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PhD thesis

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Essays on Labor and Exchange Markets in Chile

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Acknowledgments

This PhD thesis is the outcome of a journey that I started in 2012, when I decided to continue my academic formation at the Department of Economics at the University of Copenhagen in Denmark. Journeys are not always easy and mine was no exception. This, nevertheless, made it unique and unforgettable. Without the encouragement and support of many people, this would not have been possible.

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English summary

This thesis consists of four self-contained articles presented in four separate chapters dealing with persistence in Chilean macroeconomic data. Generally, the results provide empirical support for the predictions of the structural slumps theory (Phelps, 1994) in a world of imperfect knowledge (Frydman and Goldberg, 2007, 2011).

The first chapter studies real exchange rate persistence. The results show that long and persistent swings in the real exchange rate are compensated by similar movements in the interest rate spread, which restores the equilibrium in the product market when the real exchange rate moves away from its long-run benchmark value. Fluctuations in the copper price are also associated with real exchange rate swings.

The second chapter analyzes the Phillips curve and the effect of monetary policy on unemployment. The results show that the natural rate of unemployment is positively associated with the interest rate, suggesting that monetary policy might not be completely neutral over the business cycle. Additionally, trend-adjusted productivity is positively co-moving with the unemployment rate and negatively co-moving with the real exchange rate. This suggests that in periods of real appreciation, firms improve productivity by laying off the least-productive workers.

The third chapter studies the effect of the minimum wage on the employment of household workers. The results show that there is a negative and inelastic long-run relationship between hours worked and the minimum wage, but no relationship between the minimum wage and the number of workers employed. This suggests that employers in the household service sector have reduced the number of hours worked per employee instead of the overall number of employees when there has been a minimum wage increase.

The fourth chapter analyzes the effect of the monetary policy on the labor market when workers' heterogeneity is explicitly considered (age, income quintile, and economic sector). The results indicate that monetary policy does not evenly affect all workers. The primary sector is the least sensitive and the secondary sector the most sensitive to contractionary monetary shocks.

Danish summary

Denne afhandling består af fire selvstændige artikler, der er præsenteret i fire kapitler. Alle kapitlerne omhandler persistens i makroøkonomisk data fra Chile. Resultaterne er generelt i overensstemmelse med forudsigelserne fra "structural slumps" theory (Phelps, 1994) i en verden karakteriseret ved ufuldkommen viden (Frydman og Goldberg, 2007, 2011).

Det første kapitel undersøger persistens i reale valutakurser. Resultaterne viser, at lange og persistente udsving i den reale valutakurs modsvares af tilsvarende bevægelser i rentespændet, hvilket genetablerer ligevægten på varemarkedet, når den reale valutakurs bevæger sig væk fra sin langsigtede benchmark-værdi. Udsving i kobberprisen er ligeledes knyttede til udsving i den reale valutakurs.

Det andet kapitel analyserer Phillips-kurven og effekten af pengepolitik på arbejdsløshed. Resultaterne viser, at arbejdsløshedens naturlige niveau er positivt forbundet med renten, hvilket indikerer, at pengepolitikken måske ikke er fuldkommen neutral henover konjunkturcyklen. Hertil kommer, at arbejdsløsheden samvarierer positivt med den trend-justerede produktivitet, som samtidig samvarierer negativt med den reale valutakurs. Dette antyder, at i perioder med real appreciering, virksomheder forbedrer produktiviteten ved at fyre den mindst produktive del af arbejdsstyrken.

Det tredje kapitel omhandler effekten af minimumslønninger på hushjælpes beskæftigelse. Resultaterne viser, at der er et negativt og uelastisk langsigtsforhold mellem det arbejdede timetal og minimumslønnen, men ingen sammenhæng mellem minimumslønnen og antallet af beskæftigede hushjælpe. Dette indikerer, at arbejdsgiverne i hushjælpsbranchen har skåret i timetallet per ansat frem for i selve antallet af medarbejdere, når minimumslønnen er blevet hævet.

Det fjerde kapitel analyserer virkningen af pengepolitik for arbejdsmarkedsudfald, når der eksplicit tages højde for arbejderheterogenitet (alder, indkomstkvintil og økonomisk sektor). Resultaterne tyder på, at pengepolitik ikke påvirker alle arbejdere på samme måde. De primære sektor er de mindst følsomme og de sekundære sektor de mest følsomme over for kontraktive pengepolitiske stød.

Introduction

Persistence in time series has puzzled economists for a long time. This thesis uses two different econometric approaches to try to understand such persistence in Chilean macroeconomic data. To address this topic, the first three chapters use the cointegrated vector autoregressive (CVAR) model developed by Johansen (1996) and the last chapter uses the factor-augmented vector autoregressive (FAVAR) model proposed by Bernanke et al. (2005). Generally, the results provide empirical support for the structural slumps theory, developed by Phelps (1994), in a world of imperfect knowledge economics (IKE), which is a theoretical framework developed by Frydman and Goldberg (2007, 2011).

Under IKE, nominal interest rates exhibit strong persistence due to a nonstationary uncertainty premium, which in foreign exchange markets is related to the purchasing power parity (PPP) gap. Inflation rates are, however, more stable over time due to international competitiveness. Thus, the Fisher effect does not hold as a stationary condition. Furthermore, due to speculative behavior in the currency market, nominal exchange rates undergo long and persistent swings around relative prices, causing the real exchange rate to behave as a near $I(2)$ process. These persistent deviations will be reflected in the uncertainty premium, and hence in the domestic and foreign interest rates. In addition, in a world of imperfect knowledge the interest rate spread and the uncertainty premium—both near $I(2)$—cointegrate to a near $I(1)$ relationship, which cointegrates with the near $I(1)$ relative inflation rate to produce a stationary market-clearing mechanism called the uncertainty adjusted uncovered interest parity (UAUIP).

The structural slumps theory predicts that fluctuations in the real exchange rate might affect domestic real interest rates, wages, unemployment, and output. One implication is that the natural rate of unemployment is a function of the real interest rate. Phelps (1994) assumes that the unemployment rate and the real interest rate are stationary, whereas empirical studies often find that the real interest rate is indistinguishable from a unit root process. Juselius and Juselius (2012) suggest that Phelps's assumption needs to be modified to account for persistence in the data by allowing for imperfect knowledge-based expectations rather than rational expectations. The rationale is as follows: In an IKE world, the Fisher effect does not hold as a stationary condition in the sense that the nominal interest rate does not change in tandem with changes in the inflation rate. This is likely to result in speculative capital inflows, causing a real exchange rate appreciation and a worsening of the domestic competitiveness. Under these circumstances, domestic firms in the tradable sector cannot generally count on exchange rates to restore competitiveness, and will thus be prone to adjust profits rather than prices. Profits can be adjusted through improvements in labor productivity by laying off the least productive part of the labor force. Therefore, the structural slumps theory in a world of imperfect knowledge predicts an increase in both labor productivity and unemployment in periods of real appreciation and rising real interest rates.

The first chapter—*Modeling Real Exchange Rate Persistence in Chile*—studies the long and persistent swings in the real exchange rate that last, on average, six or seven years. The empirical analysis, based on an $I(2)$ CVAR model, suggests that swings in the real exchange rate are mainly associated with compensatory movements in the interest rate spread. This is one of the main predictions from the IKE theory: that the interest rate spread will move in a compensatory manner to restore the equilibrium in the product market when the nominal exchange rate is away from its long-run benchmark value. Copper is the main export commodity in Chile, and the analysis suggests that the copper price co-moves along with the real exchange rate. The analysis also indicates that the $I(2)$ trend is generated by twice-cumulated shocks to the US price and that this trend loads positively into prices and the exchange rate. However, this long-run stochastic trend in relative prices and in the nominal exchange rate cancels out. The dynamic adjustment of domestic prices, the domestic interest rate, and the exchange rate shows a rich structure of error-correcting and error-increasing behavior that is consistent with the persistence observed in the data.

The second chapter—*The Phillips Curve and the Role of Monetary Policy in Chile*—is a study of the dynamics of inflation and unemployment. The empirical analysis show that these dynamics are described by a Phillips curve when allowing for a positive co-movement between trend-adjusted productivity and the unemployment rate. The results also show that the natural rate of unemployment is time varying and is positively associated with the interest rate, suggesting that monetary policy might not be completely neutral over the business cycle. Furthermore, the empirical analysis shows that trend-adjusted productivity is negatively related to the real exchange rate. Thus, periods of real appreciation and increasing interest rates are associated with higher unemployment rates and higher productivity. Altogether, these findings provide empirical support for the structural slumps theory in an IKE world.

The third chapter—*The Effect of the Minimum Wage on Employment and Hours Worked: The Case of Household Workers in Chile*—is inspired by the Chilean government policy in 2008 that increased the minimum wage of household workers. The policy, based on gradual and predictable rates, increased the minimum wage for household workers until it reached the same level as the national minimum wage. Under perfect competition, an increase in the minimum wage should be associated with lower levels of employment because employers substitute the now more expensive labor input with other inputs (e.g., capital). However, household workers' service is labor intensive and the substitution of labor for capital is generally low. Because the literature does not give a clear consensus about how to measure employment, this chapter considers both the number of hours worked and number of employed people as a measurement of employment. The results indicate that there is an inelastic and negative long-run relationship between hours worked and the minimum wage, but no relationship between the minimum wage and the number of employed household workers. This suggests that employers tend to reduce the number of hours worked per employee instead of the number of employees when there is a minimum wage increase. Additionally, a minimum wage increase seems to be associated with an increase in the total monthly income of household workers. Furthermore, and similar to the finding in Chapter 2, the results indicate a positive relationship between the national unemployment rate and trend-adjusted productivity.

Finally, the fourth chapter—*The Impact of Monetary Policy on Labor Market with Heterogeneous Workers: The Case of Chile*—examines how monetary policy affects labor markets with a special focus on the heterogeneity of different economic sectors and demographic groups. The results indicate that after a contractionary monetary shock, the secondary sector reacts most strongly in terms of both an increased job-separation rate (the probability that an employed worker will lose his or her job in the next three months) and a decreased job-finding rate (the probability that an unemployed worker will find a job within three months). Furthermore, for the primary and tertiary sector, it is mostly older workers (55 or older) who experience an increase in the job-separation rate. Altogether, the results in this chapter indicate that the primary sector is the least sensitive to contractionary monetary shocks. The fact these shocks increase the job-separation rate, which is the main driver of the unemployment rate, indicates that the interest rate might have real effects on the economy. This result is similar to the finding in the second chapter that monetary policy might not be completely neutral over the business cycle. The fourth chapter indicates, nevertheless, that not all groups of workers are evenly affected by the monetary policy. This chapter also presents empirical evidence of a positive correlation between productivity and unemployment in the primary and tertiary sectors in the economy.

Chapter 1

Modeling Real Exchange Rate Persistence in Chile

Leonardo Salazar*

Abstract

The long and persistent swings in the real exchange rate have for a long time puzzled economists. Recent models built on imperfect knowledge economics seem to provide a theoretical explanation for this persistence. Empirical results, based on a cointegrated vector autoregressive (CVAR) model, provide evidence of error-increasing behavior in prices, exchange rates, and interest rates, which is consistent with the persistence observed in the data. The movements in the real exchange rate are compensated by movements in the interest rate spread, which restores the equilibrium in the product market when the real exchange rate moves away from its long-run benchmark value. Fluctuations in the copper price also seem to explain the deviations of the real exchange rate from its long-run equilibrium value.

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1.1 Introduction

The purchasing power parity (PPP) theory establishes that identical goods will have the same price in different economies when prices are expressed in the same currency (Krugman et al., 2011). In other words, the aggregate relative prices between two countries should be equal to the nominal exchange rate between them (Taylor and Taylor, 2004).[1]

The PPP has been broadly used in economics, in both theoretical models and empirical applications. For instance, a number of general equilibrium models use the PPP as an equilibrium condition; that is, the PPP is assumed to hold over time, and the main results in these models rely on the PPP assumption (Duncan and Calderón, 2003). In addition, estimates of PPP exchange rates are used to compare national income levels, determining the degree of misalignment of the nominal exchange rate around relative prices and the appropriate policy response (Sarno and Taylor, 2002).

However, empirical evidence shows that over time, the nominal exchange rate exhibits long and persistent swings around relative prices. Specifically, while the ratio of domestic to foreign good prices changes slowly over time, the nominal exchange rate exhibits long and persistent swings away from its benchmark value. Consequently, these persistent swings are observed on the real exchange rate. See Figure 1.1 for the Chilean case.

Long and persistent fluctuations in the real exchange rate (RER) may generate allocative effects on the economy. Indeed, the competitiveness of a country might be negatively affected by a prolonged real appreciation (Mark, 2001). Furthermore, these fluctuations might affect domestic real interest rates, wages, unemployment, and output, generating structural slumps in economies (Phelps, 1994).

Heterogeneous agent models seem to be consistent with the persistence in the data. These are mostly behavioral models in which agents are boundedly rational and use different strategies, heuristics or rule of thumb, that perform reasonably well, but these strategies might not be perfect. Furthermore, these models might explain stylized facts in financial time series such as temporary bubbles, crushes, mean reversion, and excess volatility (Hommes, 2006). Given that agents are not endowed with perfect knowledge, these models might be called “imperfect knowledge” models (Juselius and Assenmacher, 2015; Juselius, 2015).

Frydman and Goldberg (2007) developed a monetary model based on imperfect knowledge economics (IKE), known as IKE-based model, that was proposed as a solution to the puzzle of the long swings in exchange rates. Its empirical validity has been tested by Johansen et al. (2010), Juselius (2015), and Juselius and Assenmacher (2015). For instance, using a cointegrated vector autoregressive (CVAR) scenario,[2] Juselius (2015) argues, based on German-US data, that the IKE-based scenario is empirically supported

[1] This concept is known as absolute PPP.

[2] A CVAR scenario tests the empirical consistency of the basic underlying assumptions of a model rather than imposing them on the data from the outset (Juselius, 2015).

by every testable hypothesis that describes the underlying assumptions of this model.

The evidence on PPP is generally mixed and the results depend on the covered period, the variables included in the analysis, and the econometric methodology used to test the PPP hypothesis.[3]

In the case of Chile, the evidence is also mixed, and the results depend primarily on the methodology used to test the PPP hypothesis. On the one hand, when augmented Dickey-Fuller (ADF) test is used in a single equation that includes the nominal exchange rate, domestic price, and foreign price, the PPP hypothesis seems to hold. That is, RER is found to be a stationary process (Délano, 1998; Duncan and Calderón, 2003). On the other hand, if multivariate cointegration techniques are used, the results show that RER behaves as a nonstationary $I(1)$ process. However, it cointegrates with other $I(1)$ variables to a stationary process. Indeed, there is evidence of cointegration between RER, productivity, net foreign assets, government expenditures, and terms of trade (Céspedes and De Gregorio, 1999) and between RER and black exchange rates (parallel market) (Diamandis, 2003). It also seems that the stationarity of RER depends on the analyzed period; for instance, Délano (1998) shows that RER behaves as an $I(0)$ process when the period 1830–1995 is considered but as an $I(1)$ process in the period 1918–1995.

The Chilean economy, similar to other economies in South America, depends strongly on its commodities' prices. Copper is the main export commodity in Chile; it accounted for 54% of Chile's exports, 14% of fiscal revenue, and 13% of nominal GDP in 2012 (Wu, 2013). Chile has become increasingly important in the world copper market because its share of global production has increased to somewhat more than a third since the late 1960s (De Gregorio et al., 2011).

A number of studies have analyzed how copper prices affect the Chilean economy through its effects on nominal exchange rates, terms of trade, and business cycles. The results suggest that a positive shock to the copper price leads to appreciation in nominal and real exchange rates, output expansion, and an increased inflation rate (Cowan et al., 2007; Medina and Soto, 2007).

In the long run, copper prices appear to explain most of the fluctuations in the Chilean peso, but in the short run, other factors, including interest rate spread, global financial risk, and local pension funds foreign exchange derivative position, may explain these fluctuations (Wu, 2013). The fact that RER has acted as a shock absorber due to the flexible exchange rate regime, a rule-based fiscal policy, and a flexible inflation targeting system might explain why the Chilean economy has become increasingly resilient to copper price shocks in the last 25 years (De Gregorio et al., 2011).

This chapter finds, based on the estimation of a CVAR model, that the long and persistent swings in the real exchange rate are compensated by movements in the interest

[3] Duncan and Calderón (2003), and Froot and Rogoff (1995) present a thorough review of the literature on PPP testing.

rate spread, which restores the equilibrium in the product market when the real exchange rate moves away from its long-run benchmark value. Fluctuations in the copper price also seem to explain the deviations of the real exchange rate from its long-run equilibrium value. Additionally, the results indicate error-increasing behavior in prices, nominal exchange rates, and interest rates, which is consistent with the persistence in the data.

1.2 Theoretical framework

1.2.1 Parity conditions

The law of one price establishes that:

$$P^j_{d,t} = S_t P^j_{f,t} \tag{1.1}$$

where P^j_d is the domestic-currency price of good j, S is the nominal exchange defined as the domestic-currency price in a unit of foreign currency, and P^j_f is the foreign-currency price of good j. Aggregating over the price of tradable goods, equation (1.1) is expressed as:

$$P_{d,t} = S_t P_{f,t} \tag{1.2}$$

where P_d is the weighted average of domestic prices of the basket and P_f is the weighted average of foreign prices of the same basket. Equation (1.2) indicates that once the price of the same basket of tradable goods is expressed in the same currency, a unit of the domestic currency should buy the same basket of goods that a unit of the foreign currency buys. Equation (1.2) is known as the absolute PPP (Rogoff, 1996).

If p_d, p_f and s are, respectively, the natural logarithm of P_d, P_f, and S, equation (1.2) can be rewritten as:

$$p_{d,t} = p_{f,t} + s_t \tag{1.3}$$

and the long-run PPP condition is expressed as:

$$p_{d,t} - p_{f,t} - s_t = \mu + ppp_t \tag{1.4}$$

where μ is a constant that reflects differences both in units of measure and in base-year normalization of price indices (Mark, 2001), and ppp_t is a stationary error term that represents the deviations from PPP.[4] If the PPP condition holds, then by definition, the real exchange rate behaves as a stationary process, that is:

[4]In empirical testing, the PPP condition is normally replaced by $s_t = \mu + \gamma_1 p_{d,t} + \gamma_2 p_{f,t} + ppp_t$, where $\gamma_1 = -\gamma_2 = 1$ is expected.

$$q_t = s_t + p_{f,t} - p_{d,t} \sim I(0) \tag{1.5}$$

Moreover, deviations from the uncovered interest parity (UIP) condition, that is, the excess returns on foreign exchange, er_t, would be stationary,[5] so that:

$$er_t = (i_d - i_f) - \left(s^e_{t+1} - s_t\right) \sim I(0) \tag{1.6}$$

where i_d and i_f are, respectively, the domestic and foreign interest rates and the superscript e denotes an expected value.

Empirical evidence finds, however, that the real exchange rates and excess returns behave as nonstationary processes, suggesting that the assumptions behind equations (1.5) and (1.6) are untenable when using real data (see Juselius, 2010, 2015; Juselius and Assenmacher, 2015; Johansen et al., 2010; and Frydman and Goldberg, 2007).

1.2.2 Persistence in the data

This subsection[6] presents a theoretical framework, developed in Juselius and Assenmacher (2015) and based on IKE, that is consistent with the long and persistent swings in the real exchange rate. The model assumes that the nominal exchange rate is mainly driven by the expected level of prices, that is:

$$s_t = B_0 + B_{1,t}\left(p_d - p_f\right)^e_{t+1} + \nu^e_{t+1} \tag{1.7}$$

where $\nu^e_t \sim I(0)$ is an error term that captures changes in interest rates and income. The superscript e stands for an expected value, B_0 is a constant term, and $B_{1,t}$ is a time-varying coefficient that represents the weight to relative prices in financial actors' forecasts. Generally, the weight depends on how far the nominal exchange is from its long-run benchmark value. Based on (1.7), changes in the nominal exchange can be expressed as:

$$\triangle s_t = B_{1,t}\triangle\left(p_d - p_f\right)^e_{t+1} + \triangle B_{1,t}\left(p_d - p_f\right)^e_{t+1} + \triangle\nu^e_{t+1}. \tag{1.8}$$

To apply equations (1.7) and (1.8) to actual data, the expected values must be associated with actual values. Juselius (2015) provides two assumptions that allow using actual instead of expected values in the cointegrating analysis. First, the forecast error of the inflation differential is stationary, that is:

[5]If deviations from PPP are assumed to be near $I(1)$, the deviations from UIP also behave as nonstationary, near-$I(1)$ processes.

[6]This subsection is based mainly on Juselius (2010, 2015), Juselius and Assenmacher (2015), and Frydman and Goldberg (2007, 2011, 2013).

$$\triangle\left(p_d-p_f\right)^e_{t+1}-\triangle\left(p_d-p_f\right)_{t+1}\sim I(0) \tag{1.9}$$

which basically states that agents' forecasts cannot permanently deviate from actual values. Second, the relative rate of inflation between two economies is at most $I(1)$, so that:

$$\triangle\left(p_d-p_f\right)_{t+1}-\triangle\left(p_d-p_f\right)_{t}\sim I(0). \tag{1.10}$$

Under assumptions (1.9) and (1.10), the cointegration properties are robust to using actual instead of expected values. Then, based on the above assumptions, equation (1.8) can be expressed as:

$$\triangle s_t = B_{1,t}\triangle\left(p_d-p_f\right)_t+\triangle B_{1,t}\left(p_d-p_f\right)_t+\triangle\nu_t. \tag{1.11}$$

Juselius and Assenmacher (2015) argue that $\triangle B_{1,t}$ has to be large for $\triangle B_{1,t}\left(p_d-p_f\right)_t$ to have a marked effect on $\triangle s_t$. Thus, one can assume, as in Frydman and Goldberg (2007), that $\left|B_{1,t}\triangle\left(p_d-p_f\right)_t\right|\gg\left|\triangle B_{1,t}\left(p_d-p_f\right)_t\right|$, so that:

$$\triangle s_t \simeq B_{1,t}\triangle\left(p_d-p_f\right)_t+\triangle\nu_t. \tag{1.12}$$

Before estimating the above model using the CVAR, the issue of time-varying parameters must be addressed. Tabor (2014) simulates data for the process $y_t=\beta_t' x_t+\varepsilon_t$ where x_t is nonstationary $I(1)$, ε_t is a random walk and $\beta_t=\beta+Z_t$ where $Z_t=\varrho Z_{t-1}+\varepsilon_{Z,t}$ and $\varrho<1$. He shows that when a CVAR model is applied to the simulated data, the estimated cointegrated coefficient corresponds to $E\left(\beta_t\right)$. Hence, based on the results in Tabor (2014), one can argue that the CVAR model may be used to estimate average long-run relationships when the underlying data-generating process involves bounded-parameter instability.

Then, the change in the real exchange rate should behave as a near $I(1)$ process provided that $B_{1,t}=B+\rho B_{1,t-1}+\varepsilon_{B_1,t}$ with $\rho<1$, but close to one. Juselius (2014) argues that the latter behavior can be used to approximate the the change in the real exchange rate through the following process:

$$\triangle q_t = a_t+\nu_{q,t} \tag{1.13}$$

where $\nu_{q,t}\sim I(0)$ is an error term and the time-varying drift term, a_t, measures the appreciations or depreciations of the real exchange rate due to changes in individual forecasting strategies.[7] This drift is assumed to follow a mean zero stationary autoregressive process,

[7] This is consistent with the FG IKE-based model developed by Frydman and Goldberg (2007), which assumes that individuals recognize their imperfect knowledge about the underlying processes that drive outcomes. Thus, they use a multitude of forecasting strategies that are revised over time in a way that

so that:

$$a_t = \rho_t a_{t-1} + \nu_{a,t} \tag{1.14}$$

where $\nu_{a,t} \sim I(0)$ is an error term and ρ_t is a time-varying coefficient that is close to one when the real exchange rate is in the vicinity of its long-run benchmark value, and otherwise $\rho_t \ll 1$.[8] The average of this coefficient, $\bar{\rho}$, is generally close to one whenever the sample period is sufficiently long (Juselius, 2015). Then, a_t describes a persistent near $I(1)$ process, and modeling the real exchange rate as a near $I(2)$ process is consistent with swings of shorter and longer duration, implying that the length of these swings is not predictable (Frydman and Goldberg, 2007).

In the FG IKE-based model, UIP—the market clearing mechanism between the expected change in the nominal exchange rate and the nominal interest rate spread—is replaced by an uncertainty adjusted uncovered interest parity (UAUIP) condition, so that the excess returns under IKE are expressed as:

$$er_t^{IK} = (i_d - i_f)_t - \left(s_{t+1}^e - s_t\right) + up_t \tag{1.15}$$

where up_t stands for an uncertainty nonstationary premium, a measure of agents' loss averseness.[9] The interest rate spread corrected for the uncertainty premium is a minimum return that agents require to speculate in the foreign exchange market. This premium starts increasing when the nominal exchange rate moves away from its long-run benchmark value and decreases when the nominal exchange rate moves toward equilibrium. In the foreign exchange market, the uncertainty premium is related to the PPP gap (Frydman and Goldberg, 2007). Then, the excess returns under IKE are formulated as:

$$er_t^{IK} = (i_d - i_f)_t - \triangle s_{t+1}^e + f\left(p_{d,t} - p_{f,t} - s_t\right). \tag{1.16}$$

This equation suggests that in a world of imperfect knowledge, the expected change in the nominal exchange rate may not be directly related to the interest rate spread, but to the spread corrected by the PPP gap.

cannot be fully prespecified. Indeed, given the diversity of forecasting strategies, this model assumes two kinds of individuals in the foreign currency market: bulls, who speculate on the belief that the asset price will rise, and bears, who speculate on its fall.

[8] When periods where a_t is far from its benchmark value are shorter compared with the near vicinity periods, it describes a persistent but mean-reverting process.

[9] Frydman and Goldberg (2007) extend the concept of loss aversion given by Kahneman and Tversky (1979) to the concept of endogenous loss aversion, which says that the greater the potential loss, the higher the degree of loss aversion. This definition establishes that the UAUIP equilibrium exists.

1.2.3 Theory-consistent CVAR scenario results

A consequence of the UAUIP condition is that both domestic and foreign interest rates are affected by the uncertainty premium. Juselius (2015) suggests the following data-generating process to describe changes in the interest rate:

$$\triangle i_{j,t} = \omega_{j,t} + \nu_{j,t} \tag{1.17}$$

where $\nu_{j,t} \sim I(0)$ is an unanticipated interest rate shock and $j = d, f$. The term $\omega_{j,t}$ stands for changes in the domestic uncertainty premium and is assumed to follow a mean zero stationary autoregressive process:

$$\omega_{j,t} = \rho^{\omega}_{j,t}\omega_{j,t-1} + \nu^{\omega}_{j,t} \tag{1.18}$$

where $\nu^{\omega}_{j,t}$ is a stationary error term. The time-varying autoregressive coefficient, $\rho^{\omega}_{j,t}$, is assumed to be almost on the unit circle when the nominal exchange rate is in the vicinity of its long-run benchmark value—the relative price—otherwise the coefficient is strictly less than one. Nevertheless, $\bar{\rho}^{\omega}_{j} \approx 1$ provided that periods where the coefficient is close to one are much longer than otherwise. When $\bar{\rho}^{\omega}_{j} \approx 1$, (1.18) describes a near $I(1)$ domestic uncertainty premium. Consequently, under IKE, the interest rate change behaves as a persistent near $I(1)$ process, implying that nominal interest rates are near $I(2)$.

Using a CVAR scenario, Juselius (2015) demonstrates that the following hypotheses are consistent with IKE:

$$s_t \sim \text{near}\, I(2) \tag{1.19}$$

$$(p_{d,t} - p_{f,t}) \sim \text{near}\, I(2) \tag{1.20}$$

$$(i_{d,t} - i_{f,t}) \sim \text{near}\, I(2) \tag{1.21}$$

$$(s_t + p_{f,t} - p_{d,t}) \sim \text{near}\, I(2) \tag{1.22}$$

$$(i_{j,t} - \triangle p_{j,t}) \sim \text{near}\, I(2) \tag{1.23}$$

$$\{(i_{d,t} - i_{f,t}) - c_1\,(s_t + p_{f,t} - p_{d,t})\} \sim I(1) \tag{1.24}$$

$$\{(\triangle p_{d,t} - \triangle p_{f,t}) - c_2\,(i_{d,t} - i_{f,t}) + c_3\,(s_t + p_{f,t} - p_{d,t})\} \sim I(0) \tag{1.25}$$

where c_i are constant coefficients. The cointegrating relationship (1.25) corresponds to the excess returns under IKE. This indicates that the interest rate differential and the uncertainty premium, both near $I(2)$, cointegrate to a near $I(1)$ relationship and that the interest rate rate spread corrected by the uncertainty premium cointegrates with the near $I(1)$ relative inflation rate to produce a stationary market-clearing mechanism.

1.3 The CVAR model and the $I(2)$ representation

The error correction form of a VAR model is expressed as:[10]

$$\triangle \boldsymbol{x}_t = \boldsymbol{\Pi} \boldsymbol{x}_{t-1} + \sum_{i=1}^{k-1} \boldsymbol{\Gamma}_i \triangle \boldsymbol{x}_{t-i} + \boldsymbol{\Phi} \boldsymbol{D}_t + \boldsymbol{\mu}_0 + \boldsymbol{\mu}_1 t + \boldsymbol{\varepsilon}_t \tag{1.26}$$

where $\boldsymbol{x}'_t = [x_{1,t}, x_{2,t}, \ldots, x_{p,t}]$ is a p-dimensional vector of stochastic variables, $\boldsymbol{\mu}_0$ is an unrestricted constant, t is an unrestricted trend with coefficient matrix $\boldsymbol{\mu}_1$, $\boldsymbol{D}_t$ is a matrix of deterministic terms (shift dummies, seasonal dummies), and $\boldsymbol{\varepsilon}_t \sim \mathcal{N}_p(\mathbf{0}, \boldsymbol{\Omega})$. Matrices $\boldsymbol{\Gamma}_i$ and $\boldsymbol{\Pi}$ have dimension $p \times p$.

If $\boldsymbol{\Pi}$ has reduced rank, $0 < r < p$, it can be decomposed into $\boldsymbol{\Pi} = \boldsymbol{\alpha}\boldsymbol{\beta}'$, where $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$ are $p \times r$ matrices of full column rank. The orthogonal complement of matrix $\boldsymbol{z}$ is denoted as $\boldsymbol{z}_\perp$, and $\overline{\boldsymbol{z}} := \boldsymbol{z}(\boldsymbol{z}'\boldsymbol{z})^{-1}$. Now, if $\boldsymbol{\Pi}$ has reduced rank and $\boldsymbol{\alpha}'_\perp \boldsymbol{\Gamma} \boldsymbol{\beta}_\perp$ is of full rank, where $\boldsymbol{\Gamma} = \mathbf{I}_p - \sum_{i=1}^{k-1} \boldsymbol{\Gamma}_i$, then the process $\boldsymbol{x}_t \sim I(1)$.[11] The r cointegrated relationships $\boldsymbol{\beta}'\boldsymbol{x}_t$ define stationary relationships among nonstationary variables, potentially representing long-run equilibrium relationships (Juselius, 2006).

Structuring the $I(2)$ representation of the CVAR model is a bit more complicated, and additional definitions must be given. The $I(2)$ model is defined by the two following reduced rank restrictions:

$$\begin{aligned} \boldsymbol{\Pi} &= \boldsymbol{\alpha}\boldsymbol{\beta}' \\ \boldsymbol{\alpha}'_\perp \boldsymbol{\Gamma} \boldsymbol{\beta}_\perp &= \boldsymbol{\xi}\boldsymbol{\eta}' \end{aligned} \tag{1.27}$$

where $\boldsymbol{\xi}$ and $\boldsymbol{\eta}$ are $(p-r) \times s_1$ matrices, s_1 is the number of $I(1)$ trends, or unit root processes, and it is such that $p - r = s_1 + s_2$, where s_2 is the number of $I(2)$ trends, or double unit root processes, in vector $\boldsymbol{x}_t$. Whereas the first rank condition in (1.27) is associated with the variables in levels, the second rank condition is related to the differentiated variables.

$\boldsymbol{\beta}_\perp$ and $\boldsymbol{\alpha}_\perp$ can, respectively, be decomposed into $\boldsymbol{\beta}_\perp = [\boldsymbol{\beta}_{\perp 1}, \boldsymbol{\beta}_{\perp 2}]$ and $\boldsymbol{\alpha}_\perp = [\boldsymbol{\alpha}_{\perp 1}, \boldsymbol{\alpha}_{\perp 2}]$. Matrices $\boldsymbol{\alpha}_{\perp 1} = \overline{\boldsymbol{\alpha}_\perp}\boldsymbol{\eta}$ and $\boldsymbol{\beta}_{\perp 1} = \overline{\boldsymbol{\beta}_\perp}\boldsymbol{\eta}$ are of dimension $p \times s_1$. Matrices $\boldsymbol{\alpha}_{\perp 2} = \boldsymbol{\alpha}_\perp \boldsymbol{\xi}_\perp$

[10] This section is based mainly on Juselius (2006), Dennis (2006), and Johansen et al. (2010).

