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Crises, Labor Market Policy, and Unemployment

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Crises, Labor Market Policy, and Unemployment*

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Abstract

Using a sample of 97 countries spanning the period 1980–2008, we estimate that financial crises have a large negative impact on unemployment in the short term, but that this effect rapidly disappears in the medium term in countries with flexible labor market institutions, whereas the impact of financial crises is less pronounced but more persistent in countries with more rigid labor market institutions. These effects are even larger for youth unemployment in the short term and long-term unemployment in the medium term. Conversely, large upfront, or gradual but significant, comprehensive labor market policies have a positive impact on unemployment, albeit only in the medium term.

JEL Classification Numbers: E29, J60, E32, D7.

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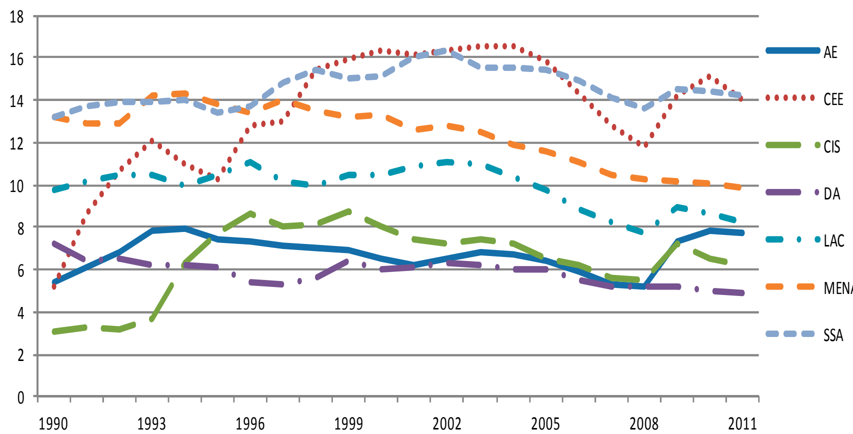
I. INTRODUCTION

Although researchers have documented the impact of real shocks on overall unemployment in OECD countries (Bruno and Sachs, 1985; Nickell, 1997; Blanchard and Wolfers, 2000; Nickell et al., 2005), very little attention has been paid to the impact of financial crises on unemployment. Yet, like preceding crises (Reinhart and Rogoff, 2009), the 2008 financial crisis has resulted in a significant, and hitherto persistent, increase in unemployment in advanced economies, as illustrated in Figure 1. While many emerging market countries have generally weathered the crisis well, youth unemployment has increased (or stopped declining) at least temporarily, in several regions, including Latin America, the Middle East, and North Africa.

Figure 1. The evolution of unemployment across the regions



Note: AE= Advanced Economies; CEE= Central and Eastern Europe; CIS= Commonwealth of Independent States; DA=Dynamic Asia; LAC= Latina America; MENA= Middle East and North Africa; SSA= Sub-Sahara Africa;

Financial crises contribute to an increase in unemployment mainly through the decline in output and investment associated with heightened uncertainty, higher risk premia (Pindyck, 1991; Pindyck and Solimano, 1993), and tighter lending standards (Hall, 2009). Hysteresis effects related to the loss of attractiveness of the unemployed can also lead to an increase in long-term and structural unemployment (Ball, 2009). More vulnerable groups such as the youth and women with limited professional experience also become increasingly at risk, with their participation rate typically declining (Duval et al., 2011).

This paper systematically measures the impact of financial crises on general, long-term, and youth unemployment in 140 countries. While the graph in Figure 1 is suggestive, we analyze formally the magnitude and persistence of the increase in unemployment resulting from financial crises. We also examine the extent to which the labor institutional and regulatory framework modulates the response of unemployment to financial crises. Following the work of Blanchard and Wolfers (2000), Nickell et al. (2005), and Bassanini and Duval (2009) on the role of institutions in explaining the unemployment response to macroeconomic or unobserved shocks, we look at the direct impact of labor market institutions on unemployment, as well as how the impact of crises varies depending on labor market institutions. We find that the flexibility of labor markets directly affects not only the magnitude but also the persistence of the impact of financial crises on unemployment.

