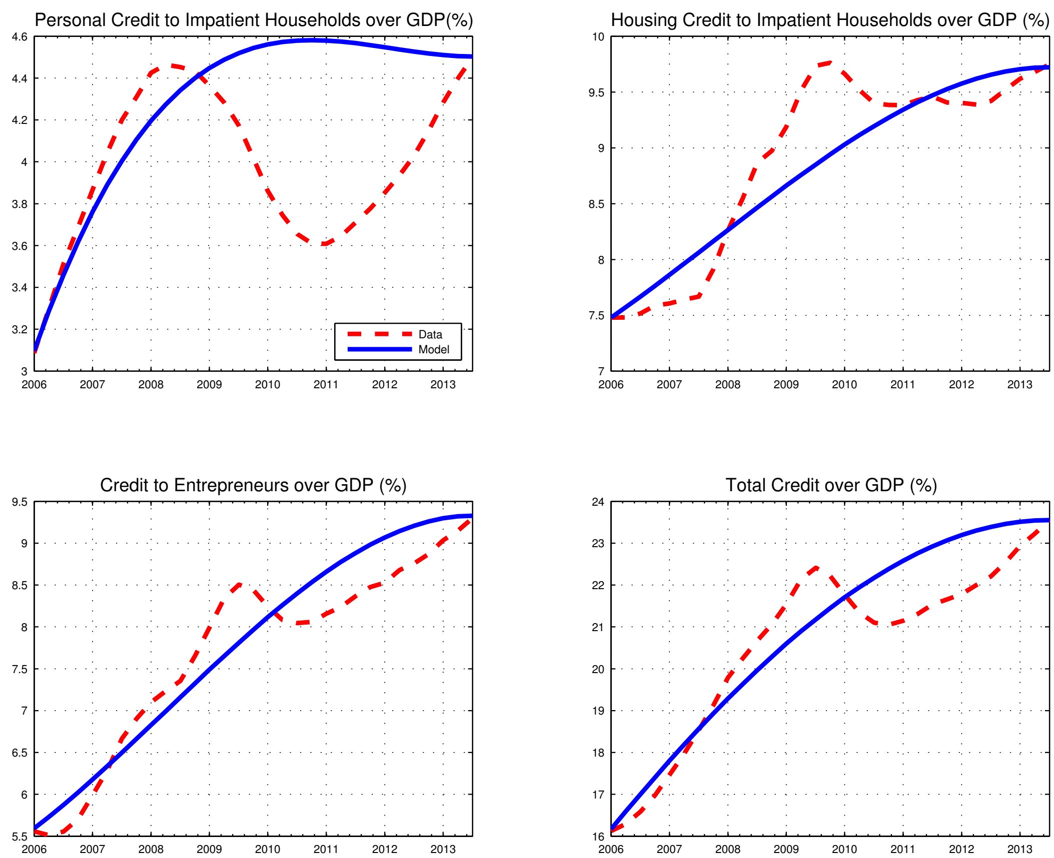

Figure A.2: Credit deepening experiment for Mexico: credit variables (model and data).