[11] If $\mathrm{rank}(\boldsymbol{\Pi}) = p$, then the process $\boldsymbol{x}_t \sim I(0)$. If $\mathrm{rank}(\boldsymbol{\Pi}) = 0$, the same process is nonstationary and not cointegrated.

and $\boldsymbol{\beta}_{\perp 2} = \boldsymbol{\beta}_{\perp}\boldsymbol{\eta}_{\perp}$ have dimension $p \times s_2$.

Using the Johansen (1997) parametrization, model (1.26) can be written in second differences as follows:

$$\begin{aligned}\triangle^2 \boldsymbol{x}_t =& \boldsymbol{\alpha}\left(\boldsymbol{\rho}'\boldsymbol{\tau}'\boldsymbol{x}_{t-1} + \boldsymbol{d}'\triangle \boldsymbol{x}_{t-1}\right) + \boldsymbol{\zeta}'\boldsymbol{\tau}'\triangle \boldsymbol{x}_{t-1} + \sum_{i=1}^{k-2} \boldsymbol{\Lambda}_i \triangle^2 \boldsymbol{x}_{t-i} + \\ & \boldsymbol{\Phi}\boldsymbol{D}_t + \boldsymbol{\mu}_0 + \boldsymbol{\mu}_1 t + \boldsymbol{\varepsilon}_t\end{aligned} \tag{1.28}$$

where $\boldsymbol{\rho} = [\mathbf{I}, \mathbf{0}]'$, $\boldsymbol{\tau} = [\boldsymbol{\beta}, \boldsymbol{\beta}_{\perp 1}]$, $\boldsymbol{d}' = -\left(\left(\boldsymbol{\alpha}'\boldsymbol{\Omega}^{-1}\boldsymbol{\alpha}\right)^{-1}\boldsymbol{\alpha}\boldsymbol{\Omega}^{-1}\boldsymbol{\Gamma}\right)\boldsymbol{\tau}_{\perp}\left(\boldsymbol{\tau}'_{\perp}\boldsymbol{\tau}_{\perp}\right)^{-1}\boldsymbol{\tau}'_{\perp}$, $\boldsymbol{\zeta} = [\boldsymbol{\zeta}_1, \boldsymbol{\zeta}_2]'$ is a matrix of medium-run adjustment, and $\boldsymbol{\Lambda}_i = -\sum_{j=i+1}^{k-1} \boldsymbol{\Gamma}_j$.

In this model, the term in $(\cdot)$ represents the long-run equilibrium or polynomially cointegrating relationships. The term $\boldsymbol{\zeta}'\boldsymbol{\tau}'\triangle \boldsymbol{x}_{t-1}$ can be interpreted as a medium-run equilibrium relationship, defining the $r + s_1$ relationship that needs to be differentiated to become stationary.

The moving average (MA) representation of the $I(2)$ model is expressed as:

$$\begin{aligned}\boldsymbol{x}_t =& \boldsymbol{C}_2 \sum_{i=1}^{t}\sum_{s=1}^{i} \underbrace{\left(\boldsymbol{\varepsilon}_s + \boldsymbol{\Phi}\boldsymbol{D}_s + \boldsymbol{\mu}_0 + \boldsymbol{\mu}_1 s\right)}_{\boldsymbol{\varepsilon}_s} + \boldsymbol{C}_1 \sum_{i=1}^{t} \underbrace{\left(\boldsymbol{\varepsilon}_i + \boldsymbol{\Phi}\boldsymbol{D}_i + \boldsymbol{\mu}_0 + \boldsymbol{\mu}_1 i\right)}_{\boldsymbol{\varepsilon}_i} + \\ & \boldsymbol{C}^*(L)\left(\boldsymbol{\varepsilon}_t + \boldsymbol{\Phi}\boldsymbol{D}_t + \boldsymbol{\mu}_0 + \boldsymbol{\mu}_1 t\right) + \boldsymbol{A} + \boldsymbol{B}t\end{aligned} \tag{1.29}$$

where $\boldsymbol{C}_2 = \boldsymbol{\beta}_{\perp 2}\left(\boldsymbol{\alpha}'_{\perp 2}\boldsymbol{\Theta}\boldsymbol{\beta}_{\perp 2}\right)^{-1}\boldsymbol{\alpha}'_{\perp 2}$, $\boldsymbol{\beta}'\boldsymbol{C}_1 = -\overline{\boldsymbol{\alpha}}'\boldsymbol{\Gamma}\boldsymbol{C}_2$, $\boldsymbol{\beta}'_{\perp 1}\boldsymbol{C}_1 = -\overline{\boldsymbol{\alpha}_{\perp 1}}'\left(\mathbf{I}_p - \boldsymbol{\Theta}\boldsymbol{C}_2\right)$, and $\boldsymbol{\Theta} = \boldsymbol{\Gamma}\overline{\boldsymbol{\beta}}\overline{\boldsymbol{\alpha}}\boldsymbol{\Gamma} + \left(\mathbf{I}_p - \sum_{i=1}^{k-2}\boldsymbol{\Lambda}_i\right)$. $\boldsymbol{A}$ and $\boldsymbol{B}$ are functions of both the initial values and the model parameters (Johansen, 1992).[12]

Matrix $\boldsymbol{C}_2$ can be expressed as $\boldsymbol{C}_2 = \breve{\boldsymbol{\beta}}_{\perp 2}\boldsymbol{\alpha}'_{\perp 2}$, where $\breve{\boldsymbol{\beta}}_{\perp 2} = \boldsymbol{\beta}_{\perp 2}\left(\boldsymbol{\alpha}'_{\perp 2}\boldsymbol{\Theta}\boldsymbol{\beta}_{\perp 2}\right)^{-1}$, so that $\boldsymbol{\alpha}'_{\perp 2}\sum_{i=1}^{t}\sum_{s=1}^{i}\boldsymbol{\varepsilon}_s$ can be interpreted as the measure of the s_2 trends which load into the variables in $\boldsymbol{x}_t$ with the weights $\breve{\boldsymbol{\beta}}_{\perp 2}$ (Juselius, 2006).

The likelihood ratio test for the joint hypothesis of r cointegrating relationships and s_1 and s_2 trends, labeled $\mathcal{H}(r, s_1, s_2)$, versus $\mathcal{H}(p)$ is given by:

$$-2\log Q\left(\mathcal{H}(r, s_1, s_2) \mid \mathcal{H}(p)\right) = -T\log\left|\tilde{\boldsymbol{\Omega}}^{-1}\hat{\boldsymbol{\Omega}}\right| \tag{1.30}$$

where $\tilde{\boldsymbol{\Omega}}$ and $\hat{\boldsymbol{\Omega}}$ are, respectively, the covariance matrices estimated under $\mathcal{H}(r, s_1, s_2)$ and $\mathcal{H}(p)$.[13]

[12] From the MA representation (1.29), it follows that the unrestricted constant, $\boldsymbol{\mu}_0$, cumulates once to a linear trend and twice to a quadratic trend. In addition, the unrestricted trend, $\boldsymbol{\mu}_1$, cumulates once to a quadratic trend and twice to a cubic trend. To avoid the latter, quadratic and cubic trends have been restricted to zero in the subsequent analysis. For further information, see Chapter 17 in Juselius (2006).

[13] The distribution of the this is found in Johansen (1995) provided that model (1.28) does not restrict deterministic components; otherwise see Rahbek et al. (1999).

1.4 Stylized facts

Figure 1.1 Panel (a) shows the evolution of the natural logarithm (log) of the nominal exchange rate, measured as Chilean pesos (CLP) per US dollar (USD) and the log of the relative prices, measured as the ratio between the Chilean consumer price index (CPI) and the US CPI. Relative prices exhibit a positive but decreasing slope, reflecting the fact that from 1986 until 1999, Chilean prices were growing faster than US prices, but after 1999 the growth in relative prices decreased. This might be associated with the partial implementation of inflation targeting in Chile in 1990, which reduced annual inflation from 26% in 1990 to 3% in 1997. In the same panel, the nominal exchange rate undergoes long and persistent swings around relative prices, suggesting that PPP may hold only as a very long-run condition.

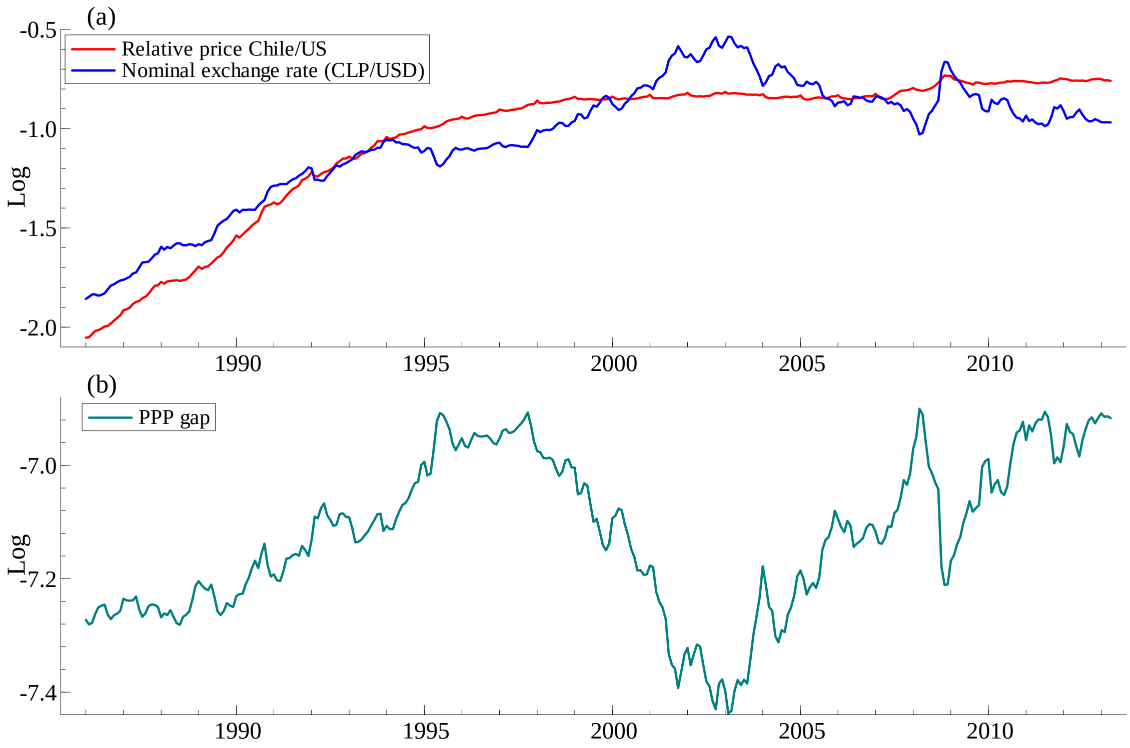
Figure 1.1 Panel (b) shows the PPP gap, defined as the difference between the log of relative prices and the log of the nominal exchange rate. The deviations exhibit long persistent swings, but it seems that the upward trend in relative prices is canceled by the upward trend in nominal exchange rate, so that the real exchange rate might fluctuate around a long-run constant level.

Figure 1.2 Panel (a) shows that relative inflation rates exhibit a high persistence, which is corroborated by the 12-month moving average. This persistence seems, however, to decrease steadily beginning in 1990, which may be associated with the implementation of inflation targeting in Chile in 1990. In the same figure, Panel (b) shows the changes in the nominal exchange rate, which seems stationary. Nevertheless, the 12-month moving average exhibits some persistence around the mean. It also appears that appreciations and depreciations are more volatile since 2000, which might be related to the free-floating exchange rate regime that was implemented by the Central Bank of Chile in September 1999. Panel (c), shows that changes in the PPP gap behave as a persistent but mean-reverting process. The 12-month moving average exhibits persistence around the mean that seems higher since 2000.

Figure 1.3 Panels (a) and (b) show, respectively, the Chilean interest rate and its first difference. The latter exhibits a large decrease in volatility since 2000. This might be associated with two major reforms that were introduced in the Chilean financial market between 2000 and 2001. While the first reform, promulgated in 2000, gave greater protection to both domestic and foreign investors, the second reform, enacted in 2001, liberalized the financial system, implying, among other things, capital account deregulation.

When the Chilean interest rate and its first difference are compared with their US counterparts, which are shown in Figure 1.3 Panels (c) and (d), an important difference in levels and volatility is noticeable. The Chilean interest rate has been historically higher than the US interest rate, and this seems to have changed since 2000. The latter is clearly reflected in the interest rate spread shown in Figure 1.3 Panel (e). The changes in the

Figure 1.1: Panel (a): Nominal exchange rate (CLP/USD) and relative prices (Chilean CPI/US CPI). Panel (b): Deviations from PPP. Monthly data 1986:1–2013:04

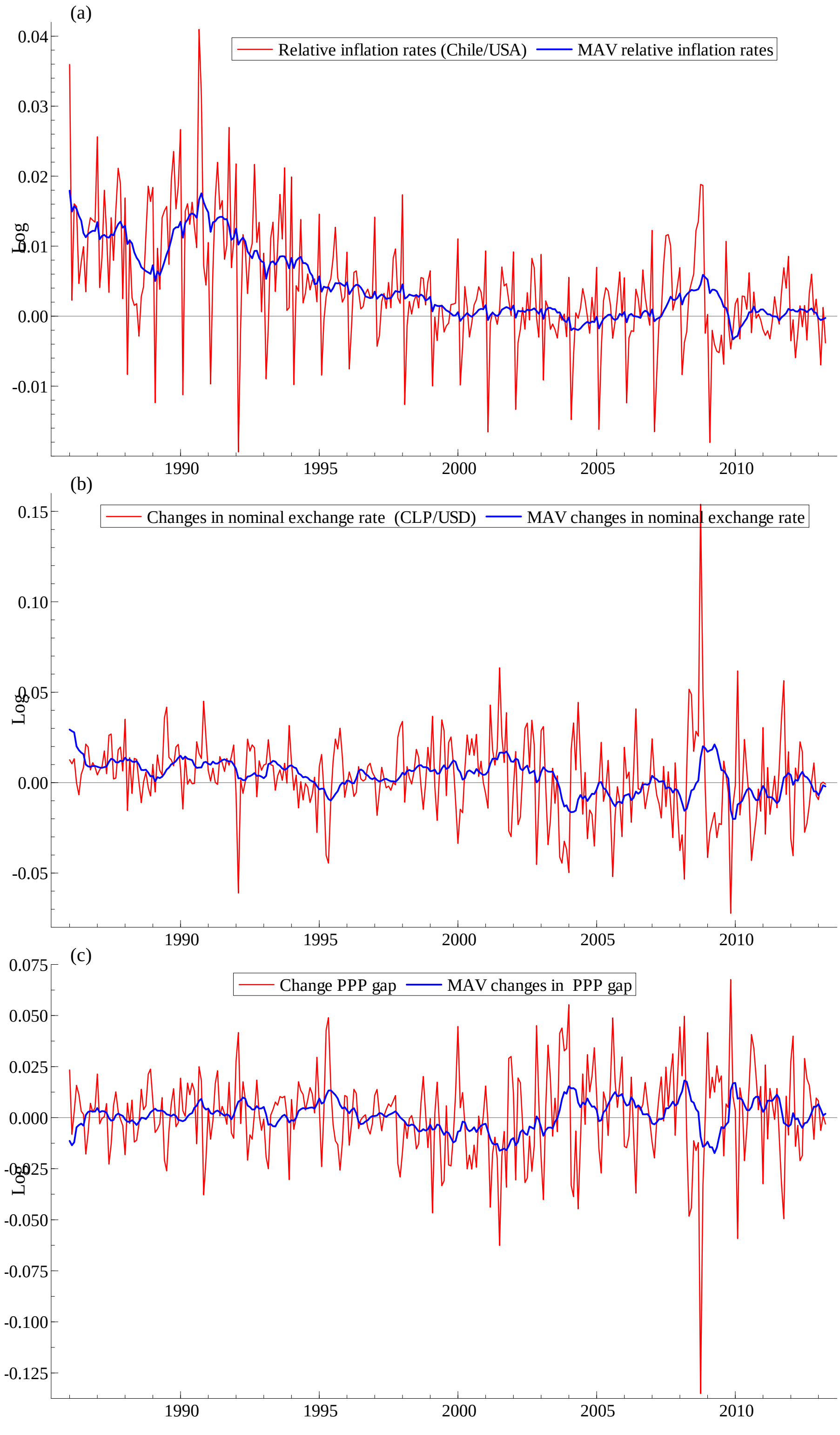


Note 1: CLP: Chilean Peso, US: United States, USD: Unites States Dollar.

Note 2: PPP gap is calculated as the difference between the log of relative prices and the log of the nominal exchange rate.

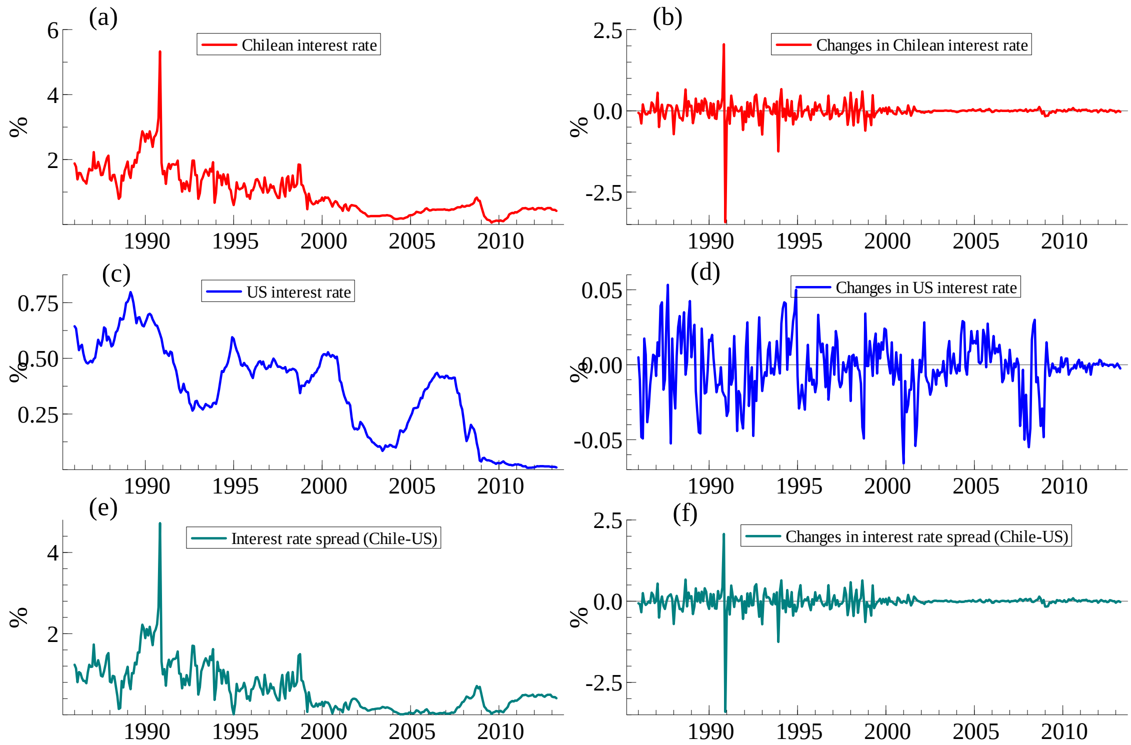
Note 3: Time series in Panel (a) have been shifted to have comparable means.

Figure 1.2: Panel (a): Relative inflation rates (Chile/US). Panel (b): Changes in nominal exchange rate (CLP/USD). Panel (c): Change in the PPP gap. Monthly data 1986:1–2013:4



Note 1: MAV is the 12-month moving average process.

Figure 1.3: Panel (a): Chilean interest rate. Panel (b): Changes in Chilean interest rate. Panel (c): US interest rate. Panel (d): Changes in US interest rate. Panel (e): Interest rate spread (Chile-US). Panel (f): Changes in interest rate spread. Monthly data: 1986:1–2013:4

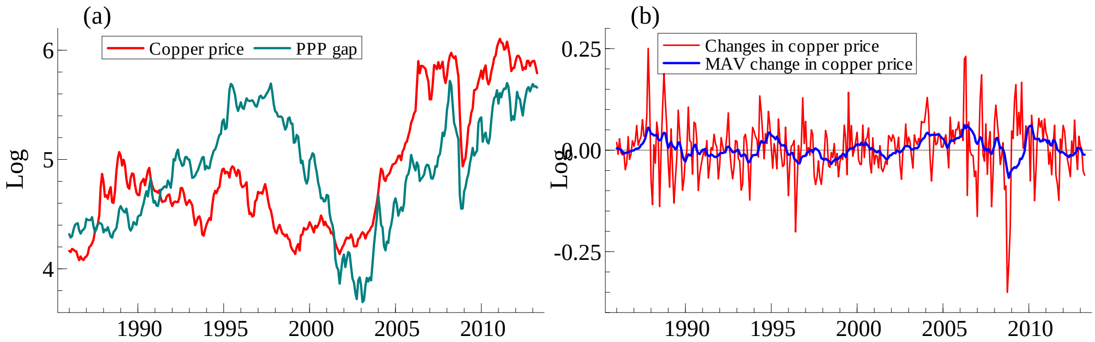


interest rate spread shown in Panel (f) seem to mimic the changes in the Chilean interest rate volatility.

Figure 1.4 Panel (a) plots the copper price and the PPP gap. Two facts are noticeable. First, it seems that both variables are positively co-moving over time, suggesting that there is a negative relationship between copper prices and real exchange rates. Second, since 2005, the copper price has been higher than in the previous years, which might be associated with an increase in world copper demand. The decrease in the copper price observed in 2008 was mainly caused by lower copper demand due to the international financial crisis. Panel (b) shows that the copper price was more volatile at the beginning and end of the sample, and its 12-month moving average suggests some persistence around its mean.

This section discussed the pronounced persistence exhibited in the data. For instance, the graphical analysis seems to suggest that nominal exchange rate, real exchange rate, and relative prices behave as a nonstationary near $I(2)$ process. However, this persistence has to be formally tested, which is done in Section 1.5.

Figure 1.4: Panel (a): Copper price and PPP gap. Panel (b): Changes in the copper price. Monthly data 1986:1–2013:4



Note 1: MAV is the 12-month moving average of the changes in copper price.

Note 2: PPP gap is calculated as the difference between the log of relative prices and the log of the nominal exchange rate.

1.5 Empirical model analysis

1.5.1 Baseline model

The monthly data cover the period 1986:1–2013:4 and the baseline model, which contains three lags,[14] is expressed as:

$$\triangle^2 \boldsymbol{x}_t = \boldsymbol{\alpha} \left[\underbrace{\tilde{\boldsymbol{\rho}}' \begin{pmatrix} \boldsymbol{\tau} \\ \boldsymbol{\tau}_0 \end{pmatrix}'}_{\tilde{\boldsymbol{\tau}}'} \underbrace{\begin{pmatrix} \boldsymbol{x}_{t-1} \\ t-1 \end{pmatrix}}_{\tilde{\boldsymbol{x}}_{t-1}} + \underbrace{\begin{pmatrix} \boldsymbol{d} \\ \boldsymbol{d}_0 \end{pmatrix}'}_{\tilde{\boldsymbol{d}}'} \underbrace{\begin{pmatrix} \triangle \boldsymbol{x}_{t-1} \\ \mathbf{1} \end{pmatrix}}_{\triangle \tilde{\boldsymbol{x}}_{t-1}} \right] + \boldsymbol{\zeta}' \tilde{\boldsymbol{\tau}}' \triangle \tilde{\boldsymbol{x}}_{t-1} + \boldsymbol{\Lambda}_1 \triangle^2 \boldsymbol{x}_{t-1} + \boldsymbol{\Phi}_p \boldsymbol{D}_{p,t} + \boldsymbol{\Phi}_s \boldsymbol{D}_{s,t} + \boldsymbol{\varepsilon}_t \tag{1.31}$$

where $\boldsymbol{x}'_t = [p_{d,t}, p_{f,t}, s_t, cp_t, i_{d,t}, i_{f,t}]$, $p_{d,t}$ is the Chilean CPI, $p_{f,t}$ is the US CPI, s_t is the nominal exchange rate, defined as CLP per USD, cp_t is the copper price, $i_{d,t}$ is the Chilean interest rate, and $i_{f,t}$ is the US interest rate.[15] All variables except interest rates are in natural logarithms. $\tilde{\boldsymbol{\rho}} = [\boldsymbol{\rho}, \mathbf{0}]$, $\mathbf{1}$ is a vector of constant terms and t is a linear trend. $\boldsymbol{D}_{p,t}$ is a (9×1) vector of intervention dummies,[16] and $\boldsymbol{D}_{s,t}$ is a (11×1) vector of centered seasonal dummies.[17] The software CATS in RATS version 2.01, together with a Beta version of CATS for OxMetrics, were used in the econometric analysis.

[14] Appendix B presents the selection of the number of lags.

[15] Appendix A presents the source, description, and transformation of the data.

[16] Appendix C specifies the intervention dummies and their estimated coefficients.

[17] Initially, the cointegration space considered a broken linear trend that started in September 1999, corresponding to the beginning of the floating exchange rate regime in Chile. However, this broken linear trend was revealed to be non-significant. The potential effect of the new regime on the nominal exchange rate was, possibly, offset by changes in the Chilean inflation rate and/or interest rate.

Table 1.1: Misspecification tests for CVAR model (1.31)

Multivariate specification tests				
Autocorrelation		Normality	ARCH	
Order 1: $\chi^2(36)$	Order 2: $\chi^2(36)$	$\chi^2(12)$	Order 1: $\chi^2(441)$	Order 2: $\chi^2(882)$
45.11 [0.14]	41.53 [0.24]	128.94 [0.00]	514.00 [0.00]	1007.42 [0.00]

Univariate specification tests						
Equation	$\triangle^2 p_{d,t}$	$\triangle^2 p_{f,t}$	$\triangle^2 s_t$	$\triangle^2 cp_t$	$\triangle^2 i_{d,t}$	$\triangle^2 i_{f,t}$
ARCH Order 3: $\chi^2(3)$	28.47 [0.00]	13.97 [0.00]	8.58 [0.03]	4.07 [0.25]	22.16 [0.00]	36.72 [0.00]
Normality $\chi^2(2)$	12.83 [0.00]	15.99 [0.00]	12.86 [0.00]	34.31 [0.00]	36.05 [0.00]	6.26 [0.04]
Skewness	0.23	0.07	−0.12	0.03	0.02	−0.17
Kurtosis	3.99	4.12	3.98	4.81	4.87	3.60
S.E.$\times 10^3$	4.38	1.96	16.68	52.17	1.51	0.15

Note 1: [·] is the p-value of the test.

Note 2: S.E. is the residual standard error.

1.5.1.1 Misspecification tests and determining the cointegration rank

Table 1.1 reports the residual misspecification tests of model (1.31).[18] The upper part indicates that the hypotheses of normality and non-ARCH of orders 1 and 2 can be rejected but not the hypothesis of non-autocorrelation. The univariate tests, reported in the lower part, show that all equations exhibit residual non-normality and that only the residuals of the copper price do not show ARCH effects. It appears that the normality problem is due to excess kurtosis rather than excess skewness. Financial variables usually exhibit non-normality and ARCH problems, but adding more dummy variables is not necessarily a solution (Juselius, 2010). Moreover, VAR estimates are robust for moderately excess kurtosis (Gonzalo, 1994; Juselius, 2006).

Table 1.2 reports the $I(2)$ trace test and shows the maximum likelihood test of the joint hypothesis of (r, s_1), which corresponds to the two rank restrictions in (1.27). The standard test procedure starts with the most restricted model, $(r = 0, s_1 = 0, s_2 = 6)$, which is reported in the first row with a likelihood ratio test of 1120.90; it then continues from this point to the right, and row by row, until the first joint hypothesis is not rejected. The first rejection corresponds to the case $(r = 2, s_1 = 2, s_2 = 2)$ with a p-value of 0.22. The case $(r = 1, s_1 = 4, s_2 = 1)$ is also not rejected, though with a lower p-value of 0.14.

Table 1.3 reports the seven largest characteristic roots for $r = 1, 2,$ and 6. The unrestricted model, $(r = 6, s_1 = 0, s_2 = 0)$, has six large roots: five almost on the unit circle and one large but less close to 1 (0.82). The case $(r = 2, s_1 = 2, s_2 = 2)$ implies six characteristic roots to be on the unit circle and leaves 0.21 as the largest unrestricted

[18]For a thorough description of the tests see Dennis (2006).

Table 1.2: Cointegration rank indices model (1.31)

$p-r$	r	$s_2=6$	$s_2=5$	$s_2=4$	$s_2=3$	$s_2=2$	$s_2=1$	$s_2=0$
6	0	1120.90 [0.00]	797.61 [0.00]	582.76 [0.00]	425.17 [0.00]	314.38 [0.00]	232.61 [0.00]	195.29 [0.00]
5	1		579.62 [0.00]	413.11 [0.00]	289.11 [0.00]	179.09 [0.00]	96.24 [0.14]	92.28 [0.02]
4	2			286.97 [0.00]	169.23 [0.00]	84.13 [0.22]	59.00 [0.53]	53.96 [0.25]
3	3				76.83 [0.30]	47.78 [0.76]	31.90 [0.82]	28.59 [0.59]
2	4					26.20 [0.93]	18.59 [0.82]	16.60 [0.45]
1	5						9.31 [0.72]	7.13 [0.34]

Note 1: p is the number of variables in vector $\boldsymbol{x}$.

Note 2: r is the number of cointegrating relationships.

Note 3: s_1 and s_2 are, respectively, the number of $I(1)$ and $I(2)$ trends.

root. The case $(r=1, s_1=4, s_2=1)$ implies the same result and is not rejected, though with a lower p-value.

Table 1.3: Model adequacy

Seven largest characteristic roots							
Unrestricted VAR $(r=6, s_1=0, s_2=0)$	0.98	0.98	0.98	0.95	0.95	0.82	0.21
$(r=1, s_1=5, s_2=0)$	1.00	1.00	1.00	1.00	1.00	0.98	0.21
$(r=1, s_1=4, s_2=1)$	1.00	1.00	1.00	1.00	1.00	1.00	0.21
$(r=2, s_1=4, s_2=0)$	1.00	1.00	1.00	1.00	0.98	0.82	0.21
$(r=2, s_1=2, s_2=2)$	1.00	1.00	1.00	1.00	1.00	1.00	0.21

Note 1: r is the number of cointegrating relationships.

Note 2: s_1 and s_2 are, respectively, the number of $I(1)$ and $I(2)$ trends.

Because two alpha coefficients associated with the second polynomial cointegration relationship are significant for the Chilean price and nominal exchange rate, the selection of model $(r=1, s_1=4, s_2=1)$ might disregard important information in the system.[19] Moreover, the IKE theory-consistent scenario is compatible with $r=2$ (Juselius, 2015).

Table 1.3 also shows that under the assumption that $\boldsymbol{x}_t \sim I(1)$, there would be one large root (0.98) left in the model when $r=1$ and two large roots (0.98 and 0.82) when $r=2$. Under such persistence, treating the process $\boldsymbol{x}_t$ as $I(1)$ is likely to yield unreliable inference (Johansen et al., 2010). Thus, based on the above discussion, the analysis considers the case $(r=2, s_1=2, s_2=2)$, which implies $\boldsymbol{x}_t \sim$ near $I(2)$.

[19]The unrestricted estimates of $\boldsymbol{\alpha}$ for the case $(r=2, s_1=2, s_2=2)$ are shown in the second and third columns in Table 1.4.

1.5.1.2 Weak exogeneity

Paruolo and Rahbek (1999) state the conditions for weak exogeneity with respect to $\boldsymbol{\theta} = \left(\underbrace{\boldsymbol{\beta}, \boldsymbol{\beta}_{\perp 1}}_{\tau}, \boldsymbol{\varsigma}\right)$, where $\boldsymbol{\varsigma} = \overline{\boldsymbol{\alpha}}'\boldsymbol{\Gamma}\overline{\boldsymbol{\beta}_{\perp 2}}$ is a component of $\boldsymbol{d}$ in (1.28).[20] Vector $\boldsymbol{x}_t$ can be decomposed into $\boldsymbol{x}_t' = \left\{\boldsymbol{x}_{1,t}', \boldsymbol{x}_{2,t}'\right\}$, where $\boldsymbol{x}_{1,t} = \boldsymbol{\kappa}_1'\boldsymbol{x}_t$ contains e "endogenous" variables and $\boldsymbol{x}_{2,t} = \boldsymbol{\kappa}_2'\boldsymbol{x}_t$ are the $(p-e)$ candidate weakly exogenous variables. $\boldsymbol{\kappa}_1$ and $\boldsymbol{\kappa}_2$ are of dimensions $p \times e$ and $p \times (p-e)$, respectively. If $r > 0$ and $s_1 > 0$, then the $(p-e)$-dimensional process $\boldsymbol{x}_{2t}$ is weakly exogenous for the cointegration parameters in $\boldsymbol{\theta}$ if and only if the condition $\boldsymbol{\kappa}_2'(\boldsymbol{\alpha}, \boldsymbol{\alpha}_{\perp 1}, \boldsymbol{\zeta}_1) = \mathbf{0}$ is satisfied. Thus, $\boldsymbol{\kappa}_2$ has to be orthogonal to the adjustment coefficients $\boldsymbol{\alpha}$ and $\boldsymbol{\zeta}$ in equation (1.28). In other words, the rows of these matrices that correspond to the weakly exogenous variables must equal zero.