We test further the impact of labor market institutions on unemployment by estimating the impact of large-scale or gradual changes in labor market institutions on unemployment. The endogeneity of labor market reforms is a potentially important issue in estimating their impact on unemployment. We attempt to address this issue by using a couple of methods, which uncover some interesting facts. In particular, we find that change in labor market institutions are less likely to occur in more centralized political regimes that also are of duration, and that large-scale changes in labor market institutions policies can reduce unemployment by the same amount that financial crises increased it, albeit only after several years. Overall, the results suggest that labor market policies may play an important role in reducing unemployment over the medium term. These policies, however, have to be properly designed to also improve the quality of employment and to minimize possible negative short-term effects, not investigated here, on inequality and job destruction.

The rest of the paper is organized as follows. The next section describes the dataset and present key descriptive statistics. Section III analyzes the impact of financial crises on unemployment. Section IV assesses the effect of labor institutions in shaping the response of unemployment to financial crises. Section V analyzes the effect of labor market policies on unemployment outcomes. Section VI concludes.

II. Data and Descriptive Statistics

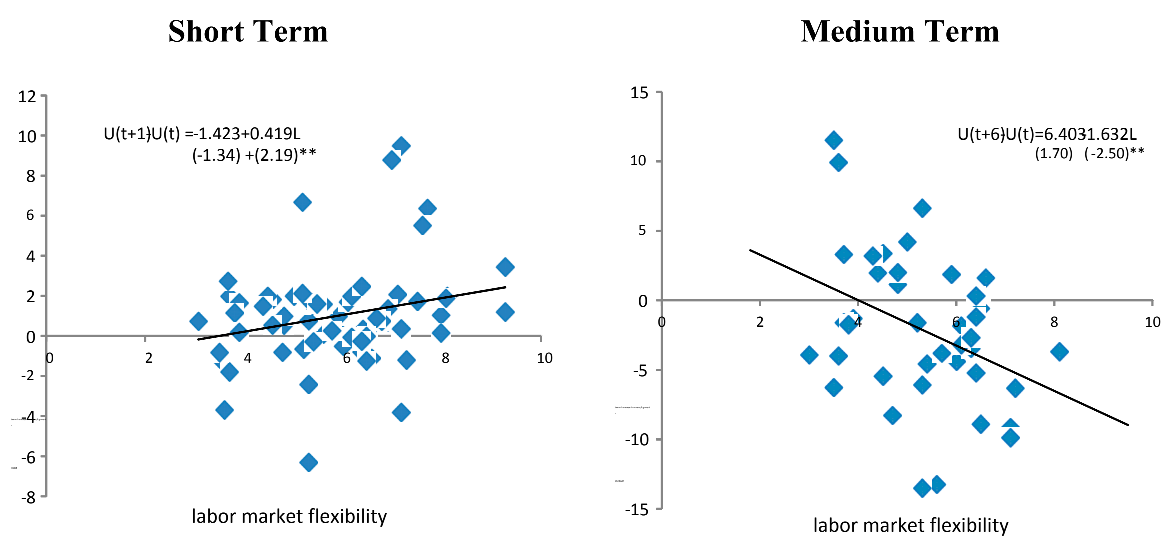
Our data set covers a panel of 97 countries from 1980 to 2008. Data for labor market flexibility are taken from the Fraser Institute's Economic Freedom of the World (EFW) database (Gwartney and Lawson, 2010), which provides a composite measure of labor market flexibility and indicators of labor market flexibility on six policy areas: (i) minimum wage (M), (ii) hiring and firing regulation (H), (iii) centralized collective wage bargaining (C), (iv) mandated cost of hiring (MCH), (v) mandated cost of work dismissal (MCW), and (vi) conscription (CO). All indicators are standardized on a 0–10 scale, with higher values of the indicator representing a more flexible labor market.

The sources of the data for the other variables used in the empirical analysis are the IMF's *World Economic Outlook* (WEO), the World Bank's World Development Indicators (WDI), the Penn World Table version 7.0 by Heston et al., (2011), the database constructed by Laeven and Valencia (2010) on financial crisis occurrences (for banking, currency, and debt crises), and the database on political institutions by Keefer (2010). The full list of variables, definitions, and sources is provided in the Annex.