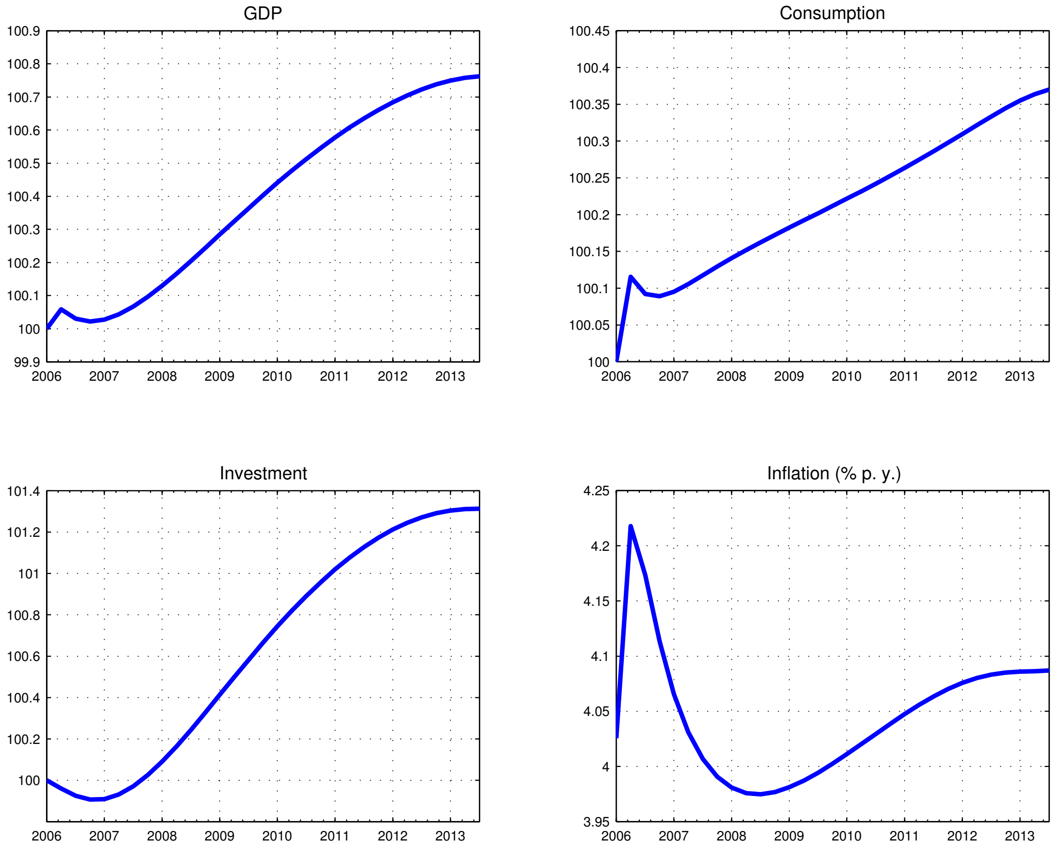

Figure A.3: Credit deepening experiment for Mexico: macro variables (model).