Table 1.4 reports the estimates of $\boldsymbol{\alpha}$, $\boldsymbol{\zeta}$, and their respective t-values, which show that copper price, *cp*, is the only variable with a row of insignificant parameters. That is, copper price seems to be a weakly exogenous variable in system (1.31). Instead of testing for weak exogeneity in system (1.31),[21] the strong exogeneity of copper price can be tested in model (1.26). The hypothesis of 15 parameters being simultaneously equal to zero in the copper price equation must be evaluated (five parameters in $\boldsymbol{\Pi}$, $\boldsymbol{\Gamma}_1$, and $\boldsymbol{\Gamma}_2$). This hypothesis cannot be rejected based on $\chi^2(15) = 13.83$ with p-value of 0.05, implying that copper price is a strong exogenous variable in the system. Thus, copper price is pushing the system but not adjusting to it. Because the copper price is internationally determined, this finding is economically plausible.

Table 1.4: Estimates of $\boldsymbol{\alpha}$ and $\boldsymbol{\zeta}$

Equation	$\boldsymbol{\alpha}$		$\boldsymbol{\zeta} = [\boldsymbol{\zeta}_1, \boldsymbol{\zeta}_2]$			
	$\boldsymbol{\alpha}_1$	$\boldsymbol{\alpha}_2$	$\boldsymbol{\zeta}_{11}$	$\boldsymbol{\zeta}_{21}$	$\boldsymbol{\zeta}_{12}$	$\boldsymbol{\zeta}_{22}$
$\triangle^2 p_{d,t}$	-0.10 (-3.65)	0.42 (10.37)	*	*	*	*
$\triangle^2 p_{f,t}$	*	*	*	*	-4.04 (-6.42)	4.07 (6.36)
$\triangle^2 s_t$	1.07 (9.86)	0.89 (5.80)	*	*	*	*
$\triangle^2 cp_t$	*	*	*	*	*	*
$\triangle^2 i_{d,t}$	*	*	*	-1.30 (-8.96)	2.80 (4.13)	-2.11 (-4.12)
$\triangle^2 i_{f,t}$	-0.004 (-3.99)	*	-0.70 (-10.77)	-0.13 (-9.43)	0.44 (8.79)	-0.45 (-8.85)

Note 1: t-values are given in $(\cdot)$.

Note 2: "*" stands for a coefficient with $|\text{t-value}| < 2.5$.

[20] For further information, see Juselius (2006), Paruolo and Rahbek (1999), and Johansen et al. (2010).

[21] The procedure to test weak exogeneity in $I(2)$ cointegrated systems has not yet been implemented in CATS.

1.5.2 Partial system

Given that the copper price was found to be a strong exogenous variable, a partial system can be modeled with vector $\boldsymbol{x}'_t = \left\{\boldsymbol{x}'_{1,t}, \boldsymbol{x}'_{2,t}\right\}$, where $\boldsymbol{x}'_{1,t} = [p_{d,t}, p_{f,t}, s_t, i_{d,t}, i_{f,t}]$ and $\boldsymbol{x}'_{2,t} = [cp_t]$. Then, equation (1.31) is reformulated as:

$$\begin{aligned} \triangle^2 \boldsymbol{x}_{1,t} =& \boldsymbol{\alpha}\left(\tilde{\boldsymbol{\rho}}'\tilde{\boldsymbol{\tau}}'\tilde{\boldsymbol{x}}_{t-1} + \tilde{\boldsymbol{d}}'\triangle\tilde{\boldsymbol{x}}_{t-1}\right) + \boldsymbol{\zeta}'\tilde{\boldsymbol{\tau}}'\triangle\tilde{\boldsymbol{x}}_{t-1} + \\ &+ \boldsymbol{\Lambda}_1\triangle^2\boldsymbol{x}_{1,t-1} + \sum_{j=0}^{1}\boldsymbol{\pi}_j\triangle\boldsymbol{x}_{2,t-i} + \boldsymbol{\Phi}_p\boldsymbol{D}_{p,t} + \boldsymbol{\Phi}_s\boldsymbol{D}_{s,t} + \boldsymbol{\varepsilon}_t \end{aligned} \tag{1.32}$$

where the left-hand side excludes the acceleration rate of the copper price and the right-hand side adds two lagged differences of the copper price.

To check the adequacy of $r = 2$ in the full model (1.31), Table 1.5 reports the $I(2)$ trace test of the partial model (1.32). The results indicate that the joint hypothesis $(r = 2, s_1 = 2, s_2 = 1)$ cannot be rejected with a p-value of 0.25, which supports the choice of $r = 2$ in the full model (1.31).[22]

In the full model (1.31), the number of $I(2)$ trends was $s_2 = 2$ and in the partial model (1.32) equals one, suggesting that one of the previous two $I(2)$ trends is now accounted for the exogenous copper price. Thus, based on these arguments, the following analysis considers the case $(r = 2, s_1 = 2, s_2 = 1)$.

Table 1.5: Cointegration rank indices model (1.32)

$p-r$	r	$s_2=5$	$s_2=4$	$s_2=3$	$s_2=2$	$s_2=1$	$s_2=0$
5	0	967.98 [0.00]	646.06 [0.00]	432.12 [0.00]	318.98 [0.00]	220.47 [0.14]	180.69 [0.02]
4	1		429.67 [0.00]	291.19 [0.00]	182.14 [0.00]	81.90 [0.01]	77.91 [0.02]
3	2			178.02 [0.00]	74.50 [0.01]	45.15 [0.25]	40.28 [0.29]
2	3				42.00 [0.19]	29.16 [0.19]	19.22 [0.54]
1	4					11.94 [0.47]	7.27 [0.54]

Note 1: p is the number of variables in vector $\boldsymbol{x}_{1,t}$.
Note 2: r is the number of cointegrating relationships.
Note 3: s_1 and s_2 are, respectively, the number of $I(1)$ and $I(2)$ trends.

1.5.2.1 Testing non-identifying hypotheses in the $I(2)$ model

- **Same restrictions on all $\tilde{\boldsymbol{\tau}}$**

The hypothesis of same restrictions on all $\tilde{\boldsymbol{\tau}}$ is formulated as $\mathbf{R}'\tilde{\boldsymbol{\tau}} = \mathbf{0}$, where $\mathbf{R}$ is of dimension $p_1 \times (p_1 - m)$, p_1 is the dimension of $\tilde{\boldsymbol{x}}$ and m is the number of free parameters. The test is asymptotically $\chi^2\left((r + s_1)(p_1 - m)\right)$ distributed (Johansen, 2006).

[22] The p-values of the trace test correspond to a basic model without exogenous variables. Thus, they can only be used as a rough indicator to select a model.

The upper part of Table 1.6 reports three hypothesis restrictions on all $\tilde{\boldsymbol{\tau}}$. The null hypothesis $\mathscr{H}_1$ entails that the nominal to real transformation may be used (Kongsted, 2005). That is, $\boldsymbol{x}_t$ that is near $I(2)$ can be transformed into the $I(1)$ vector $\breve{\boldsymbol{x}}_t' = [ppp_t, \triangle p_{d,t}, \triangle p_{f,t}, i_{d,t}, i_{f,t}, cp_t]$ without loss of information (Johansen et al., 2010). The result of $\mathscr{H}_1$ indicates that the PPP restriction can be rejected; that is, the transformation $(p_{d,t} - p_{f,t} - s_t)$ is not statistically supported.

The null hypothesis, $\mathscr{H}_2$, entails price homogeneity. That is, vector $\boldsymbol{x}_t$ can be transformed into $\breve{\breve{\boldsymbol{x}}}_t' = [p_{d,t} - p_{f,t}, s_t, \triangle p_{d,t}, i_{d,t}, i_{f,t}, cp_t]$ without loss of information. The result of $\mathscr{H}_2$ indicates that price homogeneity can be rejected; that is, the transformation $(p_d - p_f)_t$ is not statistically supported. Finally, the result of hypothesis $\mathscr{H}_3$ indicates that the restricted linear trend is no long-run excludable.

- **A known vector in $\tilde{\boldsymbol{\tau}}$**

In this case, a variable or relationship can be tested to be $I(1)$ in the $I(2)$ model. The restriction is expressed as $\tilde{\boldsymbol{\tau}} = (\mathbf{b}, \mathbf{b}_{\perp}\boldsymbol{\varphi})$ where $\mathbf{b}$ is a $p_1 \times n$ known vector, n is the number of known vectors in $\tilde{\boldsymbol{\tau}}$, and $\boldsymbol{\varphi}$ is a matrix of unknown parameters. The test is asymptotically $\chi^2\left((p_1 - r - s_1)\, n\right)$ distributed unless $\mathbf{b}$ is also a vector in $\tilde{\boldsymbol{\beta}}$ (Johansen, 1996). Thus, $\mathbf{b} \in \mathrm{sp}\left(\tilde{\boldsymbol{\beta}}\right)$ must be checked to ensure the correct distribution of the test. If the hypothesis $\tilde{\boldsymbol{\tau}} = (\mathbf{b}, \mathbf{b}_{\perp}\boldsymbol{\varphi})$ is not rejected and $\mathbf{b} \notin \mathrm{sp}\left(\tilde{\boldsymbol{\beta}}\right)$, then the analyzed variable, or relationship, can be considered $I(1)$.

The lower part of Table 1.6 reports the test results[23] of which hypotheses $\mathscr{H}_4$ to $\mathscr{H}_7$ and $\mathscr{H}_9$ are consistent with the CVAR scenario based on IKE under which nominal exchange rate, prices, relative prices, and nominal interest rate are likely to behave as a near $I(2)$ process. According to IKE, the real exchange rate is likely to behave as a near $I(2)$ process, but the result of $\mathscr{H}_8$ indicates that the hypothesis of the real exchange rate being $I(1)$ cannot be rejected based on a p-value of 0.11. This is, nevertheless, consistent with the high persistence observed in the real exchange rate. In addition, the result of $\mathscr{H}_{10}$ indicates that the copper price is likely to behave as near $I(2)$.

1.5.2.2 Testing over-identifying restrictions on the long-run structure

To identify plausible economic relationships among the variables, a set of restrictions, $\mathscr{H}_{\tilde{\beta}} : \tilde{\boldsymbol{\beta}} = (\mathbf{H}_1\boldsymbol{\vartheta}_1, \ldots, \mathbf{H}_r\boldsymbol{\vartheta}_r)$, must be imposed on $\tilde{\boldsymbol{\beta}} = \tilde{\boldsymbol{\tau}}\tilde{\boldsymbol{\rho}}$, where $\mathbf{H}_i$ is a $p_1 \times m_i$ restriction matrix, $\boldsymbol{\vartheta}_i$ is a $m_i \times 1$ vector of unknown parameters, and m_i is the number of free parameters in $\tilde{\boldsymbol{\beta}}_i$. The test is asymptotically χ^2 distributed with degrees of freedom equal to $\sum_{i=1}^{r}\left((p_1 - m_i) - (r - 1)\right)$ (Johansen et al., 2010).

[23] The hypothesis $\mathbf{b}_i \in \mathrm{sp}\left(\tilde{\boldsymbol{\beta}}\right)$ was rejected in all cases, except for the Chilean interest rate based on $\chi^2(5) = 10.42$ with a p-value of 0.06 and for the interest rate spread based on $\chi^2(5) = 6.80$ with a p-value of 0.23. Thus, the hypotheses $i_{d,t} \sim I(1)$ and $(i_{d,t} - i_{f,t}) \sim I(1)$ are not presented because the distribution of the test is not necessarily χ^2.

Table 1.6: Restrictions on $\tilde{\boldsymbol{\tau}}$

Hypothesis	Matrix restriction design	Distribution	p-value
PPP restriction $\mathscr{H}_1 : \mathbf{R}_1'\tilde{\boldsymbol{\tau}} = \mathbf{0}$	$\mathbf{R}_1' = \begin{bmatrix} 1 & 1 & 0 & 0 & 0 & 0 & 0 \\ 1 & 0 & 1 & 0 & 0 & 0 & 0 \end{bmatrix}$	$\chi^2(8) = 40.70$	0.00
Price homogeneity $\mathscr{H}_2 : \mathbf{R}_2'\tilde{\boldsymbol{\tau}} = \mathbf{0}$	$\mathbf{R}_2' = [1, 1, 0, 0, 0, 0, 0]$	$\chi^2(4) = 38.66$	0.00
Excludable trend $\mathscr{H}_3 : \mathbf{R}_3'\tilde{\boldsymbol{\tau}} = \mathbf{0}$	$\mathbf{R}_3' = [0, 0, 0, 0, 0, 0, 1]$	$\chi^2(4) = 39.14$	0.00
Chilean price $\mathscr{H}_4 : \tilde{\boldsymbol{\tau}} = (\mathbf{b}_1, \mathbf{b}_{1\perp}\boldsymbol{\varphi})$	$\mathbf{b}_1 = [1, 0, 0, 0, 0, 0, 0]$	$\chi^2(3) = 25.15$	0.00
US price $\mathscr{H}_5 : \tilde{\boldsymbol{\tau}} = (\mathbf{b}_2, \mathbf{b}_{2\perp}\boldsymbol{\varphi})$	$\mathbf{b}_2 = [0, 1, 0, 0, 0, 0, 0]$	$\chi^2(3) = 27.19$	0.00
Relative price $\mathscr{H}_6 : \tilde{\boldsymbol{\tau}} = (\mathbf{b}_3, \mathbf{b}_{3\perp}\boldsymbol{\varphi})$	$\mathbf{b}_3 = [1, -1, 0, 0, 0, 0, 0]$	$\chi^2(3) = 24.74$	0.00
Nominal exchange rate $\mathscr{H}_7 : \tilde{\boldsymbol{\tau}} = (\mathbf{b}_4, \mathbf{b}_{4\perp}\boldsymbol{\varphi})$	$\mathbf{b}_4 = [0, 0, 1, 0, 0, 0, 0]$	$\chi^2(3) = 14.15$	0.00
PPP gap $\mathscr{H}_8 : \tilde{\boldsymbol{\tau}} = (\mathbf{b}_5, \mathbf{b}_{5\perp}\boldsymbol{\varphi})$	$\mathbf{b}_5 = [1, -1, -1, 0, 0, 0, 0]$	$\chi^2(3) = 6.01$	0.11
US interest rate $\mathscr{H}_9 : \tilde{\boldsymbol{\tau}} = (\mathbf{b}_6, \mathbf{b}_{6\perp}\boldsymbol{\varphi})$	$\mathbf{b}_6 = [0, 0, 0, 0, 1, 0, 0]$	$\chi^2(3) = 10.07$	0.01
Copper price $\mathscr{H}_{10} : \tilde{\boldsymbol{\tau}} = (\mathbf{b}_7, \mathbf{b}_{7\perp}\boldsymbol{\varphi})$	$\mathbf{b}_7 = [0, 0, 0, 0, 0, 1, 0]$	$\chi^2(3) = 28.25$	0.00

Furthermore, to understand the persistence observed in the variables in the system, it is useful to study the signs and significance of the coefficients in $\boldsymbol{\beta}$, $\boldsymbol{d}$, and $\boldsymbol{\alpha}$. The error correcting- and error-increasing behavior of the variables can be analyzed using the following rules[24]

- If $d_{ij}\beta_{ij} > 0$, changes $\triangle x_{j,t}$ are equilibrium error correcting to the levels, $x_{j,t-1}$. In this case, variable j is equilibrium error correcting in the medium run.

- If $\alpha_{ij}\beta_{ij} < 0$, acceleration rates $\triangle^2 x_{j,t}$ are equilibrium error correcting to the long-run relationship, $\boldsymbol{\beta}_i'\boldsymbol{x}_t$. In this case, variable j is equilibrium error correcting in the long run.

In all other cases, there is equilibrium error increasing behavior.

The selected case, $(r = 2, s_1 = 2, s_2 = 1)$, entails two stationary polynomially cointegrating relationships, $\tilde{\boldsymbol{\beta}}_i'\tilde{\boldsymbol{x}}_t + \tilde{\boldsymbol{d}}_i'\triangle\tilde{\boldsymbol{x}}_t$, where $\tilde{\boldsymbol{\beta}}_i' = \tilde{\boldsymbol{\rho}}_i'\tilde{\boldsymbol{\tau}}_i'$ and $i = 1, 2$. Table 1.7 reports an identified long-run structure on $\tilde{\boldsymbol{\beta}}$, together with unrestricted estimates of $\tilde{\boldsymbol{d}}$ and restricted estimates of $\boldsymbol{\alpha}$,[25] which could not be rejected based on $\chi^2(7) = 5.62$ with a p-value of

[24]For further information, see Chapter 17 in Juselius (2006).

[25]The hypothesis $\mathscr{H}_{\boldsymbol{\alpha}} : \alpha_{51} = \alpha_{52} = 0$ was not rejected based on $\chi^2(2) = 0.03$ with a p-value of 0.98. This indicates that in the long run, the US interest rate has been pushing the system but not adjusting to it.

0.58. To facilitate interpretation, a coefficient in boldface (italics) stands for equilibrium error-correcting (increasing) behavior. Table 1.3 showed that all eigenvalues are inside the unit circle, so that the system is stable and any error-increasing behavior is compensated by error-correcting behavior.

Table 1.7: The estimated long-run $\tilde{\boldsymbol{\beta}}$ structure ($\chi^2(7) = 5.62$ [0.58])

	$p_{d,t}$	$p_{f,t}$	s_t	$i_{d,t}$	$i_{f,t}$	cp_t	$t \times 10^2$	c
$\tilde{\boldsymbol{\beta}}_1'$	-0.01 (-4.6)	0.01 (4.6)	0.01 (4.6)	1.00	-1.00	0.002 (3.3)	-	
$\tilde{\boldsymbol{d}}_1'$	$\mathbf{-0.51}$ (-2.1)	$*$	$\mathit{-0.44}$ (-3.9)	$*$	$*$	$*$		-0.06 (-18.2)
$\boldsymbol{\alpha}_1'$	$\mathit{-0.53}$ (-3.8)	$*$	$*$	$\mathbf{-0.28}$ (-5.7)	-			
$\tilde{\boldsymbol{\beta}}_2'$	-0.03 (-14.3)	0.22 (17.3)	-	1.00	-1.00	-	-0.32 (-16.2)	
$\tilde{\boldsymbol{d}}_2'$	$*$	$*$	-0.89 (-2.2)	$*$	$*$	$*$		-0.95 (-134.1)
$\boldsymbol{\alpha}_2'$	$\mathbf{0.74}$ (9.4)	$\mathit{0.06}$ (1.80)	0.76 (2.5)	$\mathit{0.08}$ (3.0)	-			

Note 1: t-values are given in $(\cdot)$, $*$ stands for a coefficient with $|\text{t-value}| \leq 2.0$.
Note 2: "-" is a zero restriction.
Note 3: a coefficient in boldface (italics) stands for equilibrium error-correcting (increasing) behavior.

The first polynomially cointegrating relationship, $\tilde{\boldsymbol{\beta}}_1'\tilde{\boldsymbol{x}}_t + \tilde{\boldsymbol{d}}_1'\triangle\tilde{\boldsymbol{x}}_t$, is interpreted as the excess returns under IKE and is expressed as:

$$(i_d - i_f)_t = 0.44\triangle s_t + 0.01ppp_t - 0.002cp_t + 0.51\triangle p_{d,t} + 0.06 + \hat{v}_{1,t} \tag{1.33}$$

where $\hat{v}_{1,t} \sim I(0)$ is the equilibrium error. The equation shows that the interest rate spread has been positively co-moving with the PPP gap and the copper price. This relationship resembles the excess returns under IKE, equation (1.16), where the term $(0.51\triangle p_d + 0.44\triangle s)_t$ is likely to be related to the expected change in the nominal exchange rate and the PPP gap is a measure of the uncertainty premium. Moreover, equation (1.33) indicates that the uncovered interest parity is stationary after being adjusted by the PPP gap—the uncertainty premium—and copper price.

Equation (1.33) shows that, exactly as the IKE theory predicts, movements in the interest rate spread co-move with swings in the real exchange rate. That is, the interest rate spread moves in a compensatory manner to restore the equilibrium in the product market when the nominal exchange rate has been away from its benchmark value.

The copper price also enters the relationship that describes the excess returns under IKE, though with a small coefficient. A higher copper price increases the dollar supply in Chile, generating an appreciation of the exchange rate and, consequently, a larger PPP gap. This indicates that the Chilean economy might be affected by the so called commodity super-cycle (Erten and Ocampo, 2013) through the effects that fluctuations in the copper price have on the real exchange rate and, consequently, on competitiveness.

The adjustment coefficients show that the Chilean interest rate is equilibrium error correcting in the long run. The domestic price is equilibrium error increasing in the long run but equilibrium error correcting in the medium run. Moreover, the nominal exchange rate is equilibrium error increasing in the medium run. Thus, if the exchange rate is above its long-run benchmark value, it will tend to appreciate in the medium run, generating a further increase in the equilibrium error term $\hat{v}_{1,t}$. At the same time, the inflation rate starts increasing to restore the equilibrium. In the long run, however, the domestic price will tend to decrease, which again generates further increases in $\hat{v}_{1,t}$. To restore the long-run equilibrium, the domestic interest rate starts decreasing.

The second polynomially cointegrating relationship, $\tilde{\boldsymbol{\beta}}_2'\tilde{\boldsymbol{x}}_t + \tilde{\boldsymbol{d}}_2'\triangle\tilde{\boldsymbol{x}}_t$, can be interpreted as a long-run relationship between the interest rate spread, trend-adjusted prices, and changes in the nominal exchange rate and is expressed as:

$$(i_{d,t} - i_{f,t}) = 0.03\tilde{p}_{d,t} - 0.22\tilde{p}_{f,t} + 0.89\triangle s_t + 0.95 + \hat{v}_{2,t} \tag{1.34}$$

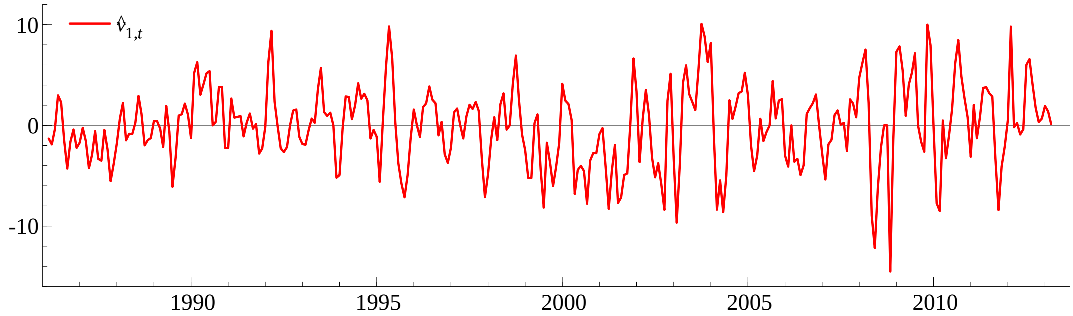
where $\tilde{p}_{f,t}$ and $\tilde{p}_{d,t}$ are, respectively, the trend-adjusted prices in US and Chile and $\hat{v}_{2,t} \sim I(0)$ is the equilibrium error. The equation shows that the interest rate spread is positively co-moving with the relative trend-adjusted level of prices and changes in the nominal exchange rate. This relationship might describe a central bank's reaction rule.

The Chilean trend-adjusted price, $\tilde{p}_{d,t}$, might tentatively be interpreted as a proxy for a long-run indicator of the inflation target. That is, given US interest rate and US trend-adjusted price, if the domestic price is above (below) its long-run trend, the central bank may use contractionary (expansionary) monetary policy to increase (decrease) the domestic interest rate. The above argument may be used to explain the relationship between the interest rate spread and the changes in the nominal exchange rate. For example, the central bank may use contractionary monetary policy to counteract inflationary pressures due to exchange rate depreciation.

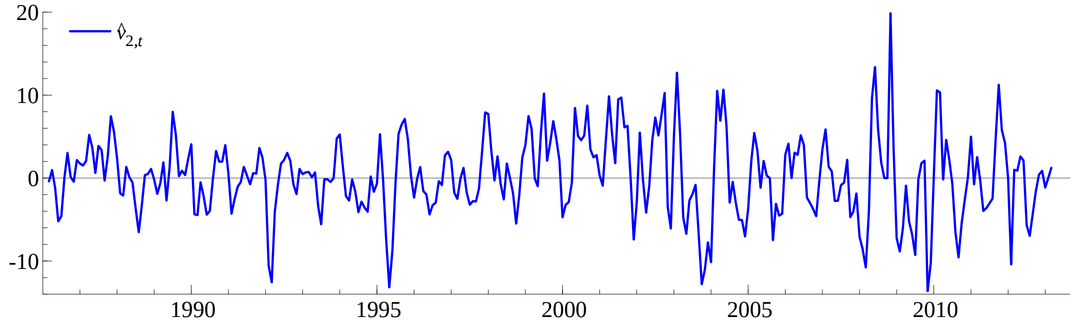
The adjustment coefficients show that when the interest rate spread has been above its long-run value, the exchange rate has appreciated in the medium run and depreciated in the long-run. Furthermore, the domestic price is equilibrium error correcting to the central bank's reaction rule in the long run, whereas the domestic interest rate and the US price are equilibrium error increasing in the long run, the latter with only a borderline significant coefficient, however. Then, if the interest rate spread is above its long-run equilibrium value, the domestic interest rate will tend to increase. This generates further increases in the equilibrium error. However, at the same time, the domestic price will tend to increase, so it starts to restore the equilibrium.

Figure 1.5 shows the graph of the polynomial cointegration relationships and despite some signs of volatility change, they seem mean-reverting.

Figure 1.5: Polynomial cointegrating relationships
$\tilde{\boldsymbol{\beta}}_1'\tilde{\boldsymbol{x}}_t + \tilde{\boldsymbol{d}}_1'\triangle\tilde{\boldsymbol{x}}_t$: Excess returns under IKE



$\tilde{\boldsymbol{\beta}}_2'\tilde{\boldsymbol{x}}_t + \tilde{\boldsymbol{d}}_2'\triangle\tilde{\boldsymbol{x}}_t$: Central bank reaction's rule



Note 1: The graphs are corrected by short-run effects. For further details, see Juselius (2006).

1.5.2.3 The Common Stochastic Trends

Table 1.8 reports the estimated $I(2)$ trend, $\boldsymbol{\alpha}_{\perp 2}$, and its respective estimated loading, $\breve{\boldsymbol{\beta}}_{\perp 2}$. The former may be interpreted as a relative price shock because it loads into prices and exchange rate rather than into exchange rate and interest rates (Juselius, 2015). The estimate of $\boldsymbol{\alpha}_{\perp 2}$ suggests, however, that only shocks to the US price have generated the $I(2)$ trend. The coefficients in $\breve{\boldsymbol{\beta}}_{\perp 2}$ indicate that the $I(2)$ trend loads into nominal exchange rate and relative prices with coefficients of the same sign but different magnitude, which is consistent with the results of hypotheses $\mathscr{H}_4$, $\mathscr{H}_5$, and $\mathscr{H}_7$ in Table 1.6 that prices and exchange rate behave as a near $I(2)$ process. Equations (1.35) and (1.36) show, respectively, the $I(2)$ properties of the relative price and PPP gap.

The relative price is expressed as:

$$(p_{d,t} - p_{f,t}) = (0.26 - 0.04)\,\boldsymbol{\alpha}_{\perp 2}'\sum_{i=1}^{t}\sum_{s=1}^{i}\hat{\boldsymbol{\varepsilon}}_s. \tag{1.35}$$

The loading coefficients to the Chilean CPI and US CPI have the same sign but not the same size. Its difference, 0.22, has to be significant because the result of hypothesis

$\mathscr{H}_6$ in Table 1.6 showed that the relative price is likely to behave as a near $I(2)$ process. The positive loading is consistent with the upward sloping trend in Figure 1.1 Panel (a).

The PPP gap is expressed as:

$$(p_{d,t} - p_{f,t} - s_t) = (0.26 - 0.04 - 0.22)\, \boldsymbol{\alpha}'_{\perp 2} \sum_{i=1}^{t} \sum_{s=1}^{i} \hat{\boldsymbol{\varepsilon}}_s. \tag{1.36}$$

The long-run stochastic trend in relative prices and nominal exchange rate cancels out. This is consistent with both the result of hypothesis $\mathscr{H}_8$ in Table 1.6, which showed that deviations from PPP are likely to behave as an $I(1)$ process, and the long swings in Figure 1.1 Panel (b).

The MA representation suggests that the Chilean economy is primarily affected by external shocks, which is natural when a small and open economy is participating in global markets. Chile has one of the most open economies in the world and also a developed financial market that is almost fully integrated into international markets.

Table 1.8: MA representation

$$\begin{bmatrix} p_d \\ p_f \\ s \\ i_d \\ i_f \end{bmatrix}_t = \underbrace{\begin{bmatrix} 0.26 \\ 0.04 \\ 0.22 \\ -0.00 \\ -0.00 \end{bmatrix}}_{\breve{\boldsymbol{\beta}}_{\perp 2}} \left[\boldsymbol{\alpha}'_{\perp 2} \textstyle\sum_{i=1}^{t} \sum_{s=1}^{i} \hat{\boldsymbol{\varepsilon}}_s\right] + \begin{bmatrix} c_{11} & c_{12} & c_{13} \\ c_{21} & c_{22} & c_{23} \\ c_{31} & c_{32} & c_{33} \\ c_{41} & c_{42} & c_{43} \\ c_{51} & c_{53} & c_{53} \end{bmatrix} \begin{bmatrix} \boldsymbol{\alpha}'_{\perp 2} \sum_{i=1}^{t} \hat{\boldsymbol{\varepsilon}}_i \\ \boldsymbol{\alpha}'_{\perp 1,1} \sum_{i=1}^{t} \hat{\boldsymbol{\varepsilon}}_i \\ \boldsymbol{\alpha}'_{\perp 1,2} \sum_{i=1}^{t} \hat{\boldsymbol{\varepsilon}}_i \end{bmatrix} + \begin{bmatrix} b_{11} \\ b_{21} \\ b_{31} \\ b_{41} \\ b_{51} \end{bmatrix} t$$

$$\boldsymbol{\alpha}'_{\perp 2} = \begin{bmatrix} \underset{(-0.92)}{-0.07} & \mathbf{1.00} & \underset{(-0.39)}{-0.03} & \underset{(0.90)}{0.14} & \underset{(0.32)}{0.35} \end{bmatrix}$$

Note 1: $(\cdot)$ is the t-value.

Note 2: c_{ij} are constant terms.

1.6 Conclusion

The long and persistent swings of the real exchange rate have for a long time puzzled economists. Recent models that build on IKE seem to provide theoretical explanations for this persistence.

This chapter has analyzed the the empirical regularities behind the PPP gap and the uncovered interest rate parity in Chile. The results, based on an $I(2)$ cointegrated vector autoregressive model, gave support for the theoretical exchange rate model based on imperfect knowledge, which assumes that individuals use a multitude of forecasting strategies that are revised over time in ways that cannot be fully prespecified. This is further supported by the results that showed a complex and fairly informative mix of error-increasing and error-correcting behavior.

The results showed that, exactly as the IKE theory predicts, movements in the interest rate spread co-move with swings in the real exchange rate. That is, the interest rate spread moves in a compensatory manner to restore the equilibrium in the product market when the real exchange rate has been away from its long-run value. This relationship describes the excess returns under IKE. The copper price also seems to explain the deviations of the real exchange rate from its long-run equilibrium value. Copper is the main export commodity in Chile and accounts for a large share in total exports; its price fluctuations seems to affect the real exchange rate through its effect on the exchange market.

Altogether, the results indicate that when the interest rate spread is corrected by the uncertainty premium (the PPP gap) and by the fluctuations in the copper price one gets a stationary market-clearing mechanism.

A Data

Table A.1 describes the variables used in this study, their sources, notations, and transformations.

Table A.1: Data Description

Variable	Description	Source	Transformation
$p_{d,t}$	Chilean Consumer Price Index	Central Bank of Chile	Natural logarithm
$p_{f,t}$	US Consumer Price Index	Bureau of Labor Statistics, United States	Natural logarithm
s_t	Nominal exchange rate (Chilean pesos per US dollar)	Central Bank of Chile	Natural logarithm
$i_{d,t}$	1-year Chilean average weighted rates of all transactions in the month by financial commercial banks in Chilean pesos (nominal). Nominal interest rates are annualized (base 360 days) using the conversion of simple interest.	Own elaboration based on data from the Central Bank of Chile	The original variable was divided by 1200 to make it comparable with monthly data
$i_{f,t}$	United States interest rate, Constant Maturity Yields, 1 Year, Average, USD	Own elaboration based on data from the International Monetary Fund	The original variable was divided by 1200 to make it comparable with monthly data
cp_t	Real copper price (USD cents./lb.)	Comisión Chilena del Cobre	Natural logarithm

Note 1: The data set and the CATS file to replicate the results are available upon request.

B Lag-length selection

Table A.2 reports the lag-length selection and lag reduction test. The upper part suggests that $k = 2$ should be selected based on SC and H-Q criteria. However, there is evidence of autocorrelation of order 1 and 2 when $k = 2$. If $k = 3$ is selected, the hypotheses of autocorrelation of orders 1 and 3 can be rejected. The lower part of Table A.2 shows that only the reduction from 4 to 3 lags cannot be rejected.

Table A.2: Lag-length selection model and lag reduction test

Lag-length selection				
Lag: k	SC	H-Q	LM(1)	LM(k)
4	-63.42	-65.35	0.34	0.52
3	-63.97	-65.62	0.13	0.38
2	$\mathbf{-64.28}$	$\mathbf{-65.71}$	0.05	0.04
1	-64.02	-65.20	0.00	0.00

Lag reduction		
Reduction from - to	Test	p-value
VAR(4) - VAR(3)	$\chi^2\,(36) = 41.55$	0.24
VAR (3) - VAR(2)	$\chi^2\,(36) = 95.74$	0.00
VAR(2)-VAR(1)	$\chi^2\,(36) = 291.88$	0.00

Note 1: SC: Schwarz Criterion, H-Q: Hannan-Quinn Criterion.