Table 1 presents descriptive statistics for the labor market flexibility indicators and the unemployment outcomes analyzed in the paper. For the composite labor market flexibility indicator we have a total of 1,214 observations, ranging from a minimum of 1.8 to a maximum of 9.5. Among the unemployment outcomes, we can notice that unemployment is mostly concentrated among young people (aged between 15 and 24). Table 2 shows that the correlation between unemployment outcomes and labor market flexibility. Indicators are in most of the cases negative and statistically significant, with hiring costs and youth unemployment having the strongest negative correlation.[3]

Figure 2 sheds some light on the role played by labor market flexibility in shaping the response of unemployment to shocks. The first panel shows the increase in unemployment rate immediately following a financial crisis and the labor market indicator. As expected, economies that are characterized by a more flexible labor market have higher increase in unemployment in the short term, reflecting higher elasticity of unemployment to output. The second panel shows the cumulative changes in unemployment following a financial crisis, and the composite indicator of labor market flexibility. The negative relation between the medium-term effect of financial crises on unemployment and labor market flexibility suggests that labor market resilience to financial crises is a positive function of labor market flexibility. Thus, while unemployment tends to increase more after a financial crisis in the short term in countries with a more flexible labor market, the medium-term effect tends to be higher for countries characterized by a more rigid labor market. This hypothesis will be formally tested in the next section.

Figure 2. Increase in Unemployment Following a Crisis vs. Labor Market Flexibility



Note: ** denote significance at 5 percent.

[3] See Bernal-Verdugo et al., (2011) for empirical evidence on the effect of labor market flexibility on different unemployment outcomes.

III. The Impact of Financial Crises on Unemployment

A. Methodology

The dynamic impact of financial crises on unemployment is estimated following the approach proposed by Jorda (2005) and Teulings and Zubanov (2010), which allows the impulse response functions to be estimated directly. For each year *k* following the onset of the downturn, the estimation equation has the following form:

$$U_{i,t+k} - U_{i,t} = \alpha_i^k + \sum_{j=1}^{2} \gamma_j^k \, \Delta U_{i,t-j} + \beta_k D_{i,t} + \delta' \boldsymbol{X}_{i,t} + \varepsilon_{i,t}^k \qquad (1)$$

where $U_{i,t+k}$ is the unemployment rate in country *i* in period $t+k$, $D_{i,t}$ is a downturn dummy that takes value 1 for the start of a financial crisis episode in period *t* in country *i* and zero otherwise, α_i represents country fixed effects that capture unobserved country-specific determinants of unemployment, $\boldsymbol{X}_{i,t}$ is a vector of control variables including the initial level of unemployment, the initial annual change of the share of urban population, the initial annual change of government spending as share of GDP, and a deterministic time trend. The coefficient γ_j captures the persistence in changes in unemployment and β_k measures the impact of the crisis on the change in unemployment for each future period *k.* Impulse Response functions (IRFs) are then obtained by least-squares dummy variable of β_k for k = 0, 1,…,6. Potential reverse causality is addressed by estimating changes in unemployment in the years that *follow* a financial crisis.[4]

An alternative way of estimating the dynamic impact of financial crises on unemployment would be to estimate an Auto-Regressive Distributed Lag (ARDL) equation for the unemployment rate and crisis dummies, and to derive the IRFs from the estimated coefficients (Romer and Romer, 1989, 2010; Cerra and Saxena, 2008). However, the IRFs derived using an ARDL specification tend to be sensitive to the choice of the number of lags, and as a result tend to be unstable. In addition, long-lasting effects of shocks may be unduly found, reflecting the use of what Cai and Den Haan (2009) call one-type-of-shock models. By contrast, the approach used in the present paper does not suffer from these problems, because lagged changes in unemployment enter only as control variables and are not used to derive the IRFs, and because the structure of the equation does not impose permanent effects. The confidence bands associated with the estimated IRFs are also easily computed using the standard deviations of the estimated coefficients β_k rather than requiring Monte Carlo simulations.

[4] See Furceri and Zdzienicka (2010, 2011), and Duval et al. (2011) for a similar approach.