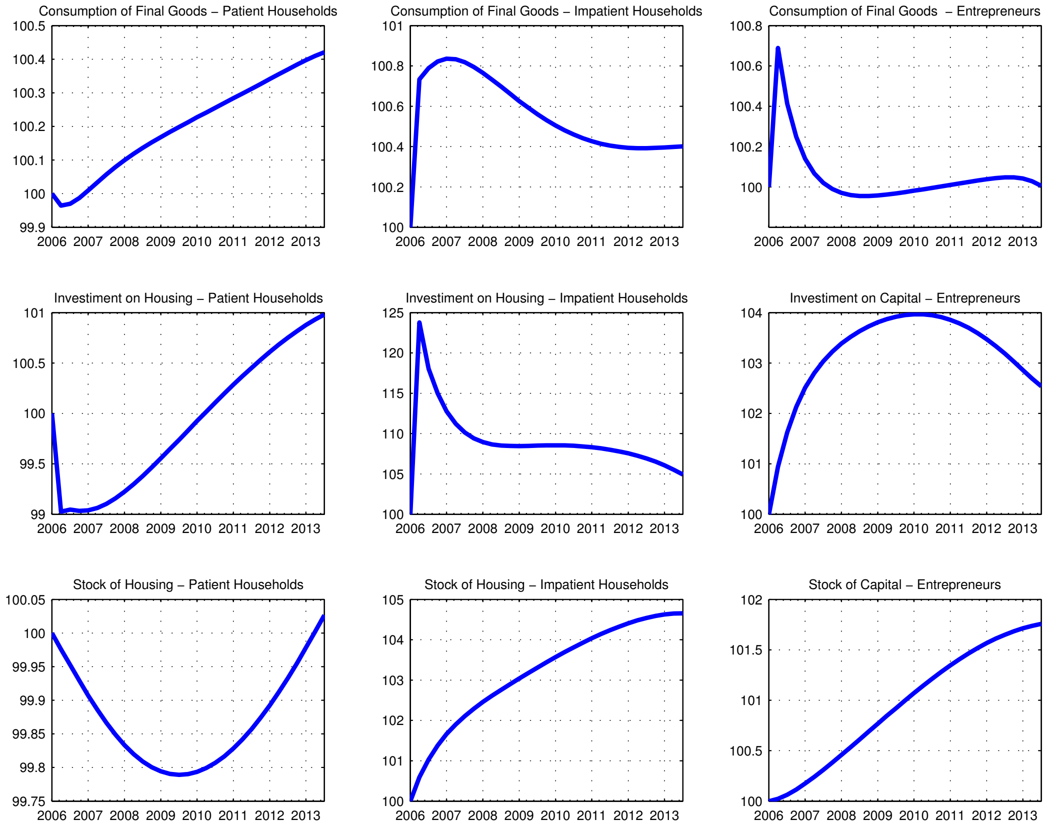

Figure A.4: Credit deepening experiment for Mexico: consumption, investment and stocks.

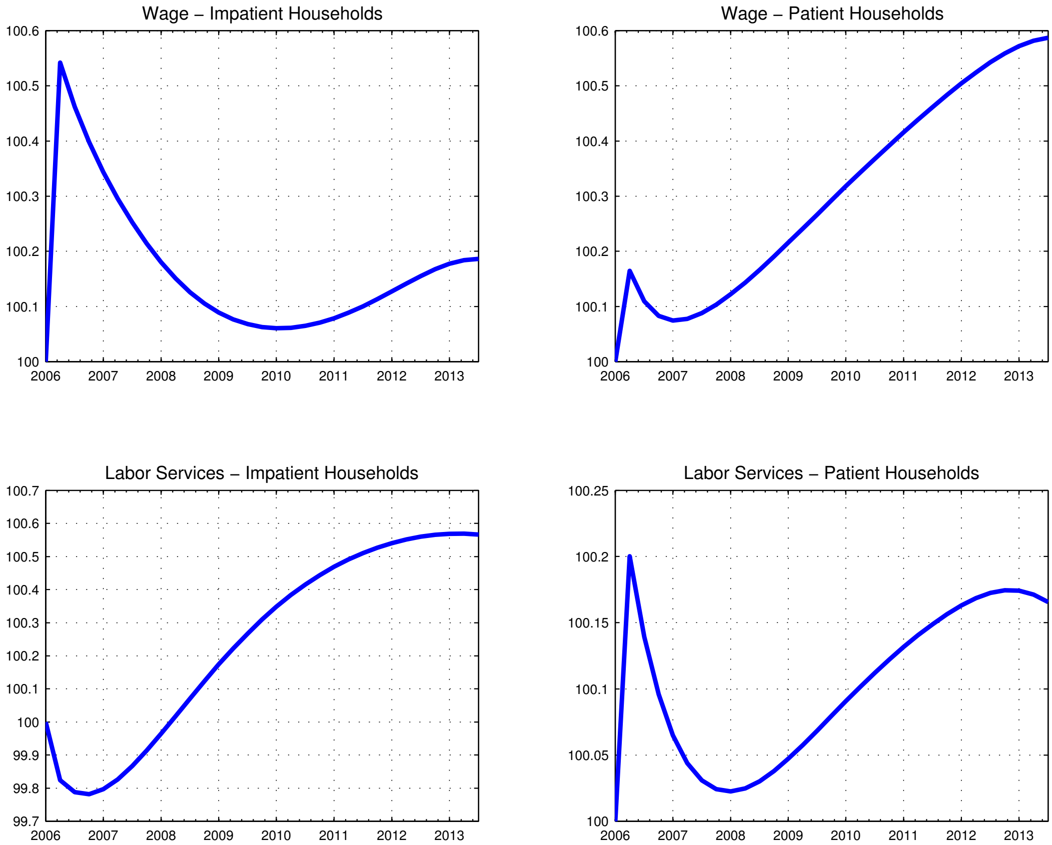

Figure A.5: Credit deepening experiment for Mexico: labor market outcomes.