Note 2: LM(i) stands for a LM-test for autocorrelation of order i.

Note 3: a number in boldface stands for the lowest criteria value.

C Dummy variables

In model (1.31), nine dummies were incorporated. Table A.3 describes the economic facts that justify the dummies, and Table A.4 reports its estimated coefficients.

Table A.3: Dummy justification

Dummy	Variable	Justification
P 1990:9	$+p_d$	The Central Bank of Chile started the partial implementation of an inflation targeting system (BCCh, 2007)
T 1990:11	$+i_d$	INA
P 1993:12	$-i_d$	INA
P 1998:9	$+i_d$	Central Bank of Chile increased the real monetary policy interest rate from 8.5% to 14% (De Gregorio et al., 2005)
P 2005:9	$+p_f$	Energy costs increased sharply. Overall, the index for energy commodities (petroleum-based energy) (BLS, 2005)
P 2006:04	$+cp$	The copper price increased in 30% in April triggered by the lower inventories and higher demand (Cochilco, 2006)
P 2008:10	$-p_f$, $+s$	The energy index fell 8.6% and the transportation index fell in 5.4% in October (BLS, 2008). The nominal exchange rate depreciated 12% due to the dollar strengthening in international markets (BCCh, 2008)
P 2008:11	$-p_f$	The overall CPI index decreased mainly due to a decrease in energy prices, particularly gasoline (BLS, 2008).
P 2010:2	$+s$, $+p_d$	The nominal exchange rate depreciated due to changes in the forward position of the pension funds (BCCh, 2010)

Note 1: P and T stand for a permanent dummy, $(0,\dots,0,1,0,\dots,0)$, and a transitory dummy, $(0,\dots,0,1,-1,0,\dots,0)$, respectively. The signs "$-$" and "$+$" stand for decreases and increases, respectively.
Note 2: INA official information regarding the variable increase or decrease is not available.

Table A.4: Estimated outlier coefficients

Dummy	$\triangle^2 p_d$	$\triangle^2 p_f$	$\triangle^2 s$	$\triangle^2 cp$	$\triangle^2 i_d$	$\triangle^2 i_f$
P 1990:9	0.01 (5.08)	*	*	*	*	*
T 1990:11	*	*	*	*	0.02 (18.44)	*
P 1993:12	*	*	*	*	-0.009 (-5.68)	*
P 1998:9	*	*	*	*	0.005 (3.32)	*
P 2005:9	*	0.01 (4.62)	*	*	*	*
P 2006:4	*	*	*	0.21 (3.98)	*	*
P 2008:10	*	-0.01 (-5.361)	0.14 (8.48)	-0.25 (-4.54)	*	*
P 2008:11	*	-0.01 (-6.73)	*	*	*	*
P 2010:2	0.01 (2.88)	*	0.07 (4.18)	*	*	*

Note 1: $(\cdot)$ is the t-value. * stands for a $|\text{t-value}| \leq 2.0$

Note 2: P and T stand, respectively, for a permanent and a transitory dummy.

Chapter 2

The Phillips Curve and the Role of Monetary Policy in Chile

Leonardo Salazar*

Abstract

In this chapter the empirical analysis finds that the dynamics of inflation and unemployment can be described by a Phillips curve when allowing for a positive co-movement between trend-adjusted productivity and unemployment. This suggests that improvements in productivity have been achieved by laying off the least productive part of the labor force. Furthermore, the natural rate of unemployment is a function of the long-term interest rate, suggesting that monetary policy might not be completely neutral over the business cycle. The results also show that there is a negative co-movement between trend-adjusted productivity and real exchange rate. This indicates that firms adjust profits through improvements in productivity in periods of real exchange rate appreciation.

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2.1 Introduction

Since Phillips (1958) observed a negative relationship between wage inflation and the unemployment rate, known as the Phillips curve, numerous studies have analyzed this relationship empirically and theoretically. Over time, the relationship between the inflation rate and some measure of the economic cycle has been analyzed, giving rise to different formulations of the Phillips curve that have been generally used for the design of economic policies and forecasting inflation.

Karanassou et al. (2010) present a thorough review of the historical development and formulations of the Phillips curve. They noted that models that analyze inflation dynamics are often classified either as standard Phillips curve models or New Keynesian Phillips curve models.

Standard Phillips curve models from the early 1960s find empirical support for a negative relationship between the inflation rate and unemployment as formulated by the traditional Phillips curve:

$$\triangle p_t = c - bu_t + \nu_{1t} \tag{2.1}$$

where $\triangle p_t$ is the inflation rate, u_t is the unemployment rate, $c > 0$ and $b > 0$ are constants, and ν_{1t} is a stationary error term. Phillips (1958) estimated equation (2.1) for the UK as did Samuelson and Solow (1960) for the United States with the expected coefficients signs. This relationship broke down in the 1970s when a positive relationship between inflation and unemployment was observed rather than a negative one; this was called stagflation. It caused Friedman (1968) and Phelps (1968) to develop a new form of the Phillips curve, the expectations-augmented Phillips curve. The idea was that the curve shifts over time as a function of changes in inflationary expectations. Thus, there is a natural rate of unemployment acting as a long-run attractor for the unemployment rate, and in the long run unemployment is independent of the rate of inflation. This formulation is given by:

$$\triangle p_t = \triangle p_t^e - b\left(u_t - u^*\right) + \nu_{2t} \tag{2.2}$$

where $\triangle p_t^e$ is the expected inflation rate, u^* is the natural rate of unemployment that is generally assumed to be constant, and ν_{2t} is a stationary error term. Thus, under the expectations-augmented Phillips curve, there is only a short-run trade-off between inflation and unemployment rate. The New Keynesian Phillips curve models relate actual and expected inflation to some measure of aggregate marginal cost instead of to unemployment.

Most of the the Phillips curve literature on Chile has dealt with the evaluation of the transmission mechanism of monetary policy (Cabrera and Lagos, 2000), the forecast of inflation (De Simone, 2001), and the estimation of NAIRU, or the non-accelerating inflation rate of unemployment (Restrepo, 2008). But some studies also analyze hysteresis as an explanation for the persistence of the unemployment rate (Solimano and Larraín, 2002; Gomes and da Silva, 2008).

These studies find that there is a significant and negative relationship between inflation and cyclical unemployment, as well as between cyclical unemployment and the inflation gap. This is interpreted as evidence for a short-run Phillips curve (Restrepo, 2008). The literature also finds that the Phillips curve seems to provide better forecasting of inflation when the explicit inflation target is used as an independent variable instead of the output gap (De Simone, 2001). Since the GDP gap as well as measures of the business cycle do not show a significant response to an increase in the monetary policy interest rate, Cabrera and Lagos (2000) argue that the Phillips curve is not a proper tool to analyze the transmission mechanism of a monetary policy.

The hysteresis hypothesis has been offered as a way to explain the persistence observed in the rate of unemployment in Chile. Solimano and Larraín (2002) show that this hypothesis cannot be rejected, and Gomes and da Silva (2008) argue that hysteresis is a better explanation of the actual unemployment than the NAIRU. Restrepo (2008) shows that the latter rate is time varying but does not offer an explanation for its cause.

Regardless of the specification of the Phillips curve, most of the available studies are based on a single-equation approach where the rate of inflation is explained by some indicator of the economic cycle (e.g., unemployment, production, marginal cost). Such an approach can be justified by the classical dichotomy that nominal variables do not affect real variables and that inflation and unemployment can be separately analyzed if there is no trade-off between them in the long run. However, the empirical evidence for this dichotomy is quite weak. For instance, Fisher and Seater (1993), King and Watson (1994), Fair (2000), and Karanassou et al. (2005) present evidence of a significant long-run trade-off between inflation and the unemployment rate. Even if the classical dichotomy holds, a single-equation approach is silent about any feedback mechanisms embedded in the data.

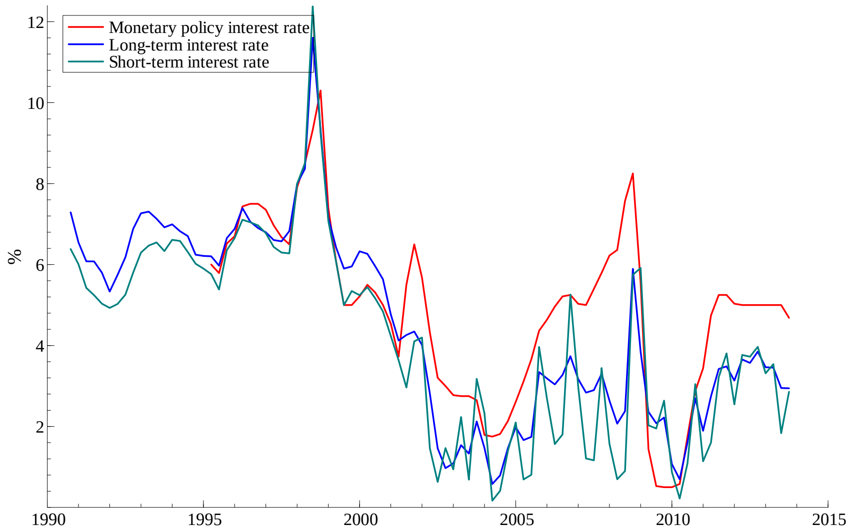
Furthermore, most empirical and theoretical studies assume the natural rate of unemployment to be an exogenous variable. This can also be justified by the classical dichotomy. Generally, in its first stage the Phillips curve is estimated under the assumption of a constant natural rate, and from the residuals of this equation, a time-varying natural rate of unemployment is then derived (Mankiw and Ball, 2002).

The results in this chapter show that the inflation-unemployment dynamic is described by a Phillips curve when allowing for a positive co-movement between trend-adjusted productivity and the unemployment rate. Furthermore, the natural rate of unemployment is a positive function of the long-term nominal interest rate, suggesting that monetary policy might not be completely neutral over the business cycle. This finding is in line with Blanchard (2003), which implicitly states that monetary policy may have real effects on the economy:

> ...if we accept the fact that monetary policy can affect the real interest rate for a decade and perhaps more, then, we must accept, as a matter of logic, that it can affect activity, be it output or unemployment, for a roughly equal time (Blanchard, 2003, p.3)

Figure 2.1 shows the monetary policy interest rate, as determined by the Central Bank of Chile, together with the long- and short-term interest rates. The latter two are calculated as the average long (1 year) and short (3 months) deposit rates in the Chilean financial system. The three interest rates are co-moving over time, suggesting that market interest rates follow the policy rate.

Figure 2.1: Monetary policy interest rate (Central Bank), long-term interest rate (financial system), and short-term interest rate (financial system). Quarterly information 1990:4-2013:4



Note 1: Data on monetary policy rate is only available since 1995:2.

Most Phillips curve studies focused on industrialized economies, particularly in European countries, and only a small proportion on developing countries. The latter is probably due to the scarceness of homogeneous datasets that cover sufficiently long periods, which may be the result of economic and political instability in these economies. Whatever the reason, the lack of research on developing economies is likely to hinder a suitable design of economic policies.

2.2 Theoretical framework

This section discusses the expectations-augmented Phillips curve in a framework developed by Hoover (2011), assuming an economy with imperfect competition and imperfect information about the current price level. In this framework, supply shocks are also allowed to shift the relationship between the unemployment rate and inflation.[1]

2.2.1 Price setting

A firm sets its price based on its expectation of the price level prevailing during the current period, taking demand and supply conditions into account. The price setting is written as:

$$\triangle p_{j,t} = \triangle p^e_{j,t} + f\left(\text{demand factors}\right) + g\left(\text{supply factors}\right) \tag{2.3}$$

where $\triangle$ is the first difference operator, $p_{j,t} = \ln\left(P_{j,t}\right)$ and $P_{j,t}$ is the price set by firm j, $p^e_{j,t} = \ln\left(P^e_{j,t}\right)$ and $P^e_{j,t}$ is the expected level of price prevailing during the current period. This price is set by firm j at the end of period $t-1$ based on an information set available at the end of the same period, $Z_{j,t-1}$. The expected price can be written as $P^e_{j,t} = E_{j,t-1}\left[P_t \middle| Z_{j,t-1}\right]$ where $E\left[\cdot\right]$ is the expectation operator. Functions $f(\cdot)$ and $g(\cdot)$ determine how demand and supply factors affect the pricing decision of firm j.

Equation (2.3) represents a single firm's price behavior. Taking the average of all firms in the economy and assuming that firm-specific supply and demand factors average out, the economy's price behavior is written as:

$$\triangle p_t = \triangle p^e_t + f\left(\begin{array}{c}\text{aggregate demand}\\ \text{factors}\end{array}\right) + g\left(\begin{array}{c}\text{aggregate supply}\\ \text{factors}\end{array}\right) \tag{2.4}$$

where $\triangle p_t$ is the current inflation rate, $\triangle p^e_t$ is the average expectation of general price inflation for all firms, $f(\cdot)$ is reflecting demand-pull inflation, and $g(\cdot)$ captures cost-push inflation.

2.2.2 The Phillips curve and the natural rate of unemployment

Functions $f(\cdot)$ and $g(\cdot)$ must be explicitly defined to apply equation (2.4) to actual data. Since Phillips (1958), an accepted and usual measure of the aggregate demand, has been the unemployment rate:

$$f\left(\text{aggregate demand factors}\right) = a - bu_t \tag{2.5}$$

where u_t is the unemployment rate, b is assumed to be a positive constant given the countercyclical behavior of the unemployment rate, and $a > 0$.

Now, assuming for the moment that aggregate supply factors can be ignored, an unemployment rate that sets actual inflation equal to expected inflation is obtained by

[1] This section presents only the model's main results. For further details, see Chapter 15 in Hoover (2011).

inserting equation (2.5) into equation (2.4):

$$u_t^* = \frac{a}{b} \tag{2.6}$$

where u_t^* is the natural rate of unemployment. If a and b are stable over time, the natural rate can be expressed without subscript t. Equation (2.4) can then be rewritten as:

$$\triangle p_t = \triangle p_t^e - \gamma\left(u_t - u^*\right) + g\,(\text{aggregate supply factors}) \tag{2.7}$$

where $\gamma = b$. Equation (2.7) is the extension of the Phillips curve that allows for supply shocks. When the unemployment rate is below (over) its natural rate, inflation tends to increase faster (slower) than expected.

Equations (2.6) and (2.7) show two classical results. First, the natural rate of unemployment is constant, and second, the Phillips curve is vertical at this level. That is, when expectations are fulfilled and the aggregate supply factors are set at their "natural" levels, there is no long-run trade-off between unemployment and inflation.

2.2.3 Discussion

According to the theoretical framework, the unemployment rate should converge to its natural rate after a supply shock. This assumption is known as the natural rate hypothesis (NRH) and posits that the Phillips curve is vertical in the long run. However, the NRH is not well supported empirically (see Gomes and da Silva, 2008, for Chilean data, and Farmer, 2013, for the US data).

Farmer (2013) proposes that the NRH can be tested in the following way. Under the assumption that expectations are rational, the number of periods (e.g., quarters) where actual inflation is above its expected value should be roughly equal to the number of periods in which the opposite situation is observed. Then, over a decade, the average inflation rate should be approximately equal to the average expected inflation. If the inflation rate over long periods is plotted together with the unemployment rate, a vertical line at the natural rate of unemployment should be visible, supporting the NRH and rational expectations.

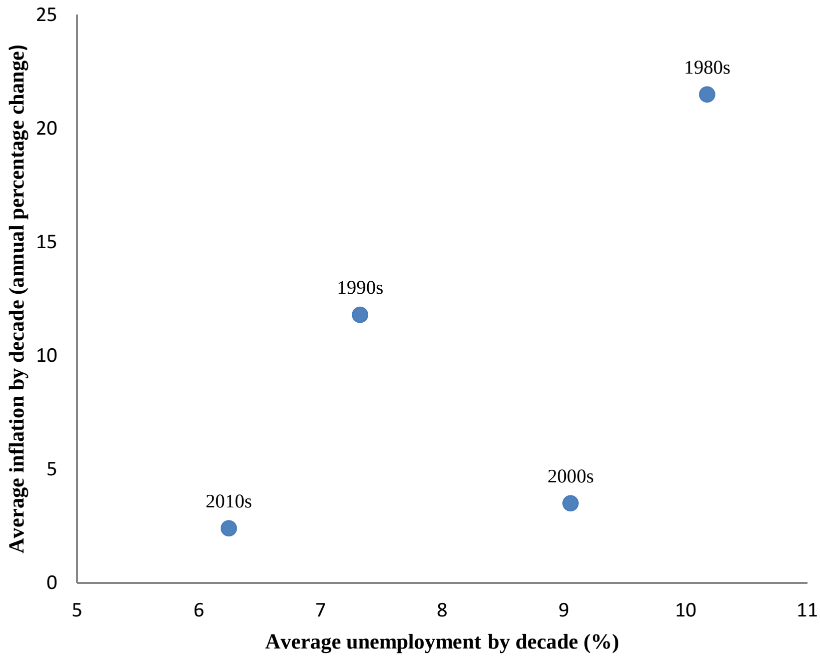
Figure 2.2 shows Chile's average inflation and unemployment rates by decade.[2] The plotted points are not vertically aligned, and there is no tendency for them to lie around a vertical line. Farmer (2013) obtains a similar result for US data and categorically concludes that because expectations are unlikely to be systematically biased over decades, the NRH is false. However, such a strong conclusion should not rely on a simple graph, and further tests must be provided.

The result that unemployment does not converge to a unique constant value in the

[2] 1980s includes the period 1986–1989 and 2010s includes the period 2010–2013.

long run might potentially explain this variable's persistence. It also suggests that the natural rate of unemployment is time varying. Phelps (1994) argues, in his structural slumps theory, that the long swings observed in European unemployment rates can be explained by fluctuations in exchange rates and real interest rates. Specifically, domestic real interest rates likely influence the natural rate of unemployment. In this case, the natural rate of unemployment is time varying, and its fluctuations are associated with the movements observed in real interest rates.

Figure 2.2: Average inflation and unemployment by decade in Chile



Note 1: 1980s includes 1986–1989, and 2010s includes 2010–2013.

Phelps (1994) provides two reasons for explaining the positive co-movement between the natural rate of unemployment and the interest rate. First, higher real interest rates increase the natural rate of unemployment by discouraging investment. Examples would be investment in the retention of workers (high interest rate reduces the probability of paying higher wages) or investment that could increase the productivity of the firm's workforce. Second, equilibrium employment will decrease with higher interest rates when government actions reduce firms' labor demand. For example, actions that affect the wealth of the working-age population will raise the real wage that workers demand (Aghion et al., 2003)

Phelps (1994) assumes a world where the unemployment rate and the real interest rate are stationary. However, empirical studies often find real interest rates to be indistinguishable from a unit root process. Such persistence in real interest rates is consistent with the theory of imperfect knowledge economics (IKE), developed by Frydman and Goldberg (2007, 2011), and Juselius and Juselius (2012) argue, based on Finish data, that the structural slumps theory based on IKE is better able to explain the persistent swings observed in the data.

Juselius and Juselius (2012) give the following rationale. Under IKE, nominal interest rates exhibit strong persistence due to a nonstationary uncertainty premium, whereas inflation rates are more stable over time due to international competitiveness. This implies that the Fisher effect does not hold as a stationary condition. The uncertainty premium is generally related to the concept of a "gap effect." In the foreign currency market the gap effect can be measured by the deviation of the real exchange rate from its long-run purchasing power parity value. In an IKE world, due to speculative behavior in the currency market, nominal exchange rates tend to move away from relative prices for long periods of time, causing the real exchange rate to behave like a near $I(2)$ process. Persistent deviations of the real exchange rate from its long-run benchmark value will be reflected in the uncertainty premium and hence in domestic and foreign interest rates.

Then an increase (decrease) in nominal interest rates, in general, is not likely to be followed by an increase (decrease) in consumer price inflation, the real interest rate will tend to rise (drop). This is likely to result in inflows (outflows) of speculative capital, causing an appreciation (depreciation) of the real exchange rate and worsening (improving) domestic competitiveness. Under this situation, domestic firms in the tradable sector cannot count on exchange rates to restore competitiveness after a shock to relative costs (e.g., a large wage rise). In this case, domestic firms will be prone to adjust profits rather than prices. Profits can be adjusted through improvements in labor productivity by laying off the least productive workers. Thus, an increase in both labor productivity and unemployment might be expected in periods of real appreciation and rising real interest rates.

Then a more adequate and general representation of the Phillips curve is given by:

$$\triangle p_t = \triangle p_t^e - \gamma \left(u_t - u_t^*\right) + \nu_{3t} \tag{2.8}$$

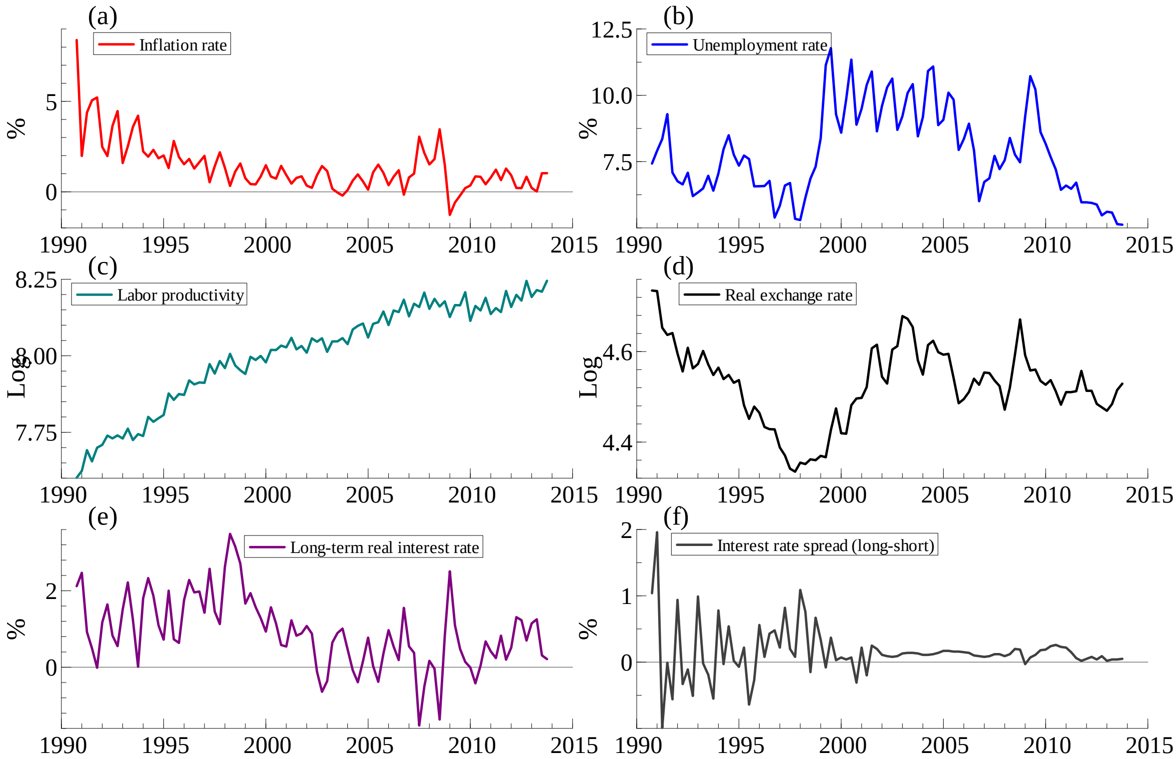
where ν_{3t} is a stationary error term and the time-varying natural rate of unemployment, u_t^*, can be expressed as $u_t^* = z\left(i_t\right)$, where i_t is the real interest rate and $z' > 0$.

2.3 Stylized facts

Figure 2.3 Panel (a) shows the quarterly inflation rate, which exhibits a gradual decrease over the sample. This might be associated with the implementation of inflation targeting in 1990. By mandate, the Central Bank of Chile is compelled to keep the annual inflation rate centered on 3% with fluctuations in the range of 2%–4%. This has been more or less the case since 2000.

Figure 2.3 Panel (b) shows the evolution of the unemployment rate, which exhibits an important increase in 1998, possibly as a result of the Asian Crisis.[3] The unemployment rate exhibits a higher mean after that crisis, suggesting that the mean of the natural rate of unemployment might have increased as well. Indeed, the rate of unemployment rose from an average rate of 6.9% to 8.4% after the Asian crisis. The unemployment rate exhibits another significant increase in 2009 when the global financial crisis hit the Chilean economy. This increase, however, seems transitory.[4]

Figure 2.3: Panel (a): Inflation rate. Panel (b): Unemployment rate. Panel (c): Real labor productivity. Panel (d): Real exchange rate. Panel (e): Long-term real interest rate. Panel (f): Interest rate spread (long and short-term interest rate). Quarterly information 1990:4–2013:04



Unemployment rate is not the only variable seemingly affected by the Asian crisis. Figure 2.3 Panel (c) shows a deceleration in the growth of real productivity around 1998, which again might indicate a slowdown in economic activity as a result of the Asian crisis. The unemployment rate and productivity exhibit seasonality caused by the agricultural activity in Chile, which is higher during the last and first quarter of each year.

Figure 2.3 Panel (d) shows the real exchange rate, and since 2000, it exhibits a higher mean. This may be related to the the free floating exchange rate regime implemented by

[3]The Asian crisis hit the Chilean economy in 1998. The tradable sector was the most affected because about 48% of the total exports were sent to Asia in 1998. The decrease in Asian demand triggered the bankruptcy of many companies, leading to a large increase in the unemployment rate.

[4]After the Asian crisis, structural reforms were introduced in the labor market to reduce the impact of domestic and international shocks.

the Central Bank of Chile in September 1999. The graph suggests that the real exchange rate may be fluctuating around a constant level, though with persistent deviations.

Two major reforms were introduced in the Chilean financial market between 2000 and 2001. While the first reform, promulgated in 2000, gave greater protection to both domestic and foreign investors, the second reform, enacted in 2001, liberalized the financial system, implying, among other things, capital account deregulation. These reforms seemed to affect the real interest rate which had exhibited a lower mean since 2001 (Figure 2.3 Panel (e)). At the same time, Figure 2.3 Panel (f), shows a large volatility reduction in the the interest rate spread between the long- and short-term rates.

2.4 The empirical model analysis

2.4.1 Baseline model

The quarterly data cover the period 1990:4–2013:4, and the information set[5] is given by $\boldsymbol{x}'_t = [\triangle p_t, u_t, ri_t, sp_t, c_t, q_t]$ where

- $\triangle p_t$ is the inflation rate.
- u_t is the unemployment rate.
- $ri_t = i^L_t - \triangle p_t$ is the long-term real interest rate, and i^L_t is the long-term nominal interest rate.
- $sp_t = i^L_t - i^S_t$ is the interest rate spread, and i^S_t is the short-term nominal interest rate. Following Juselius and Juselius (2012), this spread is used as a proxy for expected inflation.
- c_t is the real labor productivity.
- q_t is the real exchange rate.

All variables are in logs, except unemployment rate, real interest rate, and spread.

The baseline cointegrated VAR (CVAR) model for vector $\boldsymbol{x}_t$ is:

$$\triangle \boldsymbol{x}_t = \boldsymbol{\alpha}\tilde{\boldsymbol{\beta}}'\tilde{\boldsymbol{x}}_{t-1} + \boldsymbol{\Gamma}_1 \triangle \boldsymbol{x}_{t-1} + \sum_{i=0}^{1} \boldsymbol{\delta}_i d_{s98:3,t-i} + \boldsymbol{\mu}_0 + \boldsymbol{\delta}_2 \boldsymbol{D}_{p,t} + \boldsymbol{\delta}_3 \boldsymbol{S}_t + \boldsymbol{\varepsilon}_t \tag{2.9}$$

where $\tilde{\boldsymbol{x}}_t = [\boldsymbol{x}_t, d_{s00:4,t}, t_{98:3,t}, t]'$, $d_{s00:4,t}$ is a step dummy that takes the value of 1 since 2000:4, 0 otherwise; $t_{98:3,t}$ is a broken linear trend taking values of $1, 2, \ldots, 62$ since 1998:3, 0 otherwise; and t is a linear trend. $d_{s98:03,t}$ is a shift dummy that takes the value of 1 since

[5]Appendix A presents the source and transformation of the data.

1998:3, 0 otherwise, corresponding to the first difference of the broken linear trend. $\boldsymbol{\mu}_0$ is a vector of constant terms, $\boldsymbol{D}_{p,t}$ contains 14 permanent dummies $(0,\ldots,0,1,0,\ldots,0)$,[6] and $\boldsymbol{S}_t$ is a vector of centered seasonal dummies. Finally, $\boldsymbol{\varepsilon}_t$ is an i.i.d. vector of normally distributed stochastic errors.

The step dummy $d_{s00:4,t}$ accounts for deregulation in the Chilean financial markets, and the broken linear trend accounts for the productivity slowdown in 1998. Both deterministic variables, together with the linear trend, are restricted to lie in the cointegration space.

2.4.2 Misspecification tests and determining the cointegration rank

Table 2.1 reports the residual misspecification tests of model (2.9).[7] The upper part shows that the model is, in general, well behaved. The hypotheses of non-autocorrelation and non-ARCH cannot be rejected, and there are only weak signs of non-normality. The univariate tests in the lower part of Table 2.1 show that only residual ARCH and signs of non-normality are present in the interest rate spread. The ARCH problem is evident from Figure 2.3 Panel (f) and the normality problem is associated with excess of kurtosis rather than skewness. Except for these minor misspecification signs, the model seems reasonably well specified.

Table 2.1: Misspecification tests CVAR model

Multivariate tests				
Autocorrelation		Normality	ARCH	
Order 1 : $\chi^2(36)$	Order 2 : $\chi^2(36)$	$\chi^2(12)$	Order 1 : $\chi^2(441)$	Order 2 : $\chi^2(882)$
34.78 [0.53]	45.44 [0.14]	21.39 [0.05]	382.66 [0.98]	925.35 [0.15]

Univariate tests	$\triangle^2 p_t$	$\triangle u_t$	$\triangle ri_t$	$\triangle sp_t$	$\triangle c_t$	$\triangle q_t$
ARCH Order 2:$\chi^2(2)$	0.29 [0.86]	3.20 [0.20]	0.85 [0.65]	13.83 [0.00]	3.27 [0.19]	4.77 [0.10]
Normality $\chi^2(2)$	1.66 [0.44]	1.99 [0.37]	0.61 [0.74]	11.01 [0.00]	0.81 [0.67]	0.17 [0.92]
Skewness	0.05	0.24	-0.16	0.02	0.16	0.10
Kurtosis	3.33	3.34	2.99	4.47	3.06	2.77

Note 1: [·] is the p-value of the test.

The upper part of Table 2.2 reports the eigenvalues, λ_i, for the null hypothesis of $r = 0,1,2,3,4,5$ and the corresponding Bartlett-corrected $I(1)$ trace test with the p-value in brackets. The hypothesis $r = 5$ cannot be rejected based on a p-value of 0.34.

[6] Each permanent dummy takes the value of 1 in 1992:1, 1993:1, 1993:4, 1994:1, 1995:3, 1995:4, 1997:4, 2000:4, 2003:1, 2006:4, 2007:3, 2008:3, 2009:1, and 2010:1; 0 otherwise.

[7] Dennis (2006) provides a thorough description of the misspecification tests.

To check the adequacy of this choice, the lower part of Table 2.2 reports the four largest characteristics roots for $r = 6, 5, 4, 3$. The unrestricted model, $r = 6$, has only one reasonably large root, 0.75, suggesting that the model contains one unit root at most. $r = 5$ leaves a complex pair of roots with a modulus of 0.65 as the largest root in the system. Thus, the following analysis is based on $r = 5$ cointegrating vectors.

Table 2.2: Cointegrating rank and model adequacy

I(1) trace test					
$p-r$	$H_0: r =$	Eigenvalues (λ_i)	Trace test	p-value	$Q_{.95}$
6	0	0.76	333.74	[0.00]	137.90
5	1	0.57	215.05	[0.00]	107.06
4	2	0.48	142.22	[0.00]	79.91
3	3	0.43	88.14	[0.00]	54.99
2	4	0.26	38.47	[0.02]	34.88
1	5	0.11	11.12	[0.34]	17.64
Modulus of the five largest characteristic roots					
$r = 6$	0.75	0.65	0.65	0.63	
$r = 5$	1.00	0.65	0.65	0.63	
$r = 4$	1.00	1.00	0.65	0.65	
$r = 3$	1.00	1.00	1.00	0.60	

Note 1: p is the number of variables in vector $\boldsymbol{x}_t$

Note 2: $[\cdot]$ is the p-value of the trace test simulated according to model (2.9)

Note 3: $Q_{.95}$ is the 5% critical value of the trace test

2.4.3 Identification of the long-run structure

To identify plausible economic relationships among the variables, a set of restrictions, $\mathscr{H}_{\tilde{\boldsymbol{\beta}}}$: $\tilde{\boldsymbol{\beta}} = (\boldsymbol{H}_1\boldsymbol{\varphi}_1, \boldsymbol{H}_2\boldsymbol{\varphi}_2, \ldots, \boldsymbol{H}_r\boldsymbol{\varphi}_r)$, must be imposed on $\tilde{\boldsymbol{\beta}}$, where $\boldsymbol{H}_i$ is a restriction matrix of dimension $p_1 \times (p_1 - m_i)$, p_1 is the dimension of $\tilde{\boldsymbol{x}}$, m_i is the number of restrictions imposed on $\tilde{\boldsymbol{\beta}}_i$, $p_1 - m_i$ is the number of freely varying parameters, and $\boldsymbol{\varphi}_i$ is a $(p_1 - m_i) \times 1$ vector of unknown parameters. This test is asymptotically χ^2 distributed with degrees of freedom equal to $\sum_{i=1}^{r} (m_i - (r-1))$ (Johansen, 1996). When $\alpha_{ij}\beta_{ij} < 0$ (> 0), the cointegrating relation is equilibrium correcting (increasing) in the equation $\triangle \boldsymbol{x}_{j,t}$ (Juselius, 2006).