B. Results

Figure 3 shows a statistically significant and long-lasting increase in the unemployment rate following the occurrence of a financial crisis, by increasing the rate of unemployment by about 1 percent at the peak—three years after the occurrence of the crisis—and by about 0.5 percent over the medium term—six years thereafter. Our findings are consistent with the evidence reported for OECD countries for which Furceri and Mourougane (2009) report an increase in unemployment of about 0.5–1 percent associated with financial crises, and Guichard and Rusticelli (2010) an increase of structural unemployment of about ¾ percentage point.

Financial crises have a slightly different effect on youth unemployment, with the higher increase at the peak of 2 percent, only one year after the occurrence of the crisis, and a statistically insignificant effect over the medium term. This could be due to shorter time-series available for youth unemployment (compared to the overall unemployment rate), or to a particularly high medium-term negative effect of downturns on labor force participation for young workers, implying a lower medium- term effect of crises on youth unemployment than on overall unemployment (Duval et al., 2011).

By contrast, the effect of financial crises on long-term unemployment is extremely persistent but becomes statistically significant only two years after the occurrence of a crisis. Over the medium term (six years after the crisis), long-term unemployment stabilizes at a level 6 percentage points higher than the pre-crisis level, suggesting the existence of hysteresis effects.

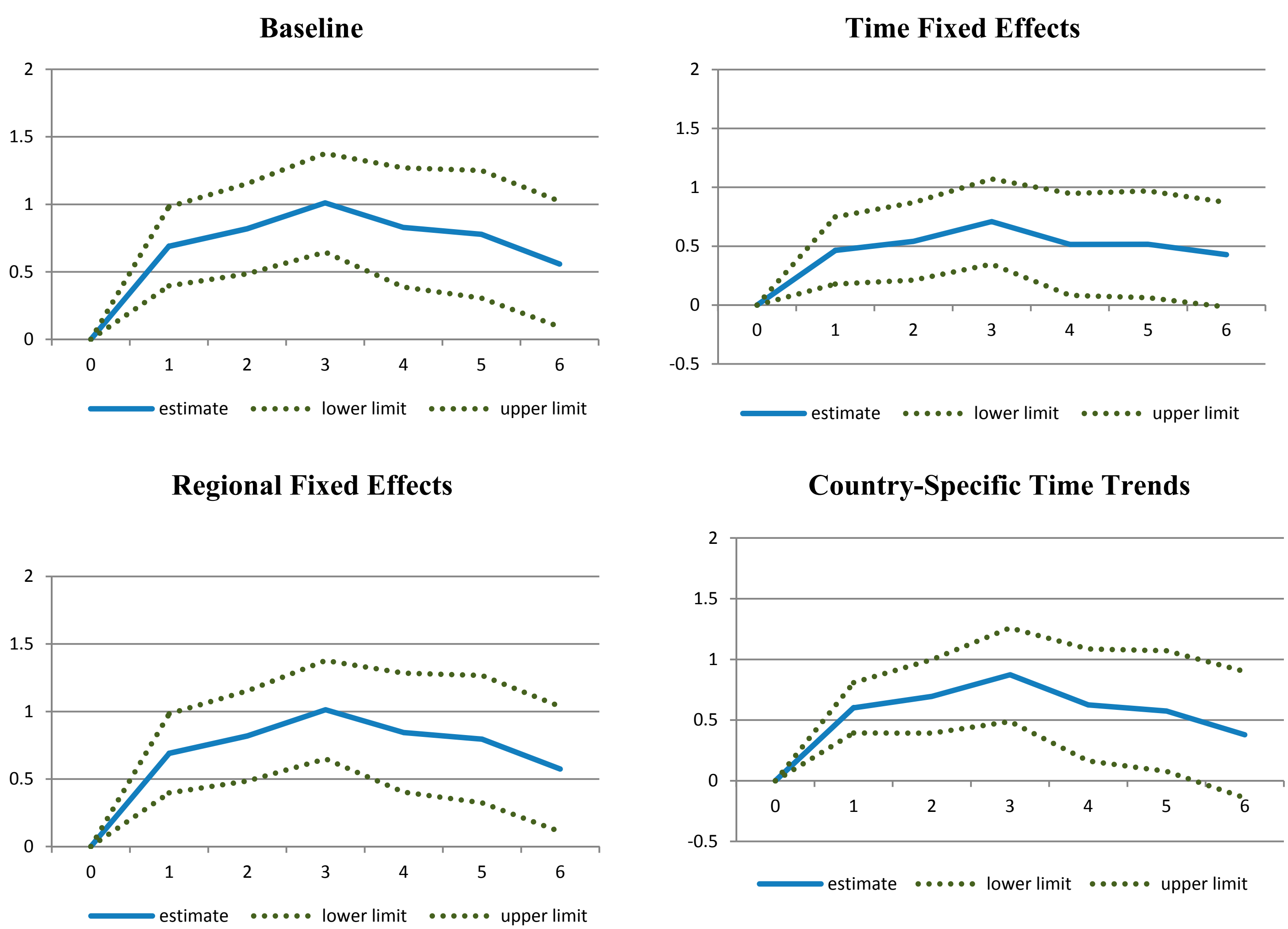
Figure 3. The Effect of Financial Crises on Unemployment Outcomes
(In percentage points)