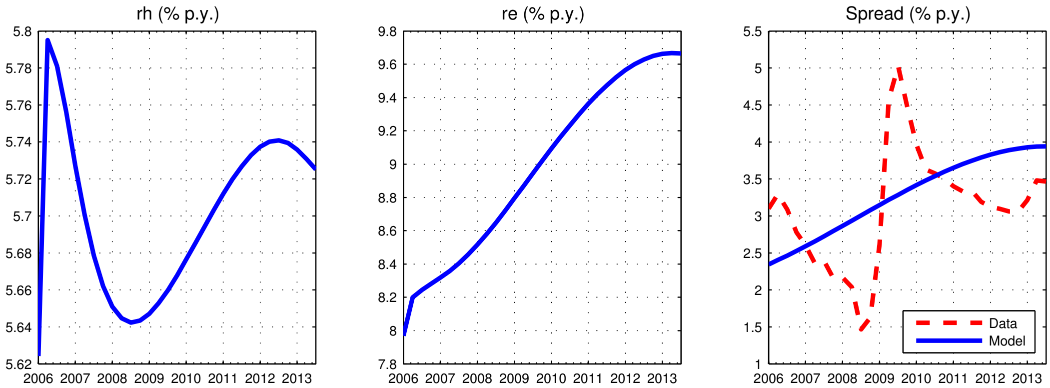

Figure A.6: Credit deepening experiment for Mexico: financial market outcomes. The spread is calculated using the bank funding rate, which is an interbank rate that banks use to lend to each other (proxy for monetary policy rate), and the implicit interest rate to enterprise which is the lending interest rate that banks charge their clients. Source: Bank of Mexico, available at www.banxico.org.mx.

	GDP		**Consumption**		**Investment**	
	Growth (data): 19.1% Growth (model): 0.7%		Growth (data): 20.5% Growth (model): 0.7%		Growth (data): 28.2% Growth (model): 1.3%	
Trend growth (% p.y.)	Above trend growth (%)	Model share (%)	Above trend growth (%)	Model share (%)	Above trend growth (%)	Model share (%)
1.0%	11.1%	6.3%	12.4%	2.4%	19.6%	6.6%
1.5%	7.3%	9.6%	8.6%	3.5%	15.5%	8.4%
2.0%	3.7%	19.0%	4.9%	6.1%	11.6%	11.2%
2.5%	0.2%	359.5%	1.4%	21.9%	7.9%	16.6%

Table A.2: Credit expansion experiment for Mexico: comparison with the real data between 2006-2013. Growth rates of GDP, consumption and investment are obtained from Bank of Mexico, available at www.banxico.org.mx.

B
Households' fixed level of human capital: mean preserving exercise

For the households' fixed level of human capital θ, we also consider a mean preserving exercise. Therefore, we change θ_1 and θ_2 to generate economies with different levels of income inequality, but the aggregate labor units is kept at the same level as in the benchmark. In this case, a lower (higher) θ_1 and a higher (lower) θ_2 increase (reduce) income inequality. Specially, for economies with higher levels of income inequality than the benchmark, we set θ_1 equals to 0.8 (0.7) and θ_2 equals to 2.28 (2.5) that corresponds to a Gini index of 47 (52). Moreover, for economies with lower levels of income inequality, we consider θ_1 equals to 1.3 (1.5) and θ_2 equals to 1.17 (0.72) to generate economies with a Gini index of 37 (32).

Figure B.1 shows these results, which are not in line with the empirical finding: consumption growth at peak is higher in economies with lower income inequality. However, this result is not monotonic. For example, the economy with Gini index of 32 shows a consumption growth at the peak of 1.81 percent, while the other one with Gini index of 37 is 1.73 percent. On the other hand, consumption grows at the peak 1.73 and 1.63 in the economies with Gini index of 47 and 52, respectively. This occurs because the results on this exercise depends on the share and on the values of θ_1 and θ_2, which influence how households are credit constrained in equilibrium.