Table 2.3 reports an identified structure on $\tilde{\boldsymbol{\beta}}$ with nine overidentified restrictions that could not be rejected based on $\chi^2(9) = 7.60$ with a p-value of 0.58.[8] To facilitate interpretation, an α_{ij} coefficient in boldface means that the cointegrating relation i is equilibrium correcting, whereas an error increasing coefficient is in italics. The results in Table 2.2 showed that all eigenvalues are inside of the unit circle. Thus, the system is stable, and any error-increasing behavior is compensated by error-correcting behavior.

[8]Appendix B indicates that the eigenvalue fluctuation test does not show signs of parameter-no constancy.

Table 2.3: The estimated long-run $\tilde{\boldsymbol{\beta}}$ structure ($\chi^2\,(9) = 7.60\ [0.58]$)

	$\triangle p_t$	u_t	ri_t	sp_t	c_t	q_t	$d_{s00:4,t}$	$t_{98:3,t}$	t
$\tilde{\boldsymbol{\beta}}_1'$	1.00	0.48 (8.54)	-	-	−0.22 (−8.09)	-	-	−0.003 (−11.83)	0.004 (12.20)
$\boldsymbol{\alpha}_1'$	**−0.51** (−6.76)	*0.22* (2.47)	0.25 (3.62)	∗	**1.07** (3.85)	−2.11 (−4.79)			
$\tilde{\boldsymbol{\beta}}_2'$	-	1.00	−1.00	-	-	-	−0.01 (−2.97)	-	0.001 (12.35)
$\boldsymbol{\alpha}_2'$	0.30 (4.95)	**−0.44** (−6.08)	*−0.25* (−4.55)	∗	∗	1.05 (2.96)			
$\tilde{\boldsymbol{\beta}}_3'$	-	-	1.00	−1.00	-	-	0.01 (4.08)	-	-
$\boldsymbol{\alpha}_3'$	0.39 (3.36)	−0.29 (−2.13)	**−0.80** (−7.55)	*−0.11* (−2.55)	∗	∗			
$\tilde{\boldsymbol{\beta}}_4'$	0.07 (3.10)	-	-	1.00	-	-	-	0.0003 (8.34)	−0.0003 (−8.34)
$\boldsymbol{\alpha}_4'$	*1.00* (4.31)	∗	−1.44 (−6.72)	**−1.18** (−13.39)	∗	∗			
$\tilde{\boldsymbol{\beta}}_5'$	-	-	-	-	1.00	1.00	−0.16 (−7.64)	-	−0.002 (−5.11)
$\boldsymbol{\alpha}_5'$	∗	∗	∗	∗	∗	**−0.58** (−7.60)			

Note 1: (·) is the t-value. ∗ stands for an alpha coefficient with |t-value| $\leq$ 2.0.

Note 2: "-" is a zero restriction.

Note 3: a coefficient in boldface stands for an equilibrium error correcting behavior.

Note 4: a coefficient in italics stands for an equilibrium error increasing behavior.

The first cointegrating vector, $\tilde{\boldsymbol{\beta}}_1'\tilde{\boldsymbol{x}}_t$, is interpreted as a Phillips curve over the business cycle and is expressed as:

$$\triangle p_t = -0.48\,(u_t - 0.45\tilde{c}_t) + \hat{v}_{1,t} \tag{2.10}$$

where $\tilde{c}_t = c_t + 0.013t_{98:3,t} - 0.018t$ stands for trend-adjusted productivity and $\hat{v}_{1,t} \sim I(0)$ measures the equilibrium error. Equation (2.10) shows that unemployment in excess of trend-adjusted productivity would lead to downward pressure on the inflation rate. This finding is consistent with the reasoning above that in a world of imperfect knowledge, firms exposed to international competition facing a positive shock to relative costs cannot, in general, count on exchange rates to restore competitiveness. Accordingly, firms will be prone to adjust profits rather than prices (Juselius and Juselius, 2012). Profits can be adjusted through improvements in productivity by laying off the least productive workers. Evidence of a positive co-movement between unemployment and productivity has also been found in Juselius (2006) for Danish data.

The adjustment coefficients show that, while the inflation rate and productivity are equilibrium error correcting to the Phillips curve, the unemployment rate is equilibrium error increasing. Thus if unemployment is above its long-run equilibrium value, the trend-adjusted productivity will tend to increase and, at the same time, the unemployment rate will tend to increase as long as enterprises improve labor productivity over its long-run trend. This introduces further increases in the equilibrium error. The inflation rate will start decreasing, which tends to restore equilibrium. Moreover, the adjustment coefficients show that long-term interest rate has been positively affected by the equilibrium error in the Phillips curve and that the real exchange rate has appreciated when the unemployment rate has been above its long-run equilibrium value.

The second vector, $\tilde{\boldsymbol{\beta}}_2'\tilde{\boldsymbol{x}}_t$, is interpreted as an unemployment relationship and is expressed as:

$$u_t = ri_t + 0.01d_{s00:4,t} - 0.001t + \hat{v}_{2,t} \tag{2.11}$$

where $\hat{v}_{2,t} \sim I(0)$ is the equilibrium error. The linear trend in (2.11) can be understood from Figure 2.3 Panel (e). The graph shows that the real interest rate has shown a tendency to decrease over time,[9] which is approximated by the deterministic trend in equation (2.11).

The adjustment coefficients show that, while the unemployment rate is equilibrium error correcting to equation (2.11), the real interest rate is equilibrium error increasing. Then, even though the interest rate is above its long-run equilibrium value, it will continue to rise. At the same time, the unemployment rate will also increase, restoring equilibrium. This seems to be greatly aggravated by the fact that the inflation rate has been positively affected by the equilibrium error $\hat{v}_{2,t}$. That is, the inflation rate shows a tendency to decrease when the real interest rate is increasing, generating further increases in the unemployment rate. In addition, the adjustment coefficients show that the real exchange rate has appreciated when the real interest rate has been above its long-run trend.

The natural rate of unemployment can be derived from equation (2.11),[10] and is expressed as:

$$u_t^* = i_t^L \tag{2.12}$$

where the natural rate of unemployment, u_t^*, is time varying and positively associated with the long-term nominal interest rate. This result indicates that monetary policy in Chile may not be completely neutral over the business cycle. That is, the Central Bank of Chile conducts its monetary policy using the interest rate as the main tool to keep the inflation rate close to its target, and equation (2.12) suggests that fluctuations in the nominal interest rate are positively associated with the natural rate of unemployment. Equation (2.12) shows that the natural rate is a function of the nominal rather than the

[9] Appendix C shows that this tendency is more evident in the nominal long-term interest rate.

[10] Equation (2.11) can be equivalently rewritten as $u_t = \left(i_t^L - \triangle p_t\right) + 0.01d_{s00:4,t} - 0.001t + \hat{v}_{2,t}$, implying that $\triangle p_t = -u_t + i_t^L + 0.01d_{s00:4,t} - 0.001t + \hat{v}_{2,t} = -\left(u_t - u_t^*\right) + 0.01d_{s00:4,t} - 0.001t + \hat{v}_{2,t}$, where $u_t^* = i_t^L$.

real interest rate, which can be explained by the fact that the rate of inflation has been fairly stable over time, as shown in Figure 2.3 Panel (a).

In an IKE world, this finding provides empirical support for Phelps's hypothesis of a Phillips curve where the natural rate of unemployment is positively co-moving with the interest rate. A time-varying natural rate of unemployment is consistent with the non-vertical scatter of the inflation and unemployment rates shown in Figure 2.2. Evidence of a natural rate of unemployment as function of the interest rate has also be found in Juselius and Ordóñez (2009) for Spanish data and in Juselius (2006) for Danish data.

Equation (2.11) also shows that unemployment is nonstationary per se and needs to be combined with another variable to obtain stationarity. This result provides another explanation of the persistence of the unemployment rate than the hysteresis hypothesis of previous studies in Chile.

The third vector, $\tilde{\boldsymbol{\beta}}_3'\tilde{\boldsymbol{x}}_t$, is interpreted as a short-term real interest rate relationship and is expressed as:

$$ri_t = sp_t - 0.01d_{s00:4,t} + \hat{v}_{3,t}$$

or, equivalently:

$$\left(i^S - \triangle p\right)_t = -0.01d_{s00:4,t} + \hat{v}_{3,t} \tag{2.13}$$

where $\hat{v}_{3,t} \sim I(0)$ measures the equilibrium error. It describes a stationary short-term real interest rate after controlling for an equilibrium mean shift in 2000:4. The latter is likely to reflect the effect of reforms introduced in the Chilean financial system. The adjustment coefficients show that, while the long-term real interest rate is equilibrium error correcting to equation (2.13), the spread is equilibrium error increasing. When the real interest rate is above its long-run equilibrium value, the spread will tend to decrease which increases the equilibrium error. At the same time, the real interest rate will tend to decrease, restoring the equilibrium. The inflation rate has been positively affected by the equilibrium error, which also helps to reduce the real interest rate.

The fourth relation, $\tilde{\boldsymbol{\beta}}_4'\tilde{\boldsymbol{x}}_t$, is interpreted as a central bank's reaction rule and is expressed as:

$$sp_t = -0.07\triangle p_t + 0.0003t_{(90:4-98:3),t} + \hat{v}_{4,t} \tag{2.14}$$

where $\hat{v}_{4,t} \sim I(0)$ measures the equilibrium error. It indicates that the central bank tends to increase the short-term interest rate to counteract inflationary pressures, thereby reducing the spread. This is consistent with the countercyclical policy of the Central Bank of Chile. The latter tends to increase its interest rate when there are signs of inflationary pressure. The term $t_{(90:4-98:3),t}$,[11] that stands for a trend between 1990:4 and 1998:3, is needed because before 1998:3 the inflation rate behaved as a trending variable, which is shown in Figure 2.3 Panel (a), but since 1998:3 the inflation rate is fluctuating around a roughly constant average, which is consistent with the full implementation of inflation targeting in 1999.

The adjustment coefficients show that while the the spread is equilibrium error correcting to the central bank's reaction rule, the inflation rate is equilibrium error increasing. When the spread is above its long-run equilibrium value, the inflation rate will tend to increase, generating further increases in the equilibrium error. At the same time, the spread will tend to narrow, restoring the equilibrium. The long-term interest rate has been negatively affected by the equilibrium error, which helps to narrow the interest rate spread.

Finally, the fifth relation, $\tilde{\boldsymbol{\beta}}_5'\tilde{\boldsymbol{x}}_t$, is interpreted as a productivity relationship and is expressed as:

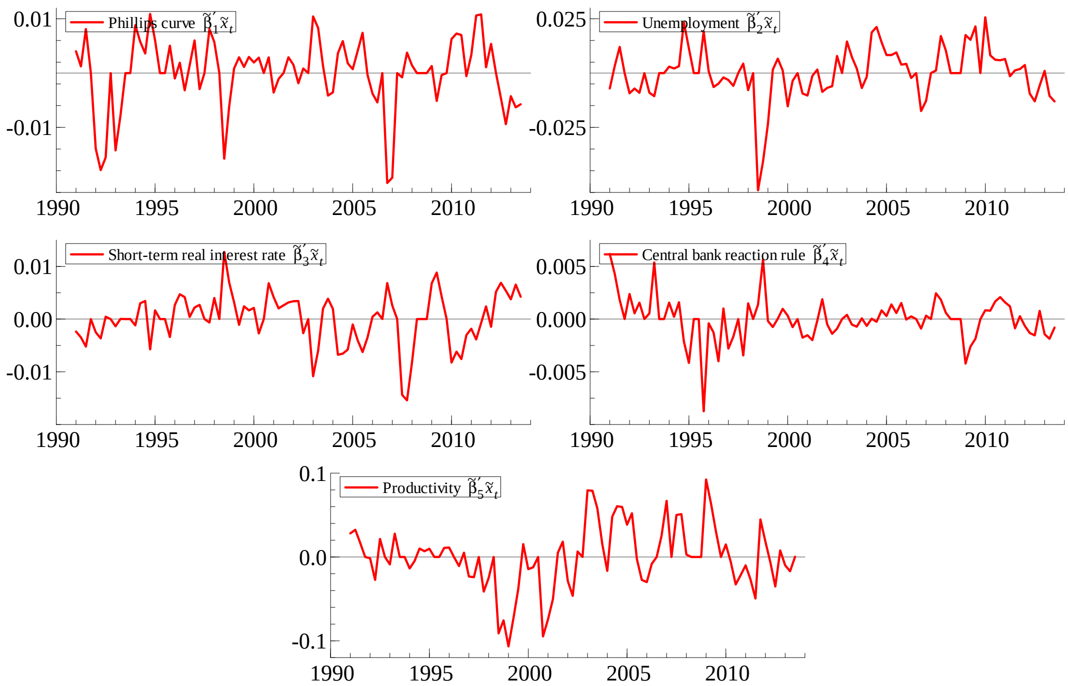
$$\breve{c}_t = -\breve{q}_t + \hat{v}_{5,t} \tag{2.15}$$

where $\breve{c}_t = c_t - 0.002t$ stands for trend-adjusted productivity, $\breve{q}_t = q_t - 0.16d_{s00:4,t}$ stands for the real exchange rate after controlling for an equilibrium mean shift, and $\hat{v}_{5,t} \sim I(0)$ measures the equilibrium error. It shows that the trend adjusted productivity has been negatively co-moving with the real exchange rate. This finding provides empirical support for the predictions of Phelps's structural slumps theory in an IKE world. That is, firms exposed to international competition that face a positive shock to relative costs will improve productivity rather than increase prices to adjust profits in periods of real appreciation.

Figure 2.4 shows the cointegrating relationships and despite some signs of persistent deviations, they seem mean-reverting.

[11] The term $t_{(90:4-98:3),t}$ is equivalent to $t - t_{98:3,t}$.

Figure 2.4: Cointegrating relations $\tilde{\boldsymbol{\beta}}'\tilde{\boldsymbol{x}}_t$



Note 1: The graphs correspond to the cointegrating relationships in model (2.9) where the effects of the short-run dynamics, $\boldsymbol{\Gamma}_1 \triangle \boldsymbol{x}_t$, have been concentrated out. For further details, see Chapter 7 in Juselius (2006).

2.5 Policy implications

According to the IKE theory, the long swings observed in the nominal exchange rate around relative prices will be associated with similar swings in the interest rate spread between the domestic and foreign interest rates. Furthermore, the structural slumps theory predicts that the domestic real interest rate may influence the unemployment rate. Then, the latter is likely to be affected by persistent swings in the real exchange rate and by the monetary policy of central banks that impact interest rates.

The main objective of the Central Bank of Chile is "safeguarding the stability of the currency and the normal functioning of the internal and external payment systems" (BCCh 1989, Section III). To achieve its objective, the Central Bank conducts its monetary policy based on inflation targeting in a free-floating exchange rate regime. The main instrument to keep inflation close to its target is the monetary policy interest rate. Changes in the latter are passed to the interbank interest rate through open-market operations, interest-bearing reserves, and discount-window policy. Commercial banks, in turn, pass these fluctuations to lending and/or deposit rates,[12] which changes consumption, savings, and investments, affecting the aggregate demand and, hence, the price level in the economy.

The results in this chapter indicate that the Central Bank, in pursuing its mandate,

[12] Figure 2.1 shows the co-movement between the monetary policy interest rate and the deposit rates.

may affect not only the unemployment rate but also the natural rate of unemployment. Moreover, the adjustment coefficients show that when the real interest rate is above its long-run equilibrium, it will tend to rise as real interest exhibits an equilibrium error-increasing behavior. Whatever the case, it seems that high and persistent unemployment might be caused by fighting inflation. Thus, Chilean monetary policy must be efficiently combined with other economic policies (e.g., fiscal policy) that stimulate employment in periods in which contractionary monetary policy is at work.

The results also suggest that firms exposed to international competition in a world of imperfect knowledge adjust profits through productivity improvements by laying off the least productive workers. Because Chile has a small and open economy that is fairly and increasingly exposed to international competition, then in periods of real exchange appreciation, policies aimed to reduce unemployment should focus on the tradable sector.

2.6 Conclusion

In this chapter, the author has empirically analyzed the dynamic of inflation and unemployment in Chile using a CVAR model. The results indicate that these dynamics may be described by a Phillips curve when allowing for a positive co-movement between trend-adjusted productivity and unemployment rate. That is, unemployment in excess of trend-adjusted productivity would lead to downward pressure on inflation. Moreover, the natural rate of unemployment is time varying and positively associated with the long-term interest rate, suggesting that monetary policy might not be completely neutral over the business cycle.

The results also show that the trend-adjusted productivity is negatively co-moving with the real exchange rate, implying that during periods of real exchange rate appreciation competitiveness is negatively affected. Consequently, firms adjust profits by increasing productivity rather than prices.

Altogether, the empirical analysis provides support for the Phelps's (1994) structural slumps theory in a world of imperfect knowledge. Phelps states that fluctuations in exchange rates and the real interest rate may explain the fluctuations in unemployment rate and that the natural rate of unemployment is a function of the domestic real interest rate in the economy. The imperfect knowledge theory states that pronounce persistence in real exchange rates are associated with a similar persistence in the real interest rate. Therefore, these theories predict that unemployment and labor productivity increase in periods of rising real interest rates and real exchange appreciation.

A Data

Table A.1: Data source, description and transformation

Variable	Description	Source	Transformation
$\triangle p_t$	Inflation rate	Own elaboration based on data from Central Bank of Chile	Log difference of the CPI
u_t	Unemployment rate	Own elaboration based on data from Central Bank of Chile and National Statistics Institute of Chile	Ratio of unemployed people to labor force
i_t^L	1-year nominal interest rate for deposits	Own elaboration based on data from Central Bank of Chile	Original series is divided by 400
i_t^S	3-months nominal interest rate for deposits	Own elaboration based on data from Central Bank of Chile	Original series is divided by 400
c_t	Real labor productivity	Own elaboration based on data from Central Bank of Chile and National Statistics Institute of Chile	Log of the ratio between real GDP and employed people
q_t	Real exchange rate	Central Bank of Chile	Log of the real exchange index

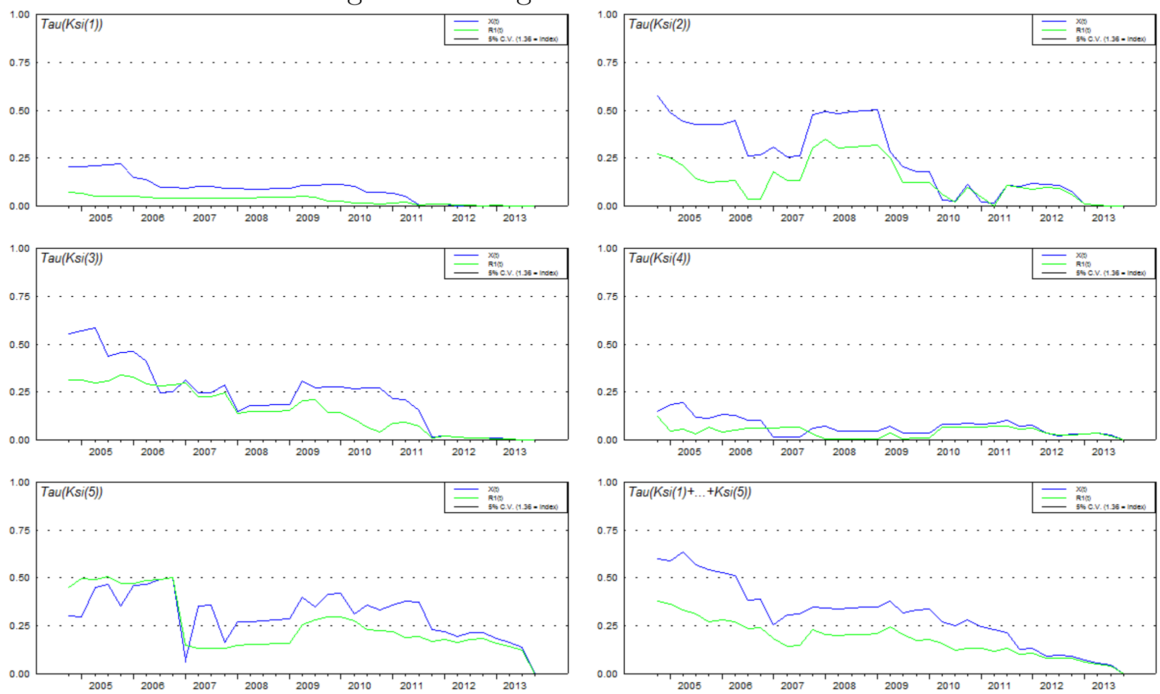
Note 1: Labor series are spliced because the Statistics Institute of Chile changed the employment survey in 2010 from Encuesta Nacional de Empleo (ENE) to Nueva Encuesta Nacional de Empleo (NENE).
Note 2: The data set and the CATS file to replicate the results are available upon request.

B Fluctuation test

Figure A.1 shows the eigenvalue fluctuation test for each individual eigenvalue λ_i, $i = 1, 2, 3, 4$,5 and for the weighted average. The individual fluctuation tests correspond to Tau (Ksi (i)) and the weighted average of them to Tau(Ksi(1)+$\cdots$+Ksi(5)).[13] When the graph is above the unit line, the parameter constancy can be rejected at the 5% level. Based on this critical value, Figure A.1 suggests that there are no signs of parameter-non constancy in the model. This is valid for the full model (2.9), corresponding to the $X(t)$ graph, and for the concentrated model, represented by $R1(t)$ graph where the short-run effects, $\mathbf{\Gamma}_1\triangle\boldsymbol{x}_t$, have been concentrated out of the full model.

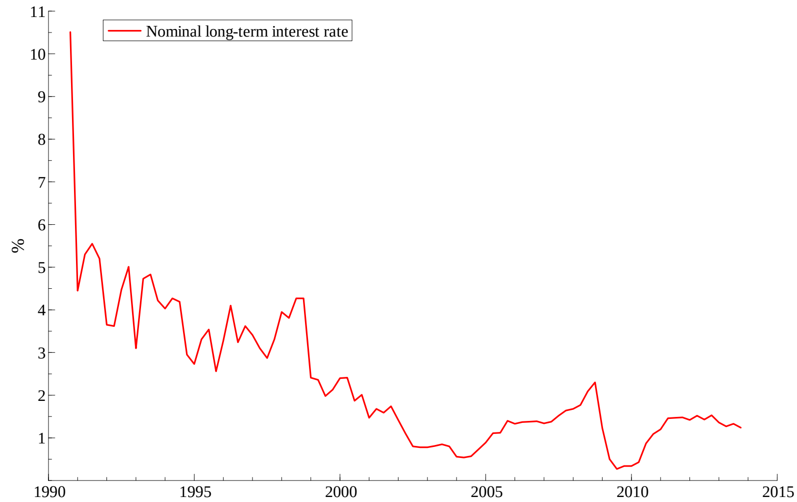
[13]For further details, see Chapter 9 in Juselius (2006).

Figure A.1: Eigenvalue fluctuation test



C Nominal interest rate

Figure A.2: Nominal long-term interest rate



Chapter 3

The Effect of the Minimum Wage on Employment and Hours Worked: The Case of Household Workers in Chile

Leonardo Salazar*

Abstract

In this chapter, the author analyzes the relationship between employment and the minimum wage of household workers. The results, based on a cointegrated vector autoregressive model, show that there is a negative and inelastic long-run relationship between the minimum wage and hours worked, suggesting that increases in the minimum wage lead to an increase of the total income in the household service sector. The results also show that in the long run there is no significant relationship between the real minimum wage of household workers and their employment, implying that employers in the household service sector tend to reduce the number of hours worked per employee instead of the overall number of employees when the minimum wage is increased.

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3.1 Introduction

A household worker is any person engaged in household work within an employment relationship whose work is performed in or for a household or households (ILO, 2013), and the main objective of the minimum wage is to secure a minimum standard of living of the workers in an economy, primarily the least-skilled workers.

In Chile, workers between 18 and 64 years old must receive as a minimum remuneration the national minimum wage. However, household workers, from 1974 to February 2009, were entitled to receive a minimum wage that was on average 73.3% of the national minimum wage. That is, for more than 34 years a minimum wage gap has existed between household workers and people receiving the national minimum wage, indicating that the minimum standard of living for household workers is lower than for other workers. This situation changed in 2008 when the Chilean Government enacted Law 20279 which, based on gradual and predictable rates, increased the minimum wage for household workers until, by March of 2011, it reached 100% of the national minimum wage. The objective of this law was to eliminate the minimum wage gap between workers that did not seem to have any reasonable foundation.

From standard economics textbooks, under perfect competition, a binding minimum wage will be associated with lower levels of employment due to two effects. First, a binding minimum wage over the competitive wage reduces employment as employers substitute the now more expensive labor input with other inputs (e.g., capital). Second, the new minimum wage increases the firm's costs, thus raising prices and reducing labor demand. Furthermore, an increase in the minimum wage may also cause a "labor-labor" substitution; that is, employers substitute less-skilled workers for more-skilled ones (Neumark, 2014).

Empirical evidence shows, however, that the prediction of the competitive model does not always hold. While there are a number of studies finding negative employment elasticities in the interval of -0.3 to -0.1 for the youth (Neumark and Wascher, 2008), other studies report a positive or zero effect of minimum wage on employment (Card and Krueger, 1995; Zavodny, 1998). These results have opened a stimulating debate about the policy effects of the minimum wage (e.g., poverty effects, distribution effects). Generally, US studies dominate the literature. Among the the available studies from 1985 to 2000, 120 studies used US data and only 22 were applied to data from other countries (Gindling and Terrell, 2007). In the case of Chile, the results show a negative relationship between minimum wage and employment in the manufacturing sector and between minimum wage and youth employment. However, for the whole economy, the literature reports employment elasticities fluctuating between -0.5 and 0.2. These fairly large variations might be explained by the choice of different methodologies, data, and analyzed periods (Ramos and Chamorro, 2013).

The literature does not give a clear consensus about how to measure employment. Most studies use the employment-to-population ratio[1] to analyze the effect of the minimum wage on employment. Because in the short run it might be easier for firms to change the number of hours worked rather than employment, this ratio might not be the best measure (Zavodny, 2000).

While some studies describing Chilean household workers exist, so far there is no research that estimates the effect of the minimum wage on their employment. The law of minimum wage for household workers was applied, essentially, without knowing the potential effects of this salary increase on employment. The purpose of this chapter is to estimate the effect of the minimum wage on the number of hours worked and on employment.[2] This is important because a quite large number of household workers are affected by a minimum wage increase. In 2013 approximately 264,159 people worked in the household service sector in Chile, and among these workers, 195,917 received 1.2 times the minimum wage or lower.[3]

Compared to mechanisms commonly observed in other sectors of the economy, the household worker service sector differs in two respects. First, it is composed mostly of low-skilled workers with low levels of education, potentially aggravating the "labor-labor" substitution. For example, in 2013 the average schooling of people in this sector was nine years. Second, the household workers' service is labor intensive, and the substitution of labor for capital is generally low. These features suggest that the effect of the minimum wage on the employment of household workers might be lower than for other groups.

International studies that analyze the effect of the minimum wage on the employment of household workers are sparse. Evidence exists only for South Africa and Brazil. The former shows that the introduction of the minimum wage caused a significant increase in real hourly wages (Hertz, 2005; Dinkelman and Ranchhod, 2012; Bhorat et al., 2013), but the effect on employment and on hours worked was less clear. While Hertz (2005) finds a negative effect on both hours worked and employment, Dinkelman and Ranchhod (2012) find no effect, and Bhorat et al. (2013) find a significant decrease in the number of hours worked, but no effect on employment. In the Brazilian case, regardless of whether the growth of the minimum wage of household workers was above the growth rate of the economy, the demand for household workers was sufficiently high to leave the number of hired workers almost unchanged (Domingues and de Souza, 2012; ILO, 2013).

To elaborate on this topic, the author examines the relationship between the minimum wage and the employment of household workers by estimating a cointegrated vector autoregression (CVAR) model for hours worked, employment, minimum wage of house-

[1] The employment-to-population ratio is the ratio of the number of employed people in the group of interest to the population aged 15 or over.

[2] In this chapter, employment is measured as employment-to-population ratio.

[3] This information was calculated using the Social and Economic Characterization survey (CASEN 2013) of the Ministry of Social Development in Chile.

hold workers, national productivity, and national unemployment rate. To the best of the author's knowledge, this is the first study of the effect of the minimum wage on employment and hours worked in the household service sector in Chile. While previous studies in Chile analyze the effect of minimum wage on the number of employed people, this study considers both hours worked and the employment-to-population ratio as a measure of employment in the household service sector.

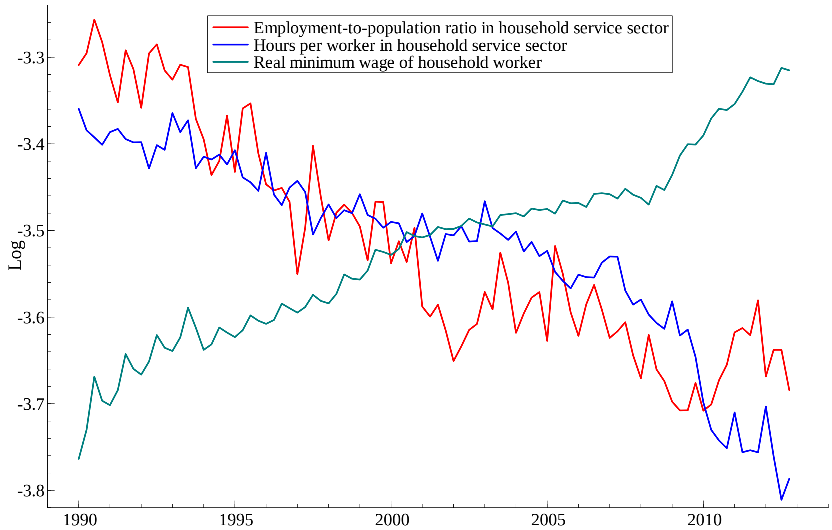
The results show that hours worked and the minimum wage of household workers are negatively related. Specifically, the author shows that an increase of 1% of the real minimum wage of household workers is associated with a decrease of 0.65% in the number of hours worked in the household service sector. However, there is no significant association between minimum wage and employment.

3.2 Stylized facts

The Chilean Labor Code (DT, 2015) establishes that the monthly wage received by a worker cannot be lower than the monthly minimum wage. The level of this varies according to the following categories: (1) workers between 18 and 65 years old (national minimum wage), (2) workers under 18 and over 65 years old, and (3) household workers. Until the first quarter of 2011, the minimum wage of household workers was always lower than the national minimum wage. This situation changed when Law 20279 was implemented, establishing that by March 2011 the household workers' minimum wage had to be equal to the national minimum wage. Accordingly, Figure 3.1 shows a large increase in the real minimum wage of household workers from 2009 to 2011. Also, between 1990 and 1995, after Chile returned to democracy, the real minimum wage of household workers exhibits a large increase due to a general policy of raising minimum wages. The main difference between these two increases is that while the latter policy raised all minimum wages at the same time, Law 20279 increased only the minimum wage for household workers.

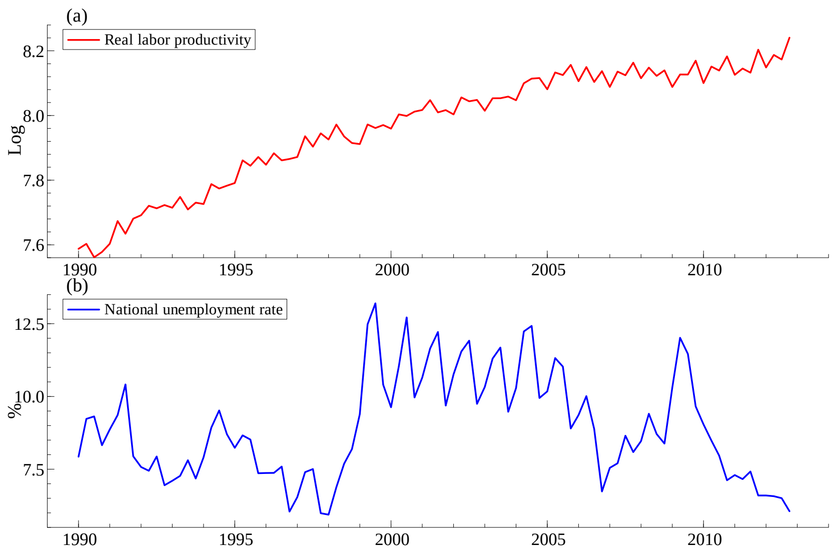
As shown in Figure 3.1, the number of hours worked in the household service sector exhibits a negative trending behavior over time. The large decrease observed around 2009 coincides with the implementation of Law 20279. It suggests that legal increases in the minimum wage for household workers was associated with reductions in the number of hours worked in the household service sector. In the same figure, employment in the household service sector also exhibits a negative trending behavior and is co-moving positively with hours per worker over time. However, from 2010 to 2012 employment shows an increase that is not observed in hours per worker. Also, the employment has a stronger seasonal pattern than hours per worker.