Note: solid line represents the estimated IRF; dotted lines represent 95 percent confidence bands.

C. Robustness Tests

To check the robustness of the results, equation (1) was re-estimated by alternatively including (i) time fixed effects, (ii) regional fixed effects, and (iii) a country-specific time trend. As shown in Figure 4, the results using these different controls remain statistically significant and broadly unchanged. As an additional robustness check, we have also restricted the estimation sample to those countries for which data for change in unemployment are available for each period *k* to control for a possible composition bias derived from estimating equation (1) over an unbalanced set of countries. The results, not reported, remain statistically significant and qualitatively unchanged.

Figure 4. The Effect of Financial Crises on Unemployment Outcomes: Robustness Check (In percentage points)



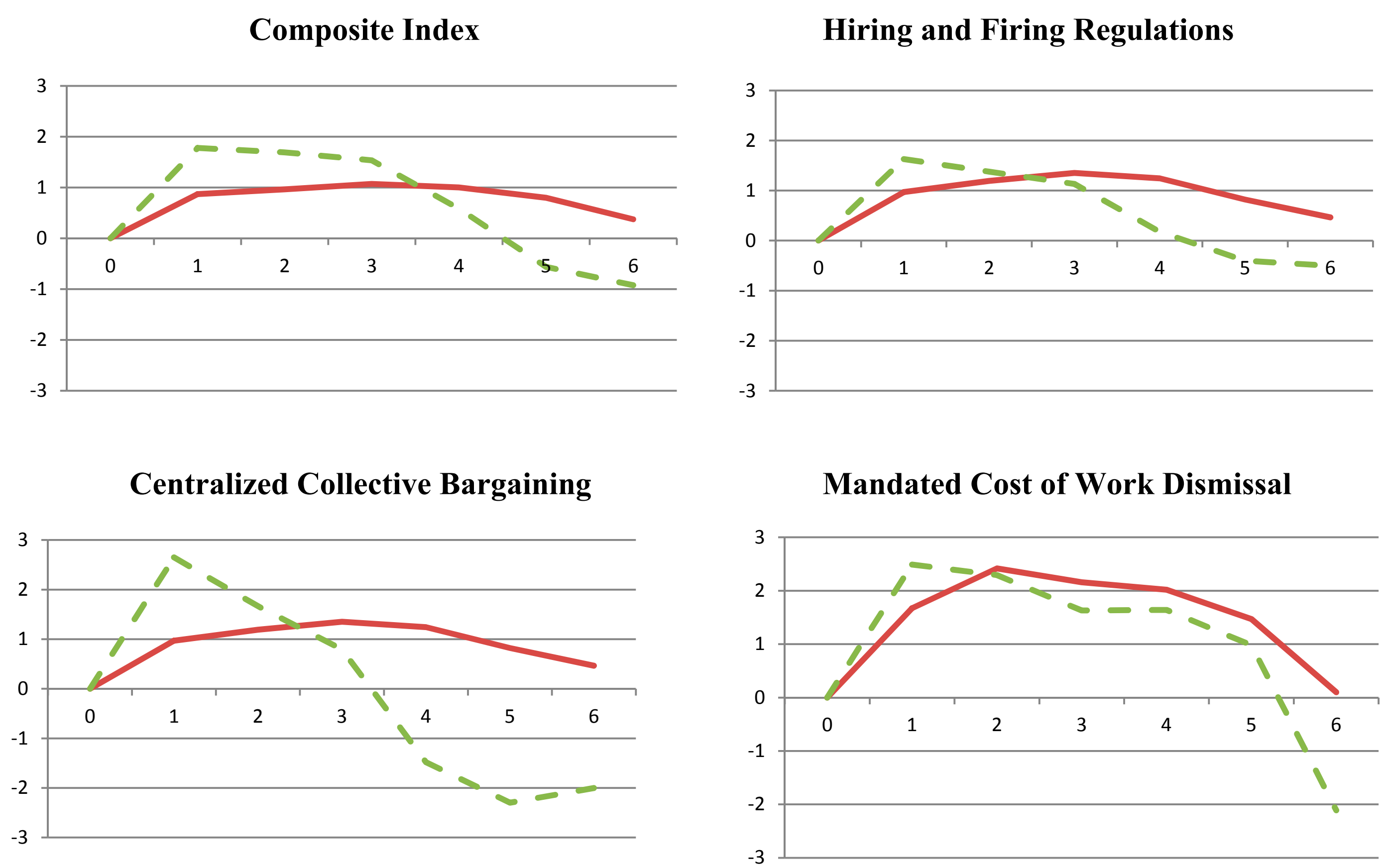
IV. Labor Market Flexibility and the Response of Unemployment to Financial Crises