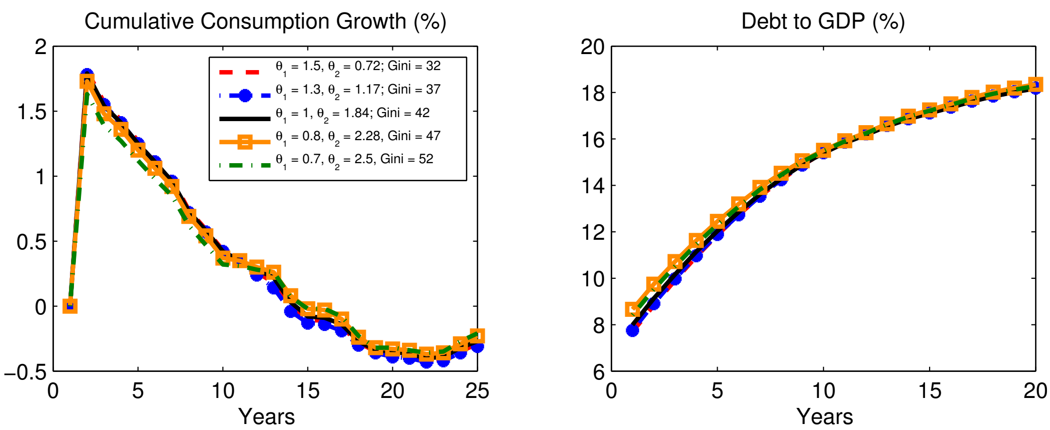

Figure B.1: σ_ε^2: Cumulative consumption growth ad Debt to GDP by income inequality.

Table B.1 decomposes these results in the contribution of assets distribution and policies functions. Note that, when we consider the assets distributions fixed, results are monotonic: consumption at the peak grows 1.86 percent in the economy with highest income inequality and grows 1.65 percent in the one with lowest income inequality. Also considering the contribution for the assets distribution, results are not monotonic: there is a negative contribution for economies with Gini index of 32, 37 and 52, however, there is a positive contribution for the one with Gini index of 47.

Consumption growth at peak (%)					
θ_1	θ_2	**Gini Index**	**Total**	**Policy**	**Distribution**
1.5	0.72	32	1.81	1.86	-0.04
1.3	1.17	37	1.73	1.81	-0.08
1	**1.84**	**42**	**1.79**	**-**	**-**
0.8	2.28	47	1.73	1.70	0.03
0.7	2.5	52	1.63	1.65	-0.02

Table B.1: θ_1 and θ_2 - Mean preserving spread: Policy functions and Assets distribution.

C
Predictor Means for the Synthetic Estimates

	Figure 3.4(a)		Figure 3.5(a)		Figure 3.6		Figure 3.7(a)	
Variable	Treated	Synthetic	Treated	Synthetic	Treated	Synthetic	Treated	Synthetic
$\Delta log(BRL)$	1.257	0.913	1.1257	0.938	1.257	0.528	0.364	0.252
$\Delta log(Volatility)$	0.379	0.737	0.379	0.371	0.379	0.364	-0.066	-0.171
$\Delta(RiskReversal)$					0.034	0.040		
$\Delta log(EquityIndex)$	-0.671	-1.687	-0.671	-0.747	-0.671	-0.692	-0.587	-0.587
$\Delta log(BondIndex)$	-0.993	-0.986	-0.993	-0.399	-0.993	-0.562	-0.268	-0.030
$CapitalFlows/GDP$	-0.002	-0.002	-0.002	-0.001	-0.002	-0.002	-0.002	-0.001
	Figure 3.8(a)		Figure 3.9		Figure 3.10		Figure 3.11	
Variable	Treated	Synthetic	Treated	Synthetic	Treated	Synthetic	Treated	Synthetic
$\Delta log(BRL)$	0.364	0.364	0.364	0.095	-0.277	-0.266	0.692	0.686
$\Delta log(Volatility)$	-0.066	-0.066	-0.066	-0.028	-0.096	-0.181	-0.152	0.301
$\Delta(RiskReversal)$			0.001	-0.003				
$\Delta log(EquityIndex)$	-0.587	-0.586	-0.587	-0.280	1.171	0.884	-0.019	-0.019
$\Delta log(BondIndex)$	-0.268	0.026	-0.268	-0.175	0.393	0.411	0.260	0.259
$CapitalFlows/GDP$	-0.002	-0.001	-0.002	-0.001	0.001	0.001	-0.001	-0.001

Table C.1: Notes: Treatment corresponds to the means for Brazil, and Synthetic to the means for its synthetic estimates in the Figure indicated by the different columns. For example, the results under the Figure 3.4(a) heading correspond to the means and synthetic for the log change in the exchange rate in the sample around the program announcement. For ease of illustration, variables are scaled to 100 times the log change in the exchange rate, equity and bond indices, and volatility, risk reversal and capital flows are measured in percentage terms.