Figure 3.1: Real minimum wage, hours per worker, and employment in the household service sector.



The national unemployment rate, which is shown in Figure 3.2 Panel (b), exhibits an important increase in 1998, possibly as a result of the Asian crisis.[4] However, unemployment is not the only variable seemingly affected by the Asian crisis. Figure 3.2 Panel (a) shows a deceleration in the growth of the real productivity around 1998, which again might indicate a slowdown in the economic activity associated with the Asian crisis. The unemployment rate and productivity exhibit seasonality caused by agricultural activity in Chile, which is higher during the last and first quarter of each year.

Figure 3.2: Panel (a): Real labor productivity. Panel (b): National unemployment rate.



[4]The Asian crisis hit the Chilean economy in 1998. The tradable sector was the most affected since about 48% of the total exports were sent to Asia in 1998. The decrease in Asian demand triggered the bankruptcy of many companies, leading to a large increase in the unemployment rate.

3.3 Theoretical framework

In a neoclassical economy, where labor is the only input used in production, a representative firm is assumed to maximize the following problem:[5]

$$\max \pi (l) = pf (l) - wl \tag{3.1}$$

where l is labor input, p is the price in the economy, w is the competitive market wage, $f(\cdot)$ is a neoclassical production function, and $\pi(\cdot)$ is a profit function. According to (3.1), a representative firm chooses the level of employment that maximizes its profit. The first order condition is the following:

$$p\frac{\partial f (l)}{\partial l} = w \tag{3.2}$$

that is, a representative firm maximizes profits by setting the value of the marginal product of labor, $p\frac{\partial f(l)}{\partial l}$, equal to the marginal cost of labor, w. A competitive firm will, therefore, change the employment level until condition (3.2) is satisfied. The level of the market-clearing equilibrium wage is w^* and the competitive level of employment l^*. The aggregate level of employment in the economy can be calculated as the number of firms times the level of employment given by (3.2).

If a binding minimum wage, w^m, is imposed so that $w^m > w^*$, condition (3.2) will not hold. Specifically, the binding minimum wage will be greater than the value of the marginal product, $w^m > p\frac{\partial f(l)}{\partial l}$. For a constant price level, the only way that a firm can satisfy condition (3.2) is by lowering the level of employment; that is, the firm reduces the total quantity of labor demanded. Hence, the neoclassical competitive model predicts a lower level of employment in the aggregate, but this magnitude depends on labor demand elasticity. The more elastic the labor demand is, the more employment will fall when the wage increases. In addition, the higher salary will tend to increase the quantity of labor supplied.

The reduction in the quantity of labor demanded implies a decrease in total labor hours demanded, but the effect on hours per worker is ambiguous. For example, employers may substitute lower-skilled workers with higher-skilled ones. If the latter works more hours than the former, the hours per workers might increase (Zavodny, 2000).

Employment surveys normally classify a person as employed if she/he worked for at least one hour during the reference week or is only temporally absent from her/his job. Then, using the number of employed people may be misleading because increases in the minimum wage may not reduce the reported employment, even though they have a significant effect on hours worked. Thus, using the number of employed people to calculate elasticities of labor demand may provide a biased estimate of its response to changes in

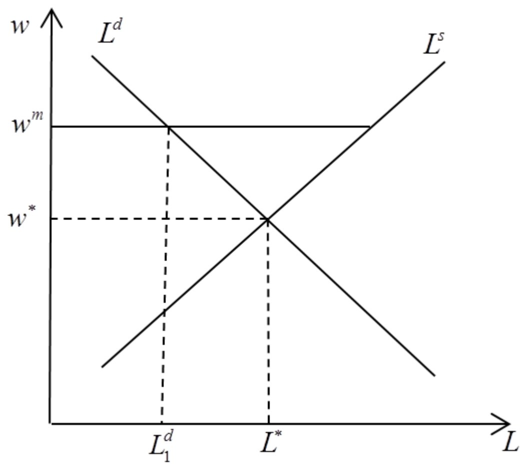
[5]This section is based on Zavodny (1998).

the minimum wage (Couch and Wittenburg, 2001).

Finally, a higher hourly wage generates two simultaneous effects on labor supply. The substitution effect, which suggests that workers offer more hours of work because the higher wage rate increases the opportunity cost of leisure, and the income effect predicts a decrease in the supply of hours worked. Workers who earn a higher income for a given number of hours can afford more leisure. The final effect depends on the relative size of both effects, but altogether it is more likely to observe variations in hours worked rather than changes in employment.

The imposition of a minimum wage is shown graphically in Figure 3.3. Given an upward-sloping labor supply curve, L^s, and a downward-sloping labor demand curve, L^d, the market-clearing equilibrium wage and the associated level of employment are w^* and L^*, respectively. If a minimum wage, $w^m > w^*$, is introduced, then the level of employment will fall until L_1^d.

Figure 3.3: Labor market under binding minimum wage



3.4 The CVAR model and the moving average representation

The error correction form of a VAR model can be written as:[6]

$$\triangle \boldsymbol{x}_t = \boldsymbol{\Pi} \boldsymbol{x}_{t-1} + \sum_{i=1}^{k-1} \boldsymbol{\Gamma}_i \triangle \boldsymbol{x}_{t-i} + \boldsymbol{\mu}_0 + \boldsymbol{\mu}_1 t + \boldsymbol{\Phi} \boldsymbol{D}_t + \boldsymbol{\varepsilon}_t \tag{3.3}$$

where $\boldsymbol{x}_t' = [x_{1,t}, x_{2,t}, \ldots, x_{p,t}]$ is a p-dimensional vector of stochastic variables, $\boldsymbol{\mu}_0$ is an unrestricted constant, t is an unrestricted trend with coefficient matrix $\boldsymbol{\mu}_1$, $\boldsymbol{D}_t$ is a matrix of deterministic terms (e.g., shift dummies, seasonal dummies), and $\boldsymbol{\varepsilon}_t \sim \mathcal{N}_p(\mathbf{0}, \boldsymbol{\Omega})$. Matrices $\boldsymbol{\Gamma}_i$ and $\boldsymbol{\Pi}$ have dimension $p \times p$.

[6]This section is mainly based on Juselius (2006) and Dennis (2006).

If $\mathbf{\Pi}$ has reduced rank, $0 < r < p$, it can be decomposed into $\mathbf{\Pi} = \boldsymbol{\alpha}\boldsymbol{\beta}'$, where $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$ are $p \times r$ matrices of full column rank. The orthogonal complement of matrix $\boldsymbol{z}$ is denoted as $\boldsymbol{z}_\perp$. Now, if $\mathbf{\Pi}$ has reduced rank and $\boldsymbol{\alpha}'_\perp \mathbf{\Gamma} \boldsymbol{\beta}_\perp$ is of full rank, where $\mathbf{\Gamma} = \mathbf{I}_p - \sum_{i=1}^{k-1} \mathbf{\Gamma}_i$, then the process $\boldsymbol{x}_t \sim I(1)$.[7] The r cointegrated relations $\boldsymbol{\beta}'\boldsymbol{x}_t$ define stationary relations among nonstationary variables, potentially representing long-run equilibrium relationships (Juselius, 2006).

The moving average (MA) representation of model (3.3) is the following:

$$\boldsymbol{x}_t = \boldsymbol{C} \sum_{i=1}^{t} (\boldsymbol{\varepsilon}_i + \boldsymbol{\mu}_0 + i\boldsymbol{\mu}_1 + \mathbf{\Xi}\boldsymbol{D}_i) + \boldsymbol{C}^* (L) (\boldsymbol{\varepsilon}_t + \boldsymbol{\mu}_0 + \boldsymbol{\mu}_1 t + \mathbf{\Phi}\boldsymbol{D}_i) + \boldsymbol{A} \tag{3.4}$$

where $\boldsymbol{C}^* (L)$ is an infinite order polynomial depending on the parameters of model (3.3) and $\boldsymbol{A}$ depends on the initial values of $\boldsymbol{x}_t$, satisfying $\boldsymbol{\beta A} = \mathbf{0}$.[8] Finally, $\boldsymbol{C}$ is of dimension $p \times p$, known as the long-run impact matrix, and defined as:

$$\boldsymbol{C} = \underbrace{\boldsymbol{\beta}_\perp \left[\boldsymbol{\alpha}'_\perp \left(\mathbf{I}_p - \sum_{i=1}^{k-1} \mathbf{\Gamma}_i \right) \boldsymbol{\beta}_\perp \right]}_{\tilde{\boldsymbol{\beta}}_\perp} \boldsymbol{\alpha}'_\perp \tag{3.5}$$

where $\boldsymbol{\alpha}_\perp$ determines the $(p - r)$ common stochastic trends, the *pushing forces*, and $\tilde{\boldsymbol{\beta}}_\perp$ can be interpreted as the loadings with which the common trends enter the variables.

3.5 Empirical model analysis

3.5.1 Baseline model

The quarterly data cover the period 1990:1–2012:4, and the information set[9] is given by $\boldsymbol{x}'_t = [e, h, w, c, u]_t$ where

- e_t is employment, measured as the log of the employment-to-population ratio of household workers.
- h_t is hours per worker, defined as the log of the number of hours worked in the household service sector weighted by the number of employed household workers.
- w_t is the log of the real minimum wage of household workers.

[7] If rank $(\mathbf{\Pi}) = p$, the process $\boldsymbol{x}_t \sim I(0)$. If the rank of rank $(\mathbf{\Pi}) = 0$ is zero, the same process is nonstationary and not cointegrated .

[8] From the MA representation (3.4), it follows that the unrestricted constant, $\boldsymbol{\mu}_0$, cumulates to a linear trend and that the unrestricted trend, $\boldsymbol{\mu}_1$, cumulates to a quadratic trend. To avoid the latter, quadratic trends have been restricted to zero in the subsequent analysis. For further information, see Chapter 6 in Juselius (2006).

[9] Appendix A presents the data source and transformations.

- c_t is the log of the national real labor productivity.
- u_t is the national unemployment rate.

The baseline cointegrated VAR model for vector $\boldsymbol{x}_t$ is the following:

$$\triangle \boldsymbol{x}_t = \boldsymbol{\alpha}\tilde{\boldsymbol{\beta}}' \tilde{\boldsymbol{x}}_{t-1} + \boldsymbol{\Gamma}_1 \triangle \boldsymbol{x}_{t-1} + \boldsymbol{\phi}_{d_S} d_{s98:3,t} + \mathbf{1} + \boldsymbol{\phi}_p \boldsymbol{D}_{p,t} + \boldsymbol{\phi}_s \boldsymbol{S}_t + \boldsymbol{\varepsilon}_t \tag{3.6}$$

where $\tilde{\boldsymbol{x}}_t = [\boldsymbol{x}_t, t_{98:3,t}, t]'$, $t_{98:3,t}$ is a broken linear trend taking values of $1, 2, \ldots, 58$ from 1998:3 to 2012:4, 0 otherwise; and t is a linear trend. The shift dummy, $d_{s98:3,t}$, takes the value of 1 since 1998:3, 0 otherwise; $\mathbf{1}$ is a vector of constant terms; $\boldsymbol{D}_{p,t}$ contains four permanent dummies $(0, \ldots 0, 1, 0, \ldots, 0)$;[10] and $\boldsymbol{S}_t$ is a vector of centered seasonal dummies.

The trend is needed to account for the average growth in productivity, in hours worked, in employment, and in the minimum wage of household workers. The broken linear trend accounts for the productivity slowdown. The linear and broken linear trends are restricted to lie in the cointegration space.

3.5.2 Misspecification tests and determining the cointegration rank

Table 3.1 reports the residual misspecification tests of model (3.6). The upper part shows that the baseline model is well specified; there is no evidence of autocorrelation, non-normality, or ARCH effects. The univariate tests, in the lower part of Table 3.1, show weak signs of non-normality in both the residuals of the real minimum wage of the household workers and employment. Except for these minor misspecification signs, the model seems reasonably well specified.

The upper part of Table 3.2 reports the eigenvalues, λ_i, for the null hypothesis of $r = 0, 1, 2, 3, 4$, cointegrating relations, and the corresponding Bartlett-corrected $I(1)$ trace test with the p-value in brackets. Additionally, the last column, α_r, reports the largest absolute t-value of the r-th vector of $\boldsymbol{\alpha}$. For example, 5.26 is the largest absolute t-value in second column of $\boldsymbol{\alpha}$.

According to the trace test, the hypothesis $r = 3$ cannot be rejected with a p-value of 0.60. To check the adequacy of this choice, the lower part of Table 3.2 reports the five largest characteristic roots for $r = 5, 4, 3, 2$. The unrestricted model, $r = 5$, has only one large root, 0.92, suggesting that the model contains one unit root at most. This is further supported by the result that $r = 4$ leaves a complex pair of roots with a modulus of only 0.76 as the largest roots in the system.

[10] Each permanent dummy takes the value of 1 in 1990:4, 1993:3, 2008:1, and 2010:1, respectively; 0 otherwise.

Table 3.1: Misspecification tests CVAR baseline model (3.6)

Multivariate tests				
Autocorrelation		Normality	ARCH	
Order 1 : $\chi^2(25)$	Order 2 : $\chi^2(25)$	χ^2 (10)	Order 1 : $\chi^2(225)$	Order 2 : $\chi^2(450)$
33.57 [0.12]	29.33 [0.25]	15.10 [0.13]	243.22 [0.19]	451.02 [0.48]

Univariate tests					
Equation	$\triangle e_t$	$\triangle h_t$	$\triangle w_t$	$\triangle c_t$	$\triangle u_t$
ARCH Order 2: χ^2 (2)	3.41 [0.18]	1.73 [0.42]	2.11 [0.35]	0.73 [0.69]	1.41 [0.49]
Normality χ^2 (2)	6.06 [0.05]	0.55 [0.76]	7.16 [0.03]	3.29 [0.19]	1.55 [0.76]
Skewness	0.60	0.02	-0.36	0.38	0.08
Kurtosis	3.27	3.07	4.18	2.68	2.52

Note 1: [·] is the p-value of the test.

Furthermore, $\boldsymbol{\alpha}_4$ contains a significant alpha coefficient with an absolute t-value of 3.64 in the employment equation and the fourth unrestricted beta vector shows that it contains information about an economically plausible relationship between employment of household workers and national unemployment rate. Therefore, based on these arguments, the following analysis considers $r = 4$.

Table 3.2: Cointegrating rank and model adequacy

$I(1)$ trace test						
$p-r$	$H_0 : r =$	Eigenvalues (λ_i)	Trace test	p-value	$Q_{.95}$	α_r
5	0	0.52	157.73	[0.00]	97.61	-
4	1	0.44	95.93	[0.00]	71.78	8.01
3	2	0.31	48.21	[0.06]	49.06	5.26
2	3	0.14	18.32	[0.60]	30.54	5.82
1	4	0.05	4.54	[0.84]	15.01	3.64

Modulus of the five largest characteristic roots					
$r = 5$	0.92	0.77	0.77	0.64	0.45
$r = 4$	1.00	0.76	0.76	0.47	0.47
$r = 3$	1.00	1.00	0.76	0.76	0.47
$r = 2$	1.00	1.00	1.00	0.53	0.53

Note 1: p is the number of variables in vector $\boldsymbol{x}_t$.

Note 2: [·] is the p-value of the trace test simulated according to model (3.6).

Note 3: $Q_{.95}$ is the 5% critical value of the trace test.

Note 3: α_r is the largest absolute t-value of the r-th column in $\boldsymbol{\alpha}$.

3.5.3 The pulling forces

To identify economically plausible relationships among the cointegrating vectors, a set of restrictions, $\mathscr{H}_{\tilde{\beta}} : \tilde{\boldsymbol{\beta}} = (\boldsymbol{H}_1\boldsymbol{\varphi}_1, \boldsymbol{H}_2\boldsymbol{\varphi}_2, \ldots, \boldsymbol{H}_r\boldsymbol{\varphi}_r)$, must be imposed on $\tilde{\boldsymbol{\beta}}$, where $\boldsymbol{H}_i$ is a restriction matrix of dimension $p_1 \times (p_1 - m_i)$, p_1 is the dimension of $\tilde{\boldsymbol{x}}_t$, m_i is the number of restrictions, $p_1 - m_i$ is the number of freely varying parameters, and $\boldsymbol{\varphi}_i$ is a matrix of unknown parameters of dimension $(p_1 - m_i) \times 1$. This test is distributed as χ^2 with degrees of freedom equal to $\sum_{i=1}^{r} (m_i - (r-1))$ (Johansen, 1996). When $\alpha_{ij}\beta_{ij} < 0$ (> 0), the cointegrating relation is equilibrium correcting (increasing) in the equation $\triangle \boldsymbol{x}_{j,t}$ (Juselius, 2006).

Table 3.3 reports an identified structure on $\tilde{\boldsymbol{\beta}}$ with three overidentified restrictions that could not be rejected based on $\chi^2(3) = 1.54$ with a p-value of 0.67.[11] To facilitate interpretation, an α_{ij} coefficient in boldface means that the cointegrating relation i is equilibrium correcting, whereas an error-increasing coefficient is in italics. The results in Table 3.2 showed that all eigenvalues are inside the unit circle. Thus, the system is stable, and any error-increasing behavior is compensated by error-correcting behavior.

Table 3.3: The estimated long-run $\tilde{\boldsymbol{\beta}}$ structure ($\chi^2(3) = 1.54$ [0.67])

	e_t	h_t	w_t	c_t	u_t	$t_{98:3,t}$	t
$\tilde{\boldsymbol{\beta}}_1'$	-	1.00	0.65 (12.91)	-	-	-	−0.002 (−3.64)
$\boldsymbol{\alpha}_1'$	∗	**−0.18** (−3.03)	**−0.57** (−8.10)	0.14 (2.44)	5.19 (2.22)		
$\tilde{\boldsymbol{\beta}}_2'$	-	-	1.00	-	0.13 (8.31)	-	-
$\boldsymbol{\alpha}_2'$	0.13 (3.41)	∗	*0.02* (1.03)	−0.07 (−3.87)	**−3.10** (−3.88)		
$\tilde{\boldsymbol{\beta}}_3'$	-	-	-	1.00	−0.02 (−8.04)	0.006 (9.81)	−0.01 (−20.18)
$\boldsymbol{\alpha}_3'$	∗	0.19 (2.27)	∗	**−0.49** (−5.99)	*−7.46* (−2.30)		
$\tilde{\boldsymbol{\beta}}_4'$	1.00	-	-	-	0.03 (6.07)	−0.004 (−2.99)	0.009 (7.30)
$\boldsymbol{\alpha}_4'$	**−0.43** (−4.03)	∗	∗	∗	**−3.15** (−1.50)		

Note 1: t-values are given in (·).

Note 2: "-" stands for a zero restriction.

Note 3: "∗" stands for and alpha coefficient with |t-value| ≤ 2.0.

Note 4: a coefficient in boldface stands for an equilibrium error correcting behavior .

Note 5: a coefficient in italics stands for an equilibrium error increasing behavior.

The first cointegrating vector, $\tilde{\boldsymbol{\beta}}_1'\tilde{\boldsymbol{x}}_t$, is interpreted as a relation for hours worked and is expressed as:

[11] Appendix B indicates that the eigenvalue fluctuation test does not show signs of parameter-no constancy.

$$h_t = -0.65w_t + 0.002t + \hat{v}_{1,t} \tag{3.7}$$

where $\hat{v}_{1,t} \sim I(0)$ measures the equilibrium error, that is, the excess of hours worked (positive or negative) in the economy at time t relative to the trend-adjusted minimum wage. Equation (3.7) shows that the number of hours worked in the household service sector are negatively associated with the trend-adjusted minimum wage of the household workers. An increase of 1% in the real minimum wage of household workers above the long-run trend is associated with a decrease of 0.65% in the number of hours worked in the household service sector. Thus, it can be interpreted as a labor demand relation for household workers.

Given that the variables in (3.7) are in logs, the coefficient of -0.65 can be interpreted as a labor demand elasticity, implying an inelastic labor demand for household workers in terms of hours worked. Considering the period in which Law 20279 was implemented, 2009–2011, the percentage change of the real minimum wage and hours worked was 37.75% and -14.26%, respectively, resulting in a positive change in total income of 23.48%.[12]

The adjustment coefficients associated with the labor demand relation show that both hours worked in the household service sector and the minimum wage of the household workers are equilibrium error correcting. That is, if the real minimum wage is above its long-run trend value, then hours of work and the real minimum wage will tend to decrease to restore the equilibrium. Furthermore, a real minimum wage above its long-run trend value will lead to an increase in real productivity and in the national unemployment rate as the latter two are positively affected by the equilibrium error in this relationship.

The second relationship, $\tilde{\boldsymbol{\beta}}_2'\tilde{\boldsymbol{x}}_t$, is interpreted as a wage relation and is expressed as:

$$w_t = -0.13u_t + \hat{v}_{2,t} \tag{3.8}$$

where $\hat{v}_{2,t} \sim I(0)$ measures the equilibrium error. It describes a negative co-movement between real minimum wage of household workers and the state of the economy as measured by the aggregate unemployment rate. Because unemployment is the key variable used to determine the minimum wage adjustments in Chile, equation (3.8) is consistent with the Chilean mechanism for adjusting the minimum wage.

The adjustment coefficients show that unemployment rate is equilibrium error correcting to equation (3.8), whereas the real minimum wage of the household worker is equilibrium error increasing, though the coefficient is not very significant. Then, if the real minimum wage of the household workers is above its long-run equilibrium value, it will tend to increase, which further increases the equilibrium error. At the same time, national unemployment will tend to decrease to restore the equilibrium. As a consequence, the employment in the household service sector will tend to increase. Moreover, keeping

[12] If the period of 2009–2012 is considered, the percentage change in total income is 24.14%.

the real minimum wage of the household workers and/or the unemployment rate above their long-run equilibrium values would lead to a decrease in labor productivity.

The third vector, $\tilde{\boldsymbol{\beta}}_3'\tilde{\boldsymbol{x}}_t$, is interpreted as a trend-adjusted productivity relationship and is written as:

$$\tilde{c}_t = 0.02u_t + \hat{v}_{3,t} \tag{3.9}$$

where $\tilde{c}_t = c_t + 0.006t_{98:3,t} - 0.01t$ stands for trend-adjusted productivity, and $\hat{v}_{3,t} \sim I(0)$ measures the equilibrium error. It shows that unemployment has been positively co-moving with trend-adjusted productivity over time. The rationale is as follows. Firms exposed to international competition facing a positive shock to relative costs cannot generally count on exchange rates to restore competitiveness in a world of imperfect knowledge. Thus, firms will be prone to adjust profits rather than prices. Profits can be adjusted through improvements in productivity by laying off the least productive workers.

The adjustment coefficients show that productivity is equilibrium error correcting to equation (3.9), whereas the national unemployment rate is equilibrium error increasing. Thus when the unemployment rate is above its long-run equilibrium value, it tends to increase as enterprises continue to improve labor productivity. This is likely to generate unemployment persistence that is consistent with observed behavior. In addition, hours worked in the household service sector have been positively affected by the equilibrium error in this relation.

The last vector, $\tilde{\boldsymbol{\beta}}_4'\tilde{\boldsymbol{x}}_t$, is interpreted as an employment relation and is expressed as:

$$e_t = -0.03u_t + 0.004t_{98:3,t} - 0.009t + \hat{v}_{4,t} \tag{3.10}$$

where $\hat{v}_{4,t} \sim I(0)$ is the equilibrium error. This relationship shows that the trend-adjusted employment, $e_t - 0.004t_{98:3,t} + 0.009t$, is negatively associated with the national unemployment rate. The adjustment coefficients show that both the employment and national unemployment rates are equilibrium error correcting to equation (3.10), the latter with a coefficient that is borderline significant. Then, if the national unemployment rate is above its long-run equilibrium value, the employment in the household service sector will tend to decrease to restore the equilibrium.

The null hypothesis of a stationary relationship between the real minimum wage and the employment of household workers was rejected based on $\chi^2(2) = 9.29$ with a p-value of 0.00 and, when a trend is included in the analysis, on $\chi^2(1) = 4.43$ with a p-value of 0.03. Thus, both the rejection of a stationary relationship between employment and minimum wage and the finding of a labor demand for household workers, equation (3.7), supports the idea that employers tend to reduce hours worked per employee instead of the number of employees when there is an increase in the minimum wage. But the results also showed that an increase in the minimum wage is associated with a less-than-proportional reduction in the number of hours worked. Therefore, a minimum wage increase seems to lead to an increase in the total monthly income of household workers without a corresponding reduction in the employment.

3.5.4 The pushing forces

Table 3.4 reports the estimates of MA representation model. It was derived under the joint restriction of a unit vector in $\boldsymbol{\alpha}$[13] for productivity, national unemployment rate, and employment. This hypothesis was not rejected based on $\chi^2(3) = 4.91$ with a p-value of 0.18, implying that shocks to these variables have a transitory effect on the real minimum wage of household workers and hours worked in the household service sector.

Table 3.4: MA representation

	Common trend $\boldsymbol{\alpha}_{\perp}$				
	$\hat{\varepsilon}_e$	$\hat{\varepsilon}_h$	$\hat{\varepsilon}_w$	$\hat{\varepsilon}_c$	$\hat{\varepsilon}_u$
$\boldsymbol{\alpha}'_{\perp}$	0.00	-3.68 (-7.67)	1.00 (2.83)	0.00	0.00
	Loading $\tilde{\boldsymbol{\beta}}_{\perp}$				
$\tilde{\boldsymbol{\beta}}'_{\perp}$	0.08	-0.18	0.26	-0.04	-2.85

Note 1: $(\cdot)$ is the t-value

Note 2: ε_i are the empirical residuals of model (3.6)

The common stochastic trend, $\boldsymbol{\alpha}'_{\perp}\sum_{i=1}^{t}\varepsilon_i$, shown in the upper part of Table 3.4, is measured by the sum of cumulated empirical shocks to the real minimum wage corrected for hours worked in the household service sector, that is the following:

$$\boldsymbol{\alpha}'_{\perp}\sum_{i=1}^{t}\boldsymbol{\varepsilon}_i = \sum_{i=1}^{t}\left(\hat{\varepsilon}_w - 3.68\hat{\varepsilon}_h\right)_i \tag{3.11}$$

where $\hat{\varepsilon}_w$ and $\hat{\varepsilon}_h$ are, respectively, the empirical residuals of the real minimum wage and hours worked. It shows that shocks to real minimum wages corrected for hours worked in the household service sector have been the main driving force in the analyzed period.

The estimates of $\tilde{\boldsymbol{\beta}}'_{\perp}$, in the lower part of Table 3.4, show how the common stochastic trend loads into the variables. These coefficients indicate that the cumulated empirical shocks to the real minimum wage corrected by hours worked loads positively, but with a tiny coefficient, into employment and negatively into unemployment and productivity.

[13]When a variable has a unit vector in $\boldsymbol{\alpha}$, shocks to this variable have only transitory effects on any of the system's variables. For further information, see Chapter 11 in Juselius (2006).

3.6 Conclusion

In this chapter, the author has empirically analyzed the relationship between employment in the household service sector and the minimum wage of household workers in Chile. The results, based on a CVAR model, indicate that there is a negative long-run relationship between hours worked in the household service sector and the real minimum wage of the household workers. The elasticity of labor demand in the household service sector is -0.65, implying that the demand for household workers is inelastic. Because of that, policies raising the minimum wage of household workers have had a positive impact on the total income in the household service sector.

In addition, the results show that there is no significant relationship between real minimum wage of the household workers and the employment. This suggests that increases in the minimum wage may not be associated with a reduction in employment. Thus, employers in the household service sector seem to reduce the number of hours worked per employee instead of the overall number of employees when there is a minimum wage increase, so that policies raising the minimum wage are likely to be effective in reducing the wage gap as they seem to have increased the total income of household workers without a negative effect on employment.

A Data

Table A.1: Data source and transformation

Variable	Description	Source	Transformation
h_t	Hours worked in the household service sector during the reference week	Own elaboration based on data from National Statistics Institute of Chile	Log of average hours weighted by the number of household workers
e_t	Employment-to-population ratio	Own elaboration based on data from National Statistics Institute of Chile	Log of the ratio between the number of employed household workers and the population aged 15 and over
w_t	Real minimum wage of the household workers	Own elaboration based on data from National Statistics Institute of Chile and Central Bank of Chile	Log of minimum wage of household workers deflated by the consumer price index
c_t	Real labor productivity	Own elaboration based on data from Central Bank of Chile and National Statistics Institute of Chile	Log of the ratio between real GDP and employed people
u_t	Unemployment rate	Own elaboration based on data from National Statistics Institute of Chile	-

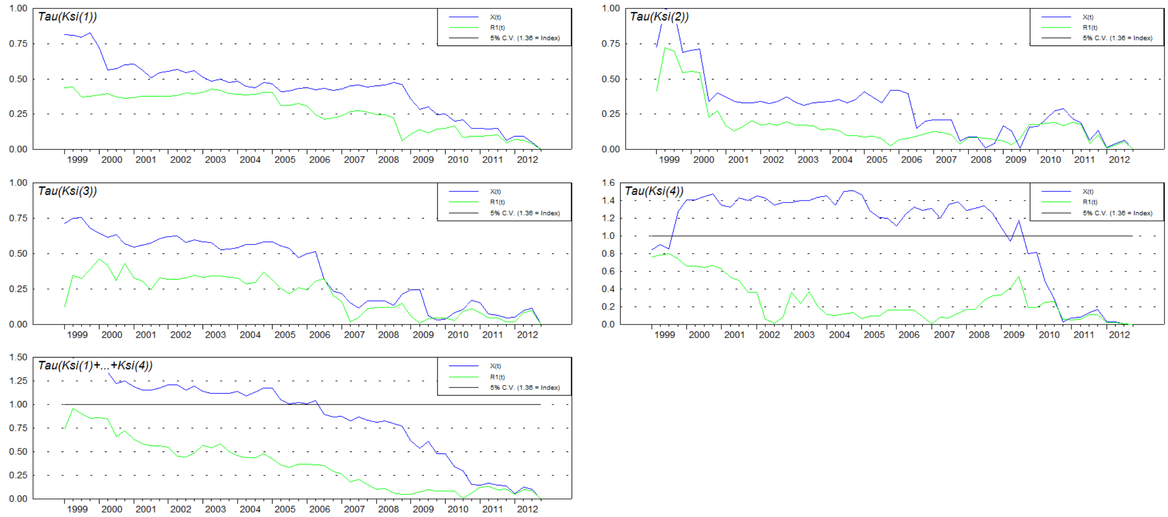
Note 1: Labor series are spliced because the Statistics Institute of Chile changed the employment survey in 2010 from Encuesta Nacional de Empleo (ENE) to Nueva Encuesta Nacional de Empleo (NENE).
Note 2: The data set and the CATS file to to replicate the results are available upon request.

B Fluctuation test

Figure A.1 shows the eigenvalue fluctuation test for each individual eigenvalue and for the weighted average. For each fluctuation test, two graphs are presented. $X(t)$ represents the full model, and $R1(t)$ plots the concentrated model where the short-run effects, $\mathbf{\Gamma}_1 \triangle \boldsymbol{x}_t$, have been concentrated out of the full model. When the graph is above the unit line, the parameter constancy can be rejected at the 5% level.[14]

The individual fluctuation tests correspond to $\mathrm{Tau}\,(\mathrm{Ksi}\,(i))$ and the weighted average to $\mathrm{Tau}\,(\mathrm{Ksi}\,(1) + \cdots + \mathrm{Ksi}\,(4))$. According to these graphs there are no signs of parameter-non constancy in the concentrated model.

Figure A.1: Eigenvalue fluctuation test



[14]For further details about the eigenvalue fluctuation test, see Chapter 9 in Juselius (2006)

Chapter 4

The Impact of Monetary Policy on a Labor Market with Heterogeneous Workers: The Case of Chile

Carlos Madeira* and Leonardo Salazar†

Abstract

We use a factor-augmented vector autoregressive (FAVAR) model to analyze the effect of a contractionary monetary shock on macroeconomic aggregates and labor market indicators for different demographic groups in Chile classified by industry, age, and income quintile. We find that for most groups the job-separation rate and idiosyncratic earnings volatility increase after interest rate rise. The response of the job-finding rate is more mixed with decreases in some groups and an increase for others after an interest rate shock. The labor market in the primary sector is the least sensitive to monetary shocks.

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4.1 Introduction

The economic effects of monetary policy have been broadly studied using different empirical and theoretical approaches, with most studies showing that monetary shocks have a real effect on output and that inflation responds negatively to a contractionary monetary policy shock (Christiano et al., 1999). Furthermore, recent empirical studies for the United States show that the welfare costs of recessions are significantly larger if one accounts for job displacement risk (Krebs, 2003, 2007) and its heterogeneous impact on different agents (De Santis, 2007). Wage volatility in the United States is counter-cyclical (Storesletten et al., 2001, 2004), especially among workers experiencing unemployment spells (McKay and Papp, 2011). Households face substantially larger earnings shocks during recessions (Storesletten et al., 2001, 2004; McKay and Papp, 2011; Davis and von Wachter, 2011; Guvenen et al., 2014), and these earnings losses are highly persistent (Davis and von Wachter, 2011).