To test the impact of labor market flexibility in shaping the effect of financial crises on unemployment, equation (1) was augmented to include labor market flexibility indicators (*L*) as a control and as an interaction term with the crisis dummy:

$$U_{i,t+k} - U_{i,t} = \alpha_i^k + \sum_{j=1}^{l} \gamma_j^k \, \Delta U_{i,t-j} + \beta_k D_{i,t} + \theta \overline{L_i} + \vartheta_k D_{i,t}\left(L_{i,t} - \overline{L_i}\right) + \delta' \boldsymbol{X}_{i,t} + \varepsilon_{i,t}^k \quad (2)$$

Figure 5 presents the response of unemployment to financial crises, obtained considering the first (solid line) and third quartile (dotted line) of the distribution for different labor market flexibility indicators. The figure shows that the short-term impact of crises on unemployment is higher in countries with more flexible labor markets, while the medium-term effect is larger in countries with more rigid labor markets. This difference in the response is statistically significant both in the short and in the medium term (Table 3). Specifically, an increase of one point in the labor market flexibility composite indicator increases the short-term effect of crises on unemployment by 0.4 percentage point, but reduces the medium term impact by about 0.6 percentage point. The results for the composite flexibility indicator are qualitatively similar for youth and long-term unemployment. Among the sub-indicators, hiring and firing regulations and centralized collective bargaining are the indicators with the largest medium term impact. As an additional robustness check, we have also restricted the estimation sample to those countries for which data for change in unemployment are available for each period *k* to control for a possible composition bias derived from estimating equation (3) over an unbalanced set of countries. The results, not reported, remain qualitatively unchanged.

Figure 5. The Role of Labor Market Flexibility in Shaping the Effect of Financial Crises on Unemployment
(In percentage points)



Note: —— rigid labor market; - - - - flexible labor market. These correspond, respectively, to the 25th and 75th percentile of the distribution of the relevant indicator.

V. LABOR MARKET POLICIES AND UNEMPLOYMENT OUTCOMES

A. The Impact of Large-Scale Changes in Labor Market Institutions

We first estimate the impact of large-scale changes in labor market institutions by assuming that a large change takes place when, for a given country in a given time, the annual change in the composite labor market flexibility indicator exceeds by two standard deviations the average annual change over all observations. Because continuous annual data for our labor market flexibility indicators are not available for the majority of the countries in the sample before 2000, the identification of major policy change has been carried out during the period 2000–08. The results, according to our identification strategy, suggest that during the last decade several countries have implemented major reforms of their labor market institutions and regulations.[5]

To estimate the dynamic impact of labor market policies on unemployment we use a specification similar to equation (1):

$$U_{i,t+k} - U_{i,t} = \alpha_i^k + \sum_{j=1}^{l} \gamma_j^k \Delta U_{i,t-j} + \beta_k R_{i,t} + \delta' \boldsymbol{X}_{i,t} + \varepsilon_{i,t}^k \tag{3}$$

where $R_{i,t}$ is a dummy that takes value 1 for the start of a large-scale change in policy in period t in country i and zero otherwise, $\boldsymbol{X}_{i,t}$ is a vector of control variables including the initial level of unemployment, the crisis dummy used in the previous section, and a deterministic time trend. β_k measures the impact of the reform on the change in unemployment for each future period k.

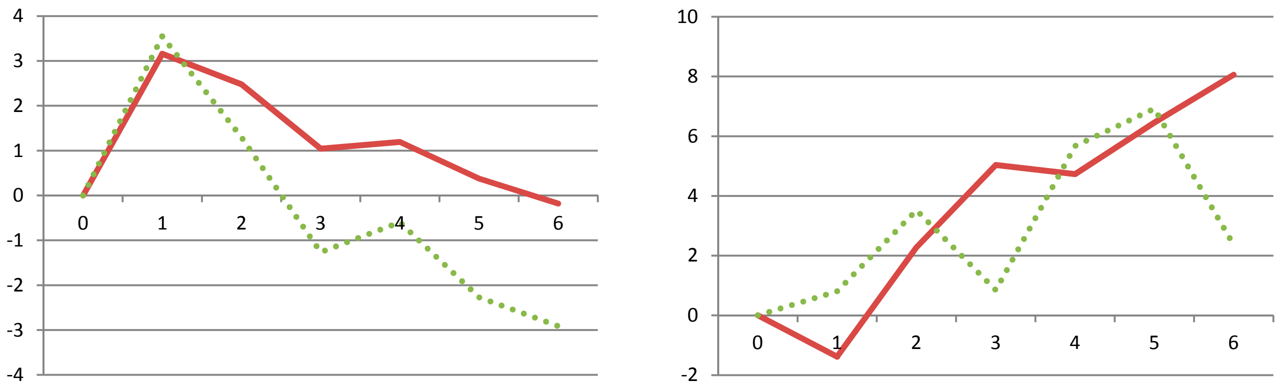
Figure 6 plots the IRFs obtained by using the least-squares dummy variable of β_k for k = 0, 1, …, 6. The figure shows that it takes, on average, about six years before the effects of labor market reforms on unemployment materialize, but the effects over the medium term are statistically significant and sizeable (Table 4).

In particular, we find that labor market policies are associated with a decrease in unemployment of about ¾ percent over the medium term, which is similar in absolute terms to the increase in unemployment associated with financial crises (Figures 7–8). In other words, large-scale changes in labor market policies have the opposite effect of a financial crisis.

[5] In particular, we have identified 48 episodes of large-scale labor market reforms mostly occurring during the past decade. The full list of episodes is available from the authors upon request.

Figure 6. The Role of Hiring and Firing Regulation in Shaping the Effect of Financial Crises on Youth and Long-Term Unemployment
(In percentage points)

Youth Unemployment **Long-term Unemployment**



Note: —— rigid labor market; - - - - flexible labor market; these correspond, respectively, to the 25th and 75th percentile of the distribution of the relevant indicator.

Figure 7. The effects of Reforms on Unemployment-OLS

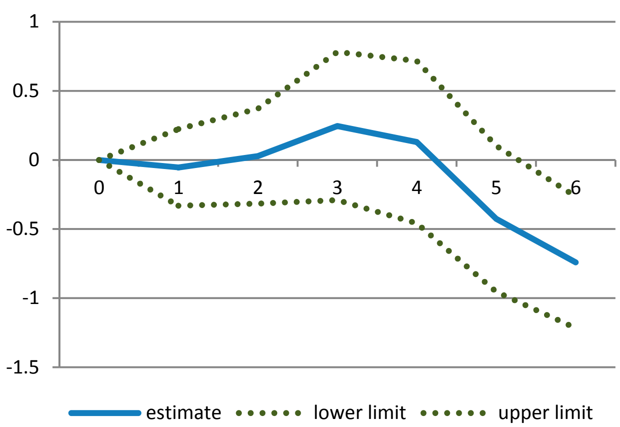