In this study we examine how different groups of workers in Chile react to the business cycle and to changes in monetary policy. We analyze the effect of a contractionary monetary shock on the job-finding and separation rates, wage volatility, and labor productivity of different groups of workers. Specifically, we classify workers into 45 distinct groups by (a) economic sector (primary, secondary, tertiary), (b) age (16–35, 36–54, and 55 or older), and (c) income quintile (with lowest income being quintile 1 and highest income being quintile 5). We then estimate a factor-augmented vector autoregressive (FAVAR) model to analyze the transmission effect of a contractionary monetary shock on the labor market experiences of the different Chilean demographic groups.

The FAVAR impulse–response functions show that the job-separation rate and wage volatility tend to increase after a contractionary monetary shock. However, economic sectors in the economy react in different ways. The secondary sector shows a significant reaction in terms of an increase in both the job-separation rate and the idiosyncratic wage volatility. The job-finding rate has a mixed response to an increase in the interest rate, although for some demographic groups in the secondary and tertiary sectors the job-finding rate does fall significantly after a contractionary monetary shock. Overall, the primary sector reacts less both in terms of employment flows, whether as job separation or creation, and in terms of wage volatility.

To measure the labor market experiences of different Chilean workers, we use a rich data set constructed by Madeira (2015). Using the National Employment Survey, which covers a sample of 35,000 households at a quarterly frequency, Madeira (2015) estimated the job-separation rate (the probability that an employed worker will lose his or her job in the next three months), the job-finding rate (the probability that an unemployed worker will find a job within three months), and their wage volatility (the standard deviation of the annual change in labor earnings). The results show that Chile has a fluid labor market

(Jones and Naudon, 2009), with unemployment inflow and outflow rates similar to those in the United States and substantially higher than those in other OECD countries (Elsby et al., 2013). Also, the average employed worker faces idiosyncratic income shocks with a standard deviation of 18%, which is roughly similar to estimates for other countries (Krueger et al., 2010).

Relative to previous studies of the business cycle in Chile such as Del Negro and Schorfheide (2008), we innovate by using measures of how monetary policy and the business cycle affect heterogeneous workers and different economic sectors. For instance, ours is the first work to use a measure of real labor productivity growth for each of the three economic sectors in Chile, and we show that productivity growth is strongly correlated for all sectors, which seems to confirm that labor markets in Chile are flexible enough to allow labor flows to equalize labor productivity across different industries. We also find that unemployment, separation and job-finding rates, and wage volatility are heterogeneous across worker types, yet there is a strong cyclical component that affects all groups. Low-income workers experience both higher unemployment rates and wage volatility. However, low-income workers have a higher job-finding rate and therefore face shorter unemployment spells, perhaps because their job matches involve less specific human capital.

In addition to being related to studies of workers' heterogeneous income shocks during the business cycle (Storesletten et al., 2001, 2004), our study relates to the empirical research on the cyclical fluctuations of the labor market (Pappa, 2009; Trigari, 2009; Mumtaz and Zanetti, 2012; Madeira, 2014). Estimating a structural vector autoregressive (SVAR) model, Ravn and Simonelli (2007) concluded that hours worked, employment, vacancies, and the vacancies-unemployment ratio decrease in response to an increase in the federal funds rate. Moreover, labor productivity first declines briefly and then increases after a few quarters. Monetary policy also affects real wages, which seems inconsistent with the high degree of nominal rigidity in the labor market. Using a standard VAR reduced form, Olivei and Tenreyro (2007) found that an expansionary monetary shock increases wages and hours. Moreover, the response of wages is mildly procyclical, whereas hours react more significantly when the shock occurs in the first or second quarter of the calendar year. Peneva (2013) showed that hourly earnings respond positively to an expansionary monetary shock, a response that is similar across both tertiary and goods sectors. Braun et al. (2009) estimated a SVAR model for the United States that includes both demand and supply shocks. They found that an expansionary monetary shock increases vacancies and the job-finding and job-creation rates, whereas it decreases the separation and job-destruction rates. Finally, they concluded that responses induced by supply shocks are more persistent than those induced by demand shocks.

This chapter is organized as follows. Section 4.2 summarizes the evolution of the labor productivity, employment flows, and wage volatility for the primary, secondary, and tertiary sectors in Chile over the last 23 years. Section 4.3 describes the structure of the

FAVAR model estimated from the macro variables and the labor market statistics for each of the 45 demographic groups in our data. Finally, Section 4.4 summarizes the main results, while Section 4.5 summarizes our conclusions.

4.2 The evolution of labor markets in Chile

We now describe the evolution of the macroeconomic series and the labor markets in Chile. Table 4.1 shows that the quarterly CPI growth rate fluctuates between values as low as 0.12% (the 10th percentile for all periods between 1996 and 2012) and values as high as 1.81% (the 90th percentile for all periods). It is worth noting that because these are quarterly values, 1.81% corresponds to an annualized inflation of 7.24%. The median and average quarterly CPI for 1996–2012 is 0.52% and 0.88%, respectively, which are values well within the bands of 2% to 4% for the annual inflation target followed by the Chilean Central Bank. All the economic sectors in Chile have exhibited a robust real productivity growth at average rates between 0.47% and 0.66%.[1] The primary sector was the industry with both the highest mean productivity growth and the most volatile one with rates ranging from as low as −2.71% (the 10th percentile observed for 1996–2012) and as high as 4.45% (the 90th percentile during the same period). Perhaps the secondary sector was the least developed of all the three main economic sectors and, therefore, the one with the largest gains to make. It is also notable that the tertiary sector has much less volatility in real productivity growth than the other sectors, which can be interpreted as the result of the primary and secondary sectors being more subject to international competition and open economy shocks.

Table 4.1: Distribution of the growth rates (%) of consumer price index and real productivity by percentile. Quarterly data 1996:1–2012:4

Variable (quarterly growth)	Mean	P10	P25	P50	P75	P90
Consumer price index (CPI)	0.88	0.12	0.52	0.80	1.21	1.81
Real productivity all sectors (PRO)	0.61	−1.06	−0.04	0.50	1.65	2.04
Real productivity primary sector (PRO1)	0.66	−2.71	−0.63	0.76	2.33	4.45
Real productivity secondary sector (PRO2)	0.47	−2.22	−0.59	0.50	2.25	3.72
Real productivity tertiary sector (PRO3)	0.62	−0.62	0.09	0.64	1.33	2.05

Note 1: Pi, $i = 10, 25, 50, 75, 90$, is the percentile used to classified the growth rate of the variables.

Figure 4.1 plots the actual evolution of quarterly growth rates of the CPI and the real productivity for each economic sector for 1996–2012. The plot shows that CPI and the real productivity of the three sectors of economic activity are relatively uncorrelated

[1] The primary sector in the Chilean economy, which considers agriculture and forestry, fishery, and mining activities, accounted for 14.8% of the real GDP in 2012. The secondary sector, corresponding to the manufacturing industry, represented 10.5% of the real GDP. Finally, the tertiary sector accounted for 66.1% of the real GDP. The percentages do not sum 100 because the real values are calculated using chained volume series that lose the additive property.

over this period. However, most of the positive and negative spikes in real productivity growth coincide for the primary, secondary, and tertiary sectors, which shows that all labor markets are strongly correlated. This is a sign that labor markets in Chile are integrated across different industries because labor flows would tend to equalize productivity across the different economic sectors.

Figure 4.1: Consumer price index (CPI) and real productivity growth by industrial sector. Quarterly data 1996:1–2012:4

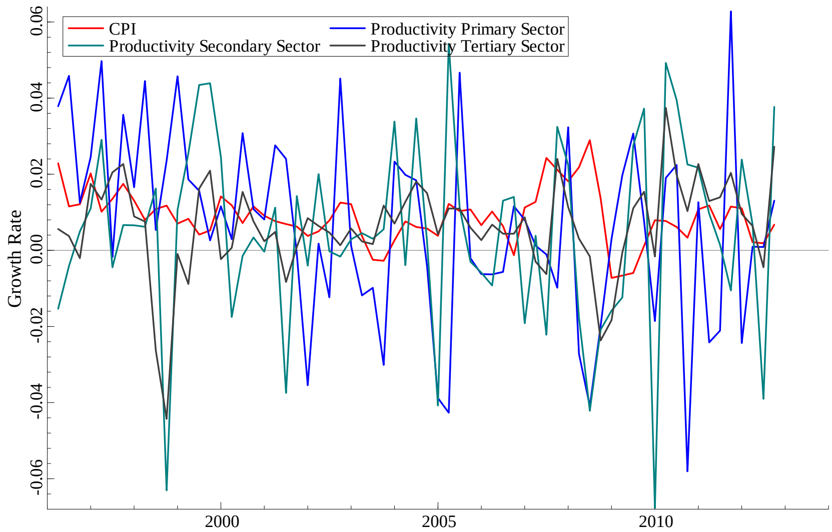


Table 4.2 reports the correlation coefficients between each pair of aggregate variables for 1996–2012. We report the correlation both for the variables in their pure form and for their Hodrick-Prescott (HP) cyclical components, which, in a few cases, give different results. In general, CPI fluctuations have a low correlation with real productivity growth. CPI growth is negatively correlated with the unemployment rate and the separation rate but is positively correlated with the job-finding rate. This is evidence for the traditional short-term Phillips curve, with inflation and unemployment correlated negatively (Christiano et al., 1999). The real productivity growth for the labor force has a low correlation with labor market variables such as unemployment, separation, and job-finding rates, which might be interpreted as evidence for short-term rigidity in the Chilean labor market. Wage volatility is positively correlated with the unemployment rate and negatively correlated with the job-finding rate, showing that in Chile idiosyncratic wage risk also increases during recessions in a manner similar to that in the United States (Storesletten et al., 2001, 2004; McKay and Papp, 2011). Both the separation and job-finding rates have a high correlation with unemployment fluctuations, which shows that, as in the case of other countries, both job creation and destruction play a role in unemployment fluctuations (Fujita and Ramey, 2009; Elsby et al., 2013).

In Table 4.3 we report the same correlation matrix for the HP cyclical components of

Table 4.2: Correlation coefficients of CPI and real productivity growth with the overall unemployment flows and wage volatility. Quarterly data 1996:1–2012:4

Correlation of standard variables (%)						
	CPI	PRO	U	EU	UE	STDI
Consumer price index (CPI)	100					
Real productivity (PRO)	−2.5	100				
Unemployment rate (U)	−27.3	−12.4	100			
Separation rate (EU)	−19.5	−2.8	55.2	100		
Job-finding rate (UE)	28.7	8.4	−71.2	13.5	100	
Wage volatility (STDI)	−20.8	10.4	38.5	13.1	−34.0	100
Correlation of Hodrick-Prescott cyclical component (%)						
	CPI	PRO	U	EU	UE	STDI
Consumer price index (CPI)	100					
Real productivity (PRO)	−3.7	100				
Unemployment rate (U)	−12.3	3.6	100			
Separation rate (EU)	−14.7	−2.4	71.2	100		
Job-finding rate (UE)	11.3	−14.6	−21.6	45.1	100	
Wage volatility (STDI)	−15.1	15.9	24.6	5.8	−26.2	100

each of the three economic sectors. Real productivity growth for the primary, secondary, and tertiary sectors is correlated with the overall economy's productivity growth of 54%, 72%, and 89%, respectively. This is strong evidence that labor flows occurs across different economic sectors and that long-term productivities tend to equalize. CPI growth is negatively correlated with the unemployment and separation rates only in the primary and secondary sectors. However, the job-finding rate is positively correlated with CPI growth in the secondary and tertiary sectors. This shows that the primary sector reacts to inflation shocks mostly in terms of job destruction, while the tertiary sector reacts to inflation shocks in terms of job creation. The secondary sector, however, reacts to inflation shocks both in terms of job creation and destruction. In all economic sectors the unemployment and job-separation rates are highly positively correlated, with coefficients between 69% and 82% for each sector, but the correlation of unemployment with the job-finding rate is much lower. This evidence seems to point out that job destruction is responsible for most of the cyclical movement in unemployment in Chile. Wage volatility is high and positively correlated with unemployment fluctuations only in the secondary and tertiary sectors. Therefore, in Chile only in the secondary and tertiary sectors does a simultaneous cycle of both high unemployment and high idiosyncratic wage volatility affect workers.

Finally, Figures 4.2, 4.3, and 4.4 show the evolution of the labor market variables (wage volatility, unemployment, separation, and job-finding rates) for the primary, secondary, and tertiary sectors, respectively. For each sector we show the separate evolution for the

Table 4.3: Correlation coefficients of CPI and real productivity growth with unemployment flows and wage volatility by economic sector. Quarterly data 1996:1–2012:4

Sector	Correlation of Hodrick-Prescott cyclical component (%)							
		CPI	PRO	PRO1	U	EU	UE	SDTI
Primary	Consumer price index (CPI)	100						
	Real productivity (PRO)	−3.7	100					
	Productivity primary sector (PRO1)	−11.7	54.3	100				
	Unemployment rate (U)	−12.8	2.6	13.1	100			
	Separation Rate (EU)	−10.6	5.3	5.3	81.8	100		
	Job-finding Rate (UE)	−3.7	−2.3	−6.1	15.5	64.1	100	
	Wage volatility (SDTI)	−7.4	18.6	0.6	3.9	0.7	−5.3	100
		CPI	PRO	PRO2	U	EU	UE	SDTI
Secondary	Consumer price index (CPI)	100						
	Real productivity (PRO)	−3.7	100					
	Productivity secondary sector (PRO2)	−14.0	72.1	100				
	Unemployment rate (U)	−26.4	−0.4	7.3	100			
	Separation rate (EU)	−27.7	−8.3	2.4	77.3	100		
	Job-finding rate (UE)	14.9	−17.5	−12.9	−36.4	18.1	100	
	Wage volatility (SDTI)	−17.4	17.0	16.6	34.3	20.0	−32.7	100
		CPI	PRO	PRO3	U	EU	UE	SDTI
Tertiary	Consumer price index (CPI)	100						
	Real productivity (PRO)	−3.7	100					
	Productivity tertiary sector (PRO3)	7.0	88.9	100				
	Unemployment rate (U)	0.1	6.3	3.5	100			
	Separation rate (EU)	−2.2	−1.8	−1.3	68.6	100		
	Job-finding rate (UE)	13.9	−16.4	−9.1	−11.1	53.8	100	
	Wage volatility (SDTI)	−14.6	14.3	13.5	23.2	0.3	−29.0	100

workers in each national income quintile (with quintile 1 representing the lowest income and 5 the highest). Several facts stand out. First, a significant seasonality exists in the unemployment, separation, and job-finding rates, which is strongest for the primary sector. Second, the shocks affecting all workers have a significant common component because unemployment, job-separation, and job-finding rates tend to move together for all income quintiles. Also, for all economic sectors the unemployment rate is lowest for workers in the income quintiles 4 and 5, with the exception of the secondary sector during the 1990s. Similarly, the highest income workers (quintile 5) show the lowest job-separation rates in all sectors and for all time periods. Finally, in all the economic sectors wage volatility is highest for the bottom income quintile, whereas workers of quintiles 3 and 4 have the lowest idiosyncratic wage volatility. The differences in annual wage volatility are in fact quite substantial, with workers in quintile 1 having a wage volatility around 40% to 50% while the workers in quintiles 3 and 4 have values of just 6% to 12%. Curiously, Figures 4.2, 4.3, and 4.4 show that wage volatility increases during both recessions (such as the 1999 and 2009 economic downturns) and expansions (such as the year 2006).

Figure 4.2: Wage volatility, unemployment, separation, and job-finding rates in the primary sector (according to the national income quintile of the workers). Quarterly data 1996:1–2012:4

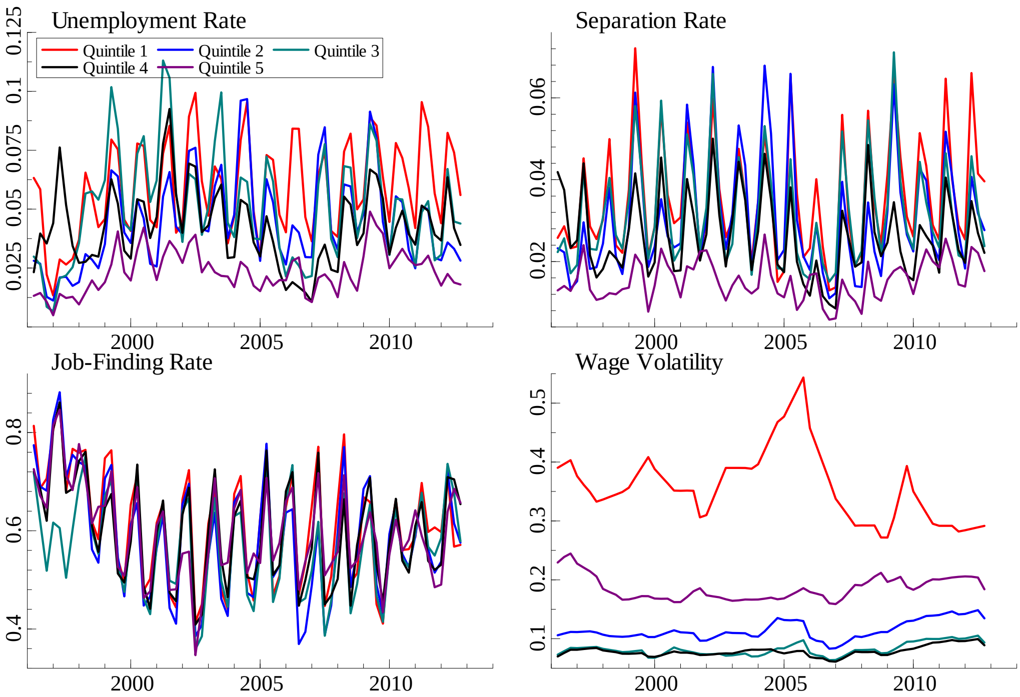


Figure 4.3: Wage volatility, unemployment, separation, and job-finding rates in the secondary sector (according to the national income quintile of the workers). Quarterly data 1996:1–2012:4

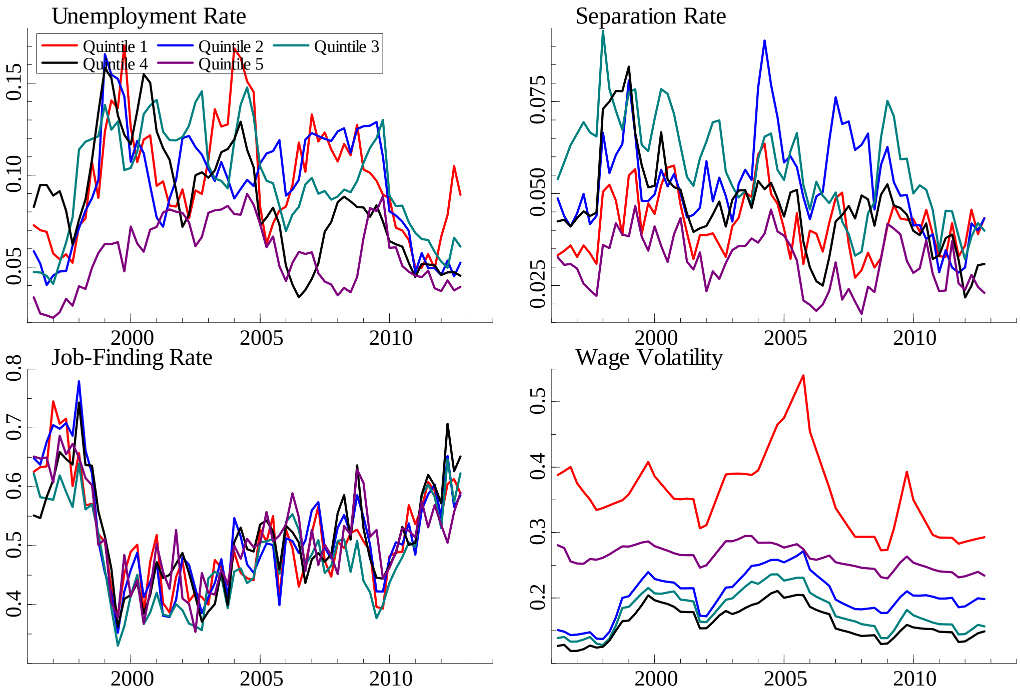
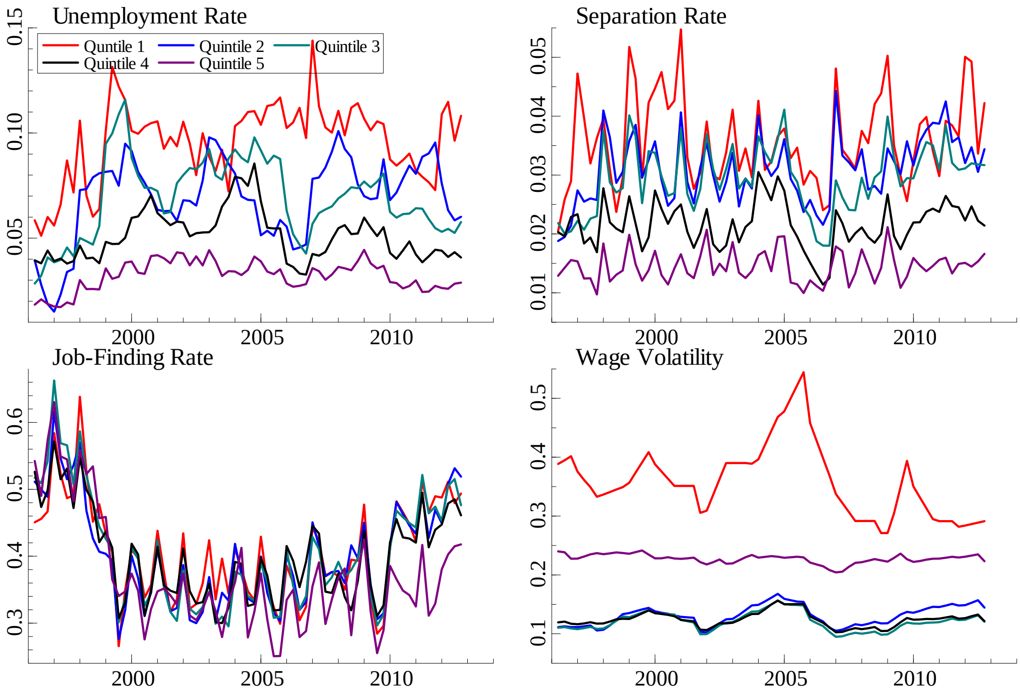


Figure 4.4: Wage volatility, unemployment, separation, and job-finding rates in the tertiary sector (according to the national income quintile of the workers). Quarterly data 1996:1–2012:4



4.3 The FAVAR model

We estimate a FAVAR model for the Chilean economy using macroeconomic data and the labor time series from Madeira (2015) for the quarterly period between 1996:2 and 2012:4. The FAVAR contains three lags and three unknown common factors,[2] and we assume that the only observable factor is the interest rate, as in Bernanke et al. (2005). The following system presents the model:

$$\begin{gathered} \begin{bmatrix} \boldsymbol{F} \\ \boldsymbol{Y} \end{bmatrix}_t = \boldsymbol{\Phi}(L) \begin{bmatrix} \boldsymbol{F} \\ \boldsymbol{Y} \end{bmatrix}_{t-1} + \boldsymbol{\mu}_t \\ \begin{bmatrix} \boldsymbol{X} \\ \boldsymbol{Y} \end{bmatrix}_t = \begin{bmatrix} \boldsymbol{\Lambda}^f & \boldsymbol{\Lambda}^y \\ \boldsymbol{0} & \boldsymbol{I} \end{bmatrix} \begin{bmatrix} \boldsymbol{F} \\ \boldsymbol{Y} \end{bmatrix}_t + \begin{bmatrix} \boldsymbol{e} \\ \boldsymbol{0} \end{bmatrix}_t \end{gathered} \tag{4.1}$$

where $\boldsymbol{F}$ is a 3-dimensional vector containing the unobservable factors, $\boldsymbol{Y}$ contains the interest rate, $\boldsymbol{\Phi}(L)$ is a lag operator of order 3, $\boldsymbol{\Lambda}^f$ is a matrix of parameters of dimension 142×3 indicating how each variable relates to the unobservable factors, and $\boldsymbol{\Lambda}^y$ is a matrix of parameters with dimension 142×1 that shows how the observable variables $\boldsymbol{X}$ relate

[2]The results of the FAVAR estimation are robust to the use of an additional unknown factor.

to the interest rate. Finally, we assume that $\boldsymbol{\mu}_t \sim \mathcal{N}_4(\mathbf{0}, \boldsymbol{\Omega})$ and $\boldsymbol{e}_t \sim \mathcal{N}_{142}(\mathbf{0}, \boldsymbol{\Gamma})$, with $\boldsymbol{\mu}_t$ and $\boldsymbol{e}_t$ being independent.

$\boldsymbol{X}$ is a 142-dimensional vector containing 135 labor series and seven macroeconomics variables. The labor series includes the job-separation rate (EU, the employment to unemployment probability), the job-finding rate (UE, the unemployment to employment probability), and the standard deviation of the total labor earnings (SDTI) of the workers for each of 45 different demographic groups.[3] We classify each group according to age (16–35, 36–54, and 55 or older), economic sector[4] (primary, secondary, and tertiary sectors), and income quintile (with quintile 1 being the lowest income group and 5 the highest income). The seven macroeconomics variables include money stock (M3), consumer price index (CPI), real exchange rate (RER), copper price (CP), and productivity in the primary, secondary. and tertiary sector (PRO1, PRO2, and PRO3, respectively).[5] In addition, we classify all variables in $\boldsymbol{X}$ as slow-moving or fast-moving variables, where the former do not contemporaneously react to the interest rate. We extract the unobservable factors from the group of slow-moving variables in $\boldsymbol{X}$.

We estimate the system of equations using joint likelihood-based Gibbs sampling. That is, we calculate the characterization of the joint posterior density, $P\left(\boldsymbol{\theta}, \boldsymbol{F}^T \left| \boldsymbol{X}^T, \boldsymbol{Y}^T\right.\right)$, by sampling from the conditional densities $P\left(\boldsymbol{F}^T \left|\boldsymbol{\theta}, \boldsymbol{X}^T, \boldsymbol{Y}^T\right.\right)$ and $P\left(\boldsymbol{\theta} \left|\boldsymbol{F}^T, \boldsymbol{X}^T, \boldsymbol{Y}^T\right.\right)$, where a superscript T indicates that the respective vector[6] includes all the sample information from period 1 until period T and $\boldsymbol{\theta} = \left[\boldsymbol{\Lambda}^f, \boldsymbol{\Lambda}^y, \boldsymbol{\Gamma}, \operatorname{vec}(\boldsymbol{\Phi}), \boldsymbol{\Omega}\right]$. We estimated the model by imposing the restrictions $\boldsymbol{\Lambda}^f \boldsymbol{D}^{-1} = \boldsymbol{\Lambda}^f$ and $\boldsymbol{\Lambda}^y + \boldsymbol{\Lambda}^f \boldsymbol{D}^{-1}\boldsymbol{B} = \boldsymbol{\Lambda}^y$ to obtain a unique identification of the factors and their loadings, with $\boldsymbol{D}$ being a non-singular matrix and $\boldsymbol{B}$ a conformable matrix.

To account for the heterogeneity observed in the Chilean labor market, all variables in vector $\boldsymbol{X}$ must be simultaneously analyzed. Given the high dimension of $\boldsymbol{X}$, using a vector of autoregression (VAR) approach to analyze the transmission mechanism of a monetary shock would provide meaningless results because degrees of freedom are lost. VAR systems are normally estimated when the number of variables is relatively small (6–8 variables) and the size of the sample is large; the present case satisfies none of these conditions. Bernanke et al. (2005) suggested that a natural solution to the degrees-of-freedom problem is to use common factors to summarize large amount of information about the economy. Then augmenting a standard VAR by factors is a feasible strategy

[3]We classify the 45 mutually exclusive groups according to the workers' ages, income quintiles and economic sectors. Let $i = 1, \ldots 45$ be expressed by a 3-dimensional vector $z =$\{economic sector (m), age (n), income-quintile (q)\}, with each variable assuming a set of discrete values: $m =$\{1: Primary, 2: Secondary, 3: Tertiary\}, $n =$\{1: 16–35, 2: 36–54, 3: $\geq$55\}, and $q =$\{1, 2, 3, 4, 5\}. Then $i = 1, .., 45$ corresponds to the following mutually exclusive values of matrix z: $1(z = [1, 1, 1])$, $2(z = [1, 1, 2]), \ldots,$ 45 $(z = [3, 3, 5])$.

[4]For further information see footnote 1.

[5]Appendix A describes the series, data sources, and transformations.

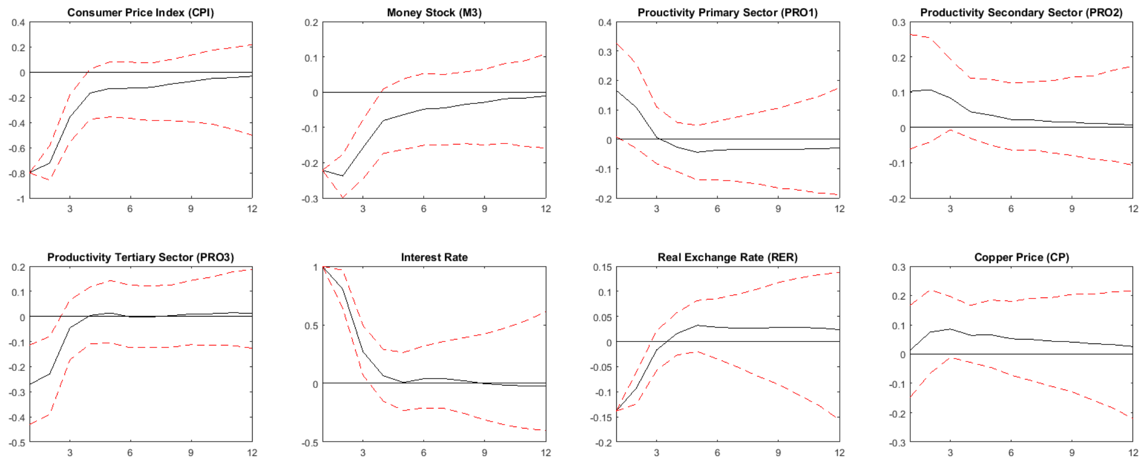
[6]For example, $\boldsymbol{Z}^T = [\boldsymbol{Z}_1, \boldsymbol{Z}_2, \ldots, \boldsymbol{Z}_T]$.

for analyzing large data set where the number of variables is much greater than the size of the sample. In addition, given the heterogeneity, the factors might capture some economic concept related to the labor variables.[7]

4.4 Empirical results

We now present an impulse–response analysis of the transmission mechanism of a contractionary monetary shock. After estimating the FAVAR model, we analyze how a positive standard deviation shock to the interest rate affects the labor productivity (PRO) of each economic sector (primary, secondary, tertiary), the job-finding (UE) and job-separation (EU) rates, the wage volatility (SDTI), and the macroeconomic variables of money aggregate, CPI, and real exchange rate. Table 4.4 reports the summary of the responses, classified according to whether the variable experiences an increase, decrease, or no response after a contractionary monetary shock. Figure 4.5 shows the impulse response function for the macroeconomic variables in vector $\boldsymbol{X}$.

Figure 4.5: Impulse–response function of consumer price index, money stock M3, productivity in primary, secondary and tertiary sectors, real exchange rate, and copper price to one standard deviation positive shock to the interest rate



Note 1: The black graph is the response, in standard deviation units, to one standard deviation positive shock to the interest rate.

Note 2: The dotted red graphs are the 90% confidence interval of the impulse–response function.

Note 3: By construction the impulse–response function of the interest rate starts at one (one standard deviation).

The upper part of Table 4.4 reports the response of the seven macroeconomic variables in vector $\boldsymbol{X}$. The value in brackets is the estimated duration time in quarters of the response to the contractionary monetary shock. In general, all the macroeconomic

[7]The average R^2 of the system describing $\boldsymbol{X}$ in (4.1) is 65.6% and 53.2% for the job-finding rate (EU) and wage volatility (SDTI), respectively. This suggests that two of the three estimated factors are measuring some economic concept related to these labor variables.

Table 4.4: Response summary of macroeconomics variables and labor variables to one standard deviation positive shock to the interest rate

Variable	Increases	Decreases	No response
Macroeconomic variables			
Real money stock (M3)		(4)	
Consumer price index (CPI)		(4)	
Real exchange rate (RER)		(3)	
Real copper price (CP)			(0)
Real productivity primary sector (PRO1)	(1)		
Real productivity secondary sector (PRO2)			(0)
Real productivity tertiary sector (PRO3)		(3)	
Number of responses in labor variables			
Job-separation rate (EU)	20	2	23
Job-finding rate (UE)	10	10	25
Wage volatility (SDTI)	22	11	12

Note 1: The number in $(\cdot)$ is the length of the response, expressed in quarters.

Note 2: There are 45 mutually exclusive groups per labor variable (EU, UE, and SDTI). Each group is classified by economic sector, age, and income quintile. For further information, see Footnote 3.

variables show correctly the expected sign response. For example the consumer price index (CPI), money stock (M3), and real exchange rate (RER) show a negative response lasting on average either three or four quarters to a contractionary monetary shock.[8] The copper price does not exhibit a significant response to an increase in the interest rate, which we expected because the copper price is internationally determined. Productivity in the primary sector (PRO1) reacts positively to an increase in the interest rate, but this response lasts only one quarter, whereas the productivity in the tertiary sector shows a negative response that lasts three quarters. Finally, productivity in the secondary sector does not show a significant response.

Table 4.4 shows that in 20 of the 45 groups, job-separation rate (EU) reacts positively to a contractionary monetary shock and that only two groups exhibit a negative response. Therefore, an unemployment increase typically follows and interest rate increase. Furthermore, the results indicate that in 10 of the 45 groups the job-finding rate reacts negatively to an increase in the interest rate, whereas in the same number of groups the opposite

[8]The economic response lasting three or four quarters is consistent with most of the VAR studies (Christiano et al., 1999). Furthermore, the fact that we do not find a price puzzle is worth emphasizing. Customarily, empirical data analysis finds a positive correlation between inflation and the interest rate, because policy makers typically increase interest rates to counteract periods of increasing inflation. Some studies use sign restrictions in the SVAR to identify shocks according to expected priors, which restrict the information to get rid of price puzzle from the outset. We do not use this approach for three reasons. First, our sample is relatively short, and the identification of shocks using sign restrictions requires long time series for identification. Second, a VAR or SVAR works relatively well when the number of variables is not so big, but our study uses more than 100 labor market time series. Finally, when restrictions are imposed from the outset, it is difficult to discern which results are due to the assumptions made and which are due to the empirical facts (Christiano et al., 1999; Uhlig, 2005).

Table 4.5: Response of labor variables by economic sector to one standard deviation positive shock to the interest rate

Sector	Variable	Number of each type of response		
		Increase	Decrease	No response
Primary	Job-separation rate (EU)	6	1	8
	Job-finding rate (UE)	3	1	11
	Wage volatility (SDTI)	3	8	4
Secondary	Job-separation rate (EU)	10	0	5
	Job-finding rate (UE)	1	5	9
	Wage volatility (SDTI)	10	0	5
Tertiary	Job-separation rate (EU)	4	1	10
	Job-finding rate (UE)	6	4	5
	Wage volatility (SDTI)	9	3	3

Note 1: In each sector there are 45 mutually exclusive groups (15 per labor variable) classified by age and income quintile. For further details, see Footnote 3.

reaction is observed. This seems to imply that recessions not only induce flows into unemployment but also create employment reallocation so that some sectors actually benefit from an increased number of workers looking for vacancies during a recession (Davis et al., 1998; Shimer, 2012; Elsby et al., 2013).

Given that labor intensity differs across economic sectors, how the labor variables respond to monetary shocks may depend on the economic sector analyzed. For example, the service sector is more labor intensive than the other sectors, suggesting that this sector might be more affected. To analyze this statement, Table 4.5 summarizes the responses of each labor market variable (job-separation rate, job-finding rate, and wage volatility) by economic sector.[9]

Table 4.5 shows that the secondary sector is the most affected in the economy when we look at the response of the job-separation rate (EU). In 10 out of 45 groups in the secondary sector, an increase in the separation rate occurs after a contractionary monetary shock, compared to six and four groups in the primary and service sector, respectively. Table 4.4 showed that job-finding rate reacts negatively to an interest rate increase for 10 of the 45 different demographic groups, and Table 4.5 finds that the probability that a worker will be hired after a contractionary monetary shock is lower both in the secondary and tertiary sectors (where five and four demographic groups are affected, respectively).

Finally, Table 4.4 showed that the standard deviation of earnings increases in 22 out of the 45 groups. Most of these groups belong to the secondary sector in the economy, as Table 4.5 shows. This result matches empirical evidence found for the United States, which shows that idiosyncratic income risk increases during recessions (Storesletten et al.,

[9]In Appendix B we show the impulse–response function of the job-separation rate, job-finding rate, and wage volatility for all combinations of age, income quintile, and economic sector.

2001, 2004; McKay and Papp, 2011; Davis and von Wachter, 2011; Guvenen et al., 2014).

The following analysis classifies the response of the labor flow variables and wage volatility (EU, UE, and SDTI) by demographic group and economic sector. Each demographic group is represented by the pair (n, q), where n is an age classification, $n = \{1\text{: } (16\text{–}35),\ 2\text{: } (36\text{–}54),\ 3\text{: } (\geq 55)\}$, and q is the worker's income quintile, $q = \{1,\ 2,\ 3,\ 4,\ 5\}$. For example, (3,5) is the group of workers aged 55 or older whose income belongs to the fifth quintile. Therefore, for each labor flow variable there are 45 mutually exclusive groups.

Table 4.6 shows the employment in each demographic group, presenting the information relative to the employment in both the whole economy and each economic sector. The tertiary sector, with a rate of 63.2%, concentrates most of the employment in the economy, followed by the secondary sector (22.2%) and the primary sector (14.6%). Within each sector, the group of mid-age workers (36–54) has the most employed people (45.8%, 48.9%, and 48.6% in the primary, secondary, and tertiary sector, respectively). Moreover, the group of older workers (55 or older) with income in the first quintile has the lowest rate in each sector (2.2%, 1.3%, and 1.6% in the primary, secondary, and tertiary sector, respectively).

Tables 4.7, 4.8, and 4.9 report the qualitative responses for each economic sector about whether a variable increases, decreases or shows no response after a contractionary monetary shock. Basically, these tables expand Table 4.5 by showing exactly the age group and income quintile of the workers affected by monetary shocks within each economic sector. In addition, these tables provide a basic summary of the qualitative responses of the individual impulse–response graphs shown in Appendix B.

Some results stand out. Within the primary sector, Table 4.7 suggests that after a contractionary monetary shock, the separation rate increases for the oldest workers (55 or older) in the income quintiles 3 and 5 and for the mid-age workers (36–54) in income quintiles 1, 2 and 3. Therefore, the separation rate change is most clear-cut for the oldest workers with income equal to or above that of the middle class and for the mid-age workers with income equal to or below that of the middle class. Then, a contractionary monetary shock is likely to affect negatively the employment of 49.1% of people working in the primary sector (see Table 4.6) Furthermore, after a contractionary monetary shock, the primary sector workers in the age interval 16–35 and in the lowest-income quintile are less likely to find a job. In the primary sector wage volatility also increases for the workers in the lowest-income quintile; this applies for all age groups. However, in the primary sector wage volatility actually declines for older workers (55 or older) and mid-age workers (36–54) across all income ranges.

Within the secondary sector, Table 4.8 shows that almost all groups of workers suffer an increase in the job-separation rate after a contractionary monetary shock. Moreover, mid-age workers (36–54) in the mid-income range (quintiles 2 and 3) and the youngest

Table 4.6: Participation of employment (%) of demographics groups in total employment and employment by sector (average 1996:2–2012:4)

Group (age,quintile)	Sector Primary [864,089.5] 14.6%	Secondary [1,319,939.5] 22.2%	Tertiary [3,750,892.9] 63,2%	Total [5,934,921.9]
(1,1)	8.2	3.7	5.6	5.6
(1,2)	10.9	10.0	8.3	9.0
(1,3)	7.9	10.8	8.5	8.9
(1,4)	3.9	8.4	7.6	7.3
(1,5)	2.7	5.0	6.2	5.4
Subtotal	33.6	37.9	36.2	36.2
(2,1)	5.8	3.9	5.5	5.2
(2,2)	10.4	7.3	7.2	7.7
(2,3)	12.7	11.1	8.8	9.9
(2,4)	8.8	14.2	11.7	11.8
(2,5)	8.1	12.4	15.4	13.6
Subtotal	45.8	48.9	48.6	48.2
(3,1)	2.2	1.3	1.6	1.6
(3,2)	4.3	1.8	1.8	2.2
(3,3)	5.2	2.5	2.3	2.8
(3,4)	4.8	4.0	3.7	3.9
(3,5)	4.1	3.8	5.7	5.0
Subtotal	20.6	13.4	15.1	15.5
Total	100	100	100	100

Note 1: [·] is the average number of employed people.

Note 2: (n, q) is a (age, quintile) group. $n = \{1\text{: (16–35)}, 2\text{: (36–54)}, 3\text{: } (\geq 55)\}$, and q is the worker's income quintile, $q = \{1, 2, 3, 4, 5\}$.

workers (16–35) suffer a decrease in their job-finding rate. Therefore, the secondary sector suffers a double impact of a contractionary monetary shock that increases job destruction (73.4% of people working in the secondary sector are likely to move into unemployment) and decreases job creation (40.5% of people working in the secondary sector are likely to experience a decrease in the probability of finding a job). Also, almost all groups of workers in the secondary sector experience an increase in their wage volatility after a contractionary shock.

Table 4.9 shows that in the tertiary sector only the oldest workers (55 and older) suffer an increase in their separation rate after a contractionary monetary shock (this represents 11.4% of the people working in the tertiary sector), whereas the youngest workers (16–35) are the only ones that experience a decrease in their job-finding rate; this group represents 30% of the people working in the tertiary sector. These groups also see an increase in wage volatility after a contractionary shock. Therefore, the results for the secondary and tertiary sectors show that idiosyncratic wage risk is counter-cyclical in Chile just as for the United States (Storesletten et al., 2001, 2004; McKay and Papp, 2011).

Thus, after a contractionary monetary shock, the FAVAR analysis shows that job-separation rate will increase relatively more in the secondary sector, job-finding rate will drop, and people between the ages of 16 and 35 are the most affected in the secondary and tertiary sectors. These results suggest that in periods where contractionary monetary policy is at work, other economic policies (e.g., fiscal policy) should stimulate employment in the secondary sector and, particularly, employment of the youth in both the tertiary and secondary sector. Given that the youth group normally consists of unskilled and inexperienced people and that an increase in the interest rate is likely to decrease the probability of finding a job for this group, an economic policy that increases the human capital of the youth (e.g., job training, higher education) and stimulates the hiring of people between 16 and 35 is likely to counteract the negative effects of a contractionary monetary shock.

Table 4.7: Response of labor variables in the primary sector by age and income quintile to o one standard deviation positive shock to the interest rate

	Job-separation rate (EU) response
Increase	(1,2), (2,1), (2,2), (2,3), (3,3), (3,5)
Decrease	(3,4)
No response	(1,1), (1,3), (1,4), (1,5), (2,4), (2,5), (3,1), (3,2)
	Job-finding rate (UE) response
Increase	(1,4), (2,3), (2,4)
Decrease	(1,1)
No response	(1,2), (1,3), (1,5), (2,1), (2,2), (2,5), (3,1), (3,2), (3,3), (3,4), (3,5)
	Wage volatility (SDTI) response
Increase	(1,1), (2,1), (3,1)
Decrease	(1,2), (2,2), (2,4), (2,5), (3,2), (3,3), (3,4), (3,5)
No response	(1,3), (1,4), (1,5), (2,3)

Note 1: (n, q) is a (age, quintile) group. $n = \{1\text{: } (16–35), 2\text{: } (36–54), 3\text{: } (\geq 55\ )\}$, and q is the worker's income quintile, $q = \{1, 2, 3, 4, 5\}\}$.

Table 4.8: Response of labor variables in the secondary sector by age and income quintile

	Job-separation rate (EU) response
Increase	(1,3), (1,4), (1,5), (2,1), (2,3), (2,4), (2,5), (3,1), (3,3), (3,5)
Decrease	-
No response	(1,1), (1,2), (2,2), (3,2), (3,4)
	Job-finding rate (UE) response
Increase	(3,2)
Decrease	(1,1), (1,2), (1,4), (2,2), (2,3)
No response	(1,3), (1,5), (2,1), (2,4), (2,5), (3,1), (3,3), (3,4), (3,5)
	Wage volatility (SDTI) response
Increase	(1,1), (1,2), (1,3), (1,4), (1,5), (2,1), (2,2), (2,3), (2,4), (3,1)
Decrease	-
No response	(2,5), (3,2), (3,3), (3,4), (3,5)

Note 1: (n, q) is a (age, quintile) group. $n = \{1\text{: } (16\text{–}35), 2\text{: } (36\text{–}54), 3\text{: } (\geq 55)\}$, and q is the worker's income quintile, $q = \{1, 2, 3, 4, 5\}$.

Table 4.9: Response of labor variables in the service sector by age and income quintile

	Job-separation rate (EU) response
Increase	(3,1), (3,2), (3,3), (3,5)
Decrease	(1,4)
No response	(1,1), (1,2), (1,3), (1,5), (2,1), (2,2), (2,3), (2,4), (2,5), (3,4)
	Job-finding rate (UE) response
Increase	(2,4), (3,1), (3,2), (3,3), (3,4), (3,5)
Decrease	(1,1), (1,2), (1,3), (1,4)
No response	(1,5), (2,1), (2,2), (2,3), (2,5)
	Wage volatility (SDTI) response
Increase	(1,1), (1,2), (1,3), (1,4), (2,1), (3,1), (3,2), (3,3), (3,4)
Decrease	(2,2), (2,3), (2,5)
No response	(1,5), (2,4), (3,5)

Note 1: (n, q) is a (age, quintile) group. $n = \{1\text{: } (16\text{–}35), 2\text{: } (36\text{–}54), 3\text{: } (\geq 55)\}$, and q is the worker's income quintile, $q = \{1, 2, 3, 4, 5\}$.

4.5 Conclusion

We analyze how monetary policy affects labor markets with a special focus on the heterogeneity of different economic sectors and demographic groups. We find that there is indeed heterogeneity of how different economic sectors react to monetary shocks. While fluctuations in real labor productivity growth are strongly correlated across different sectors, clear differences are evident in terms of the behavior of employment flows. Labor productivity growth in each economic sector has a low correlation with business cycle fluctuations in the unemployment rate and flows into and out of unemployment. We also show that fluctuations in unemployment rates have a high correlation with changes in job-separation rates. This seems to support the empirical evidence found for other countries that job destruction plays a crucial role in explaining cyclical unemployment fluctuations (Elsby et al., 2013).

After a contractionary monetary shock, the secondary sector reacts most strongly in terms of both an increased job-separation rate and a decreased job-finding rate. Also, real labor productivity falls in the tertiary sector after an interest rate shock, in contrast to the primary and secondary sectors. For the primary and tertiary sectors, it is mostly the older workers (55 or older) who experience an increase in the job-separation rate, while in the secondary sector the impact on job destruction is felt across all ages and income levels.

Finally, we find that the idiosyncratic volatility of labor earnings increases in both the secondary and tertiary sectors after a contractionary monetary shock, confirming similar results found for the United States (Storesletten et al., 2004; McKay and Papp, 2011). This increase in idiosyncratic earnings risk is also found for the lowest-income workers in the primary sector.

A Data

In Appendix A and B, the index $i = 1,..,45$ corresponds to 45 mutually exclusive groups classified according to the workers' ages, income quintiles, and economic sectors. Let i be expressed by a 3-dimensional vector z ={economic sector (m), age (n), income-quintile (q)}, with each variable assuming respectively a set of discrete values: m ={1: Primary, 2: Secondary, 3: Tertiary}, n ={1: 16–35, 2: 36–54, 3: ≥55}, and q ={1, 2, 3, 4, 5}. Then $i = 1,..,45$ corresponds to the following mutually exclusive values of matrix z : $1(z = [1, 1, 1])$, $2(z = [1, 1, 2])$, . . . , 45 $(z = [3, 3, 5])$.

Table A.1: Data source, description and transformation

Variable	Description	Source	Transformation	Slow code
M3	Real money stock M3	Central Bank of Chile	log difference	0
CPI	Consumer Price Index	Central Bank of Chile	difference of growth rate	1
I1	1-year real interest rate	Central Bank of Chile	-	0
PRO1	Real productivity primary sector	Our own calculation	log difference	1
PRO2	Real productivity secondary sector	Our own calculation	log difference	1
PRO3	Real productivity service sector	Our own calculation	log difference	1
CP	Real copper price	Central Bank of Chile	log difference	1
RER	Real exchange rate	Central Bank of Chile	log difference	0
EUi	Job-separation rate	Madeira (2015)	-	1
UEi	Job-finding rate	Madeira (2015)	-	1
SDTIi	Mean Standard Deviation of the idiosyncratic annual change in total labor income	Madeira (2015)	difference	1

Note 1: All variables are seasonally adjusted using the Census X13 program.

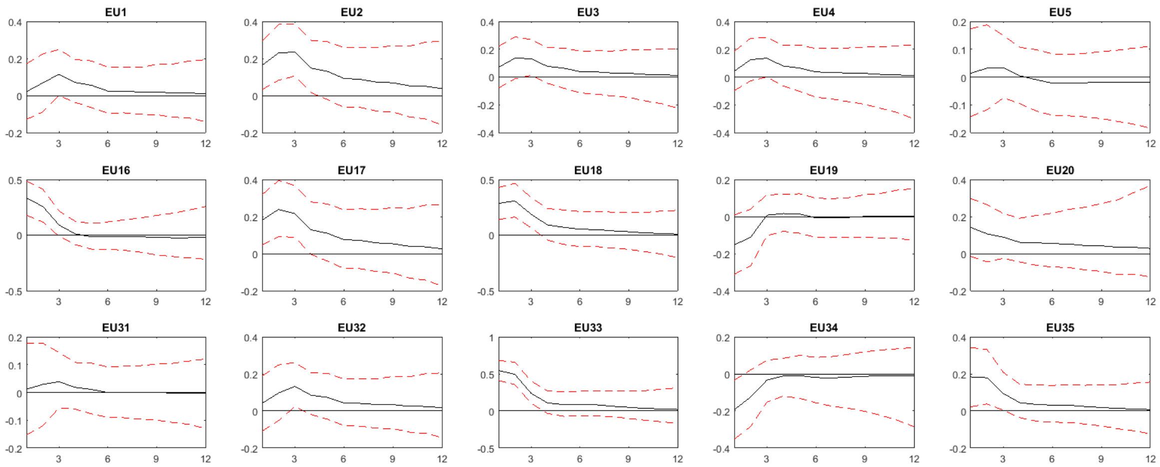
Note 2: Slow code: 1 stands for a slow-moving variable and 0 for a fast-moving variable. A slow moving variable does not contemporaneously react to the interest rate.

Note 3: The 1-year interest rate is the nominal average weighted interest rate of the financial system for operations of 90 days to 1 year, deflated by the inflation rate based on the log difference of the CPI.

Note 4: Real Productivity is obtained by the ratio of the total aggregate value-added of each economic sector (published by the Central Bank of Chile) divided by the number of workers in each sector as given by the quarterly National Employment Survey (ENE) calculated by Madeira (2015).

B Impulse–response functions

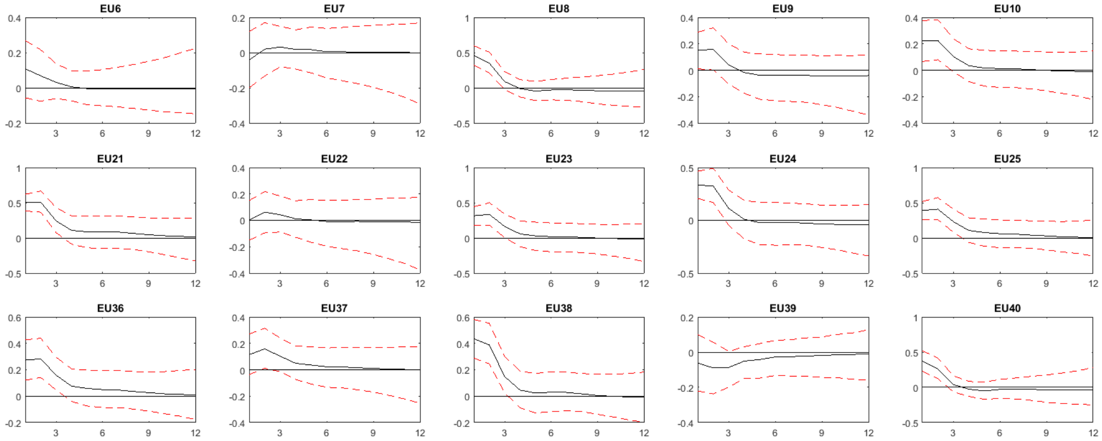
Figure A.1: Impulse–response of the job-separation rate (EUi) to a positive shock to the interest rate (one standard deviation) in the primary sector



Note 1: The black graph is the response, in standard deviation units, to one standard deviation positive shock to the interest rate.

Note 2: The dotted red graphs are the 90% confidence interval of the impulse–response function.

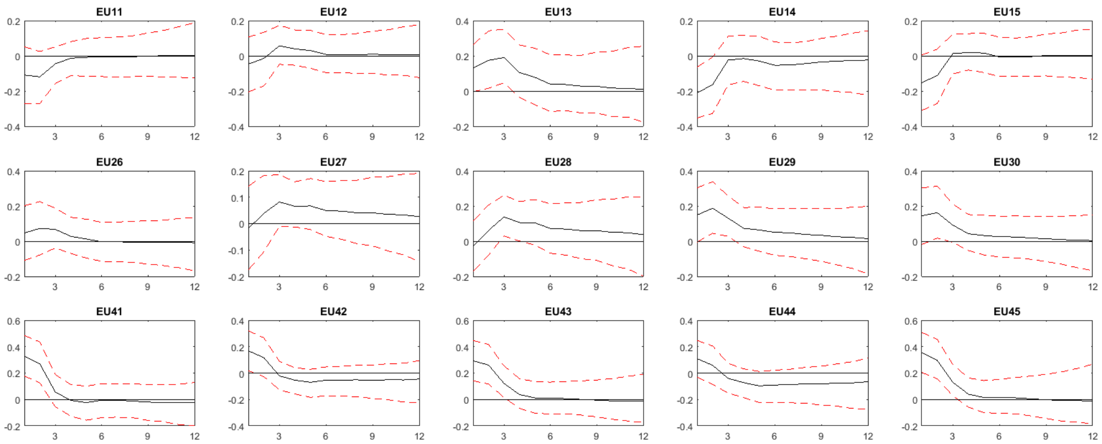
Figure A.2: Impulse–response of the job-separation rate (EUi) to a positive shock to the interest rate (one standard deviation) in the secondary sector



Note 1: The black graph is the response, in standard deviation units, to one standard deviation positive shock to the interest rate.

Note 2: The dotted red graphs are the 90% confidence interval of the impulse–response function.

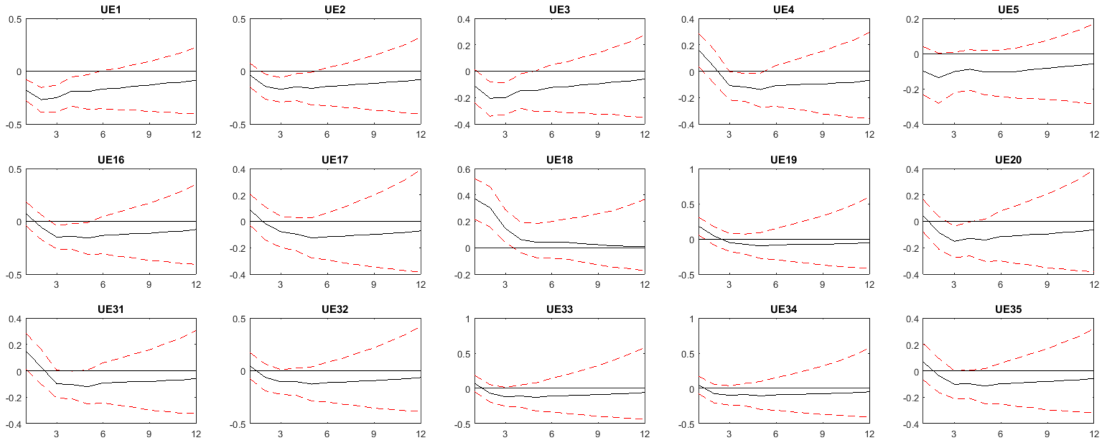
Figure A.3: Impulse–response of the job-separation rate (EUi) to a positive shock to the interest rate (one standard deviation) in the tertiary sector



Note 1: The black graph is the response, in standard deviation units, to one standard deviation positive shock to the interest rate.

Note 2: The dotted red graphs are the 90% confidence interval of the impulse–response function.

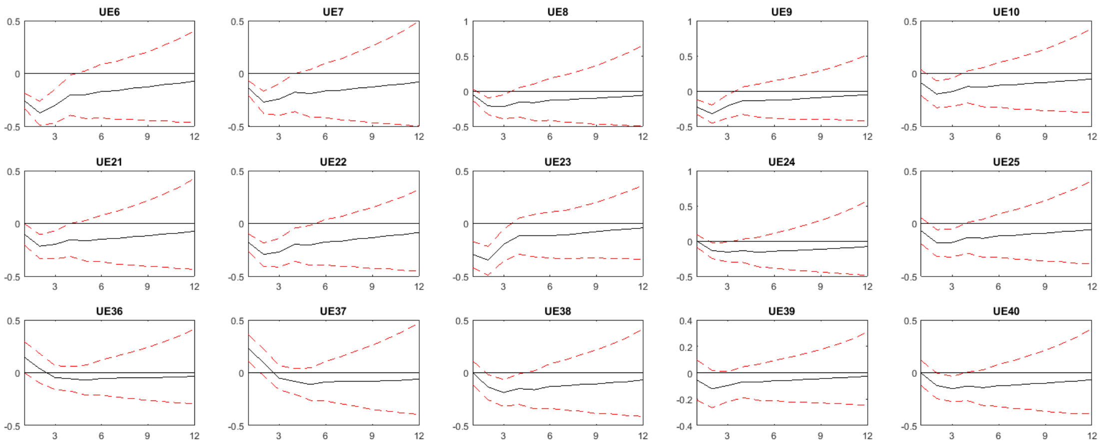
Figure A.4: Impulse–response of job-finding rate (UEi) to a positive shock to the interest rate (one standard deviation) in the primary sector



Note 1: The black graph is the response, in standard deviation units, to one standard deviation positive shock to the interest rate.

Note 2: The dotted red graphs are the 90% confidence interval of the impulse–response function.

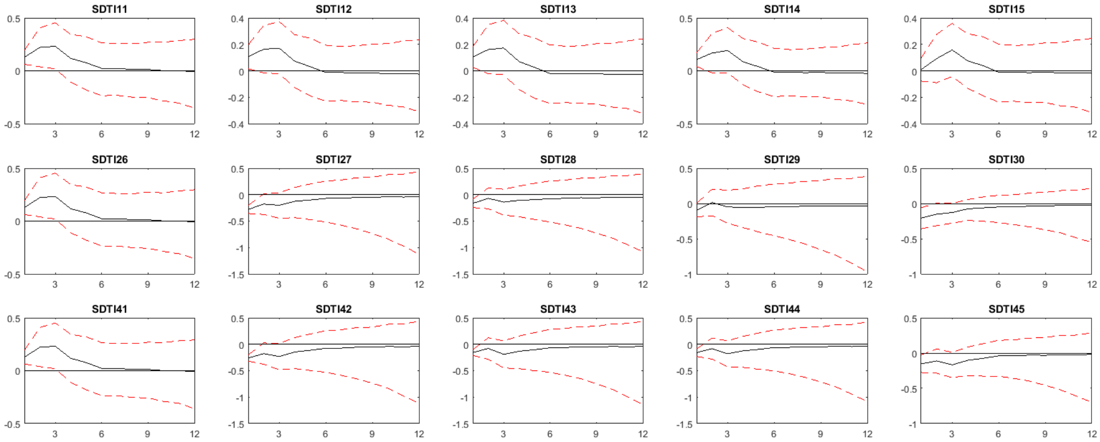
Figure A.5: Impulse–response of job-finding rate (UEi) to a positive shock to the interest rate (one standard deviation) in the secondary sector



Note 1: The black graph is the response, in standard deviation units, to one standard deviation positive shock to the interest rate.

Note 2: The dotted red graphs are the 90% confidence interval of the impulse–response function.

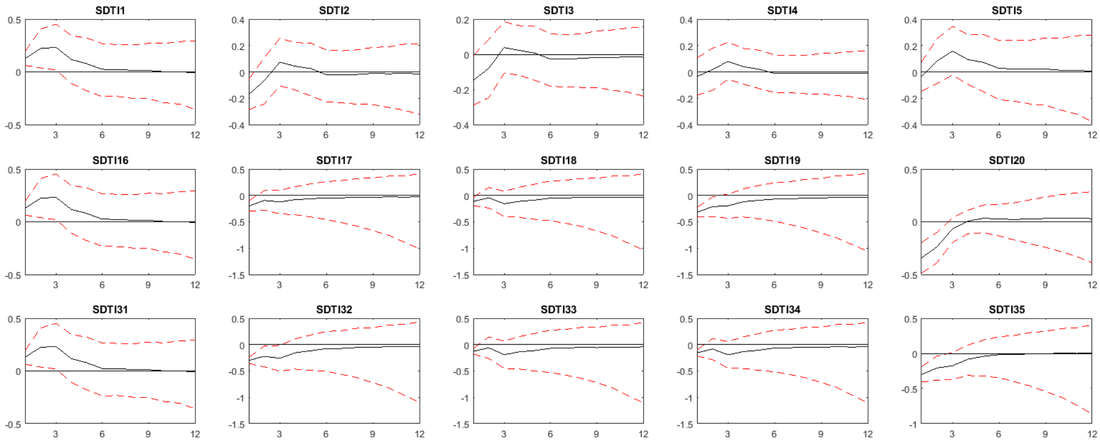
Figure A.6: Impulse–response of job-finding rate (UEi) to a positive shock to the interest rate (one standard deviation) in the tertiary sector



Note 1: The black graph is the response, in standard deviation units, to one standard deviation positive shock to the interest rate.

Note 2: The dotted red graphs are the 90% confidence interval of the impulse–response function.

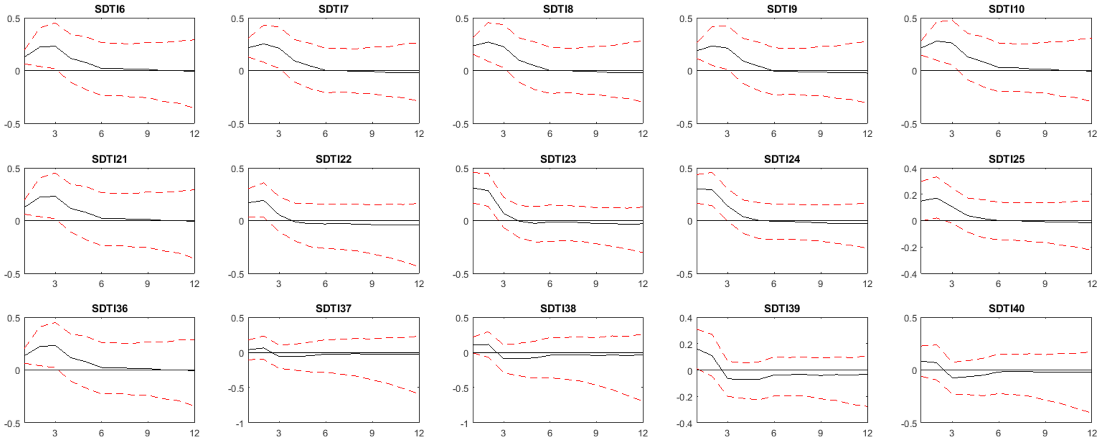
Figure A.7: Impulse–response of the wage volatility (SDTIi) to a positive shock to the interest rate (one standard deviation) in the primary sector



Note 1: The black graph is the response, in standard deviation units, to one standard deviation positive shock to the interest rate.

Note 2: The dotted red graphs are the 90% confidence interval of the impulse–response function.

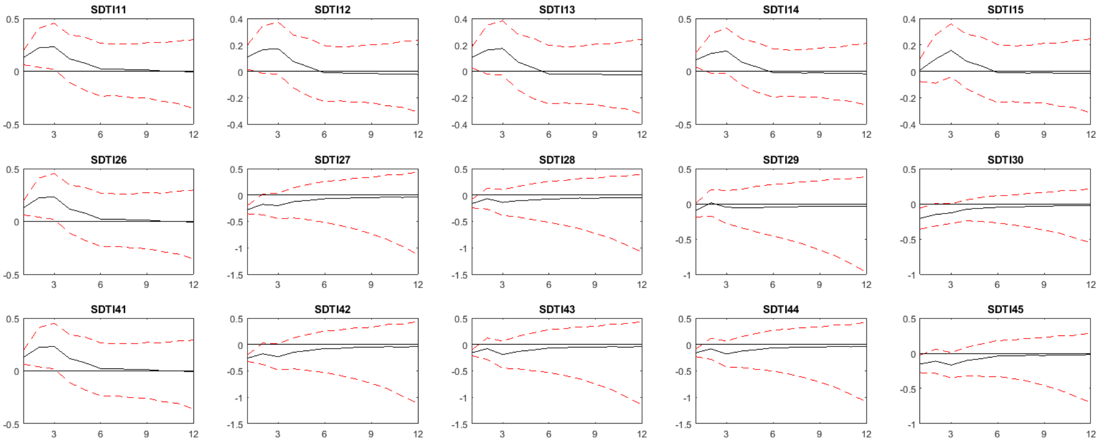
Figure A.8: Impulse–response of the wage volatility (SDTIi) to a positive shock to the interest rate (one standard deviation) in the secondary sector



Note 1: The black graph is the response, in standard deviation units, to one standard deviation positive shock to the interest rate.

Note 2: The dotted red graphs are the 90% confidence interval of the impulse–response function.

Figure A.9: Impulse–response of the wage volatility (SDTIi) to a positive shock to the interest rate (one standard deviation) in the tertiary sector



Note 1: The black graph is the response, in standard deviation units, to one standard deviation positive shock to the interest rate.

Note 2: The dotted red graphs are the 90% confidence interval of the impulse–response function.

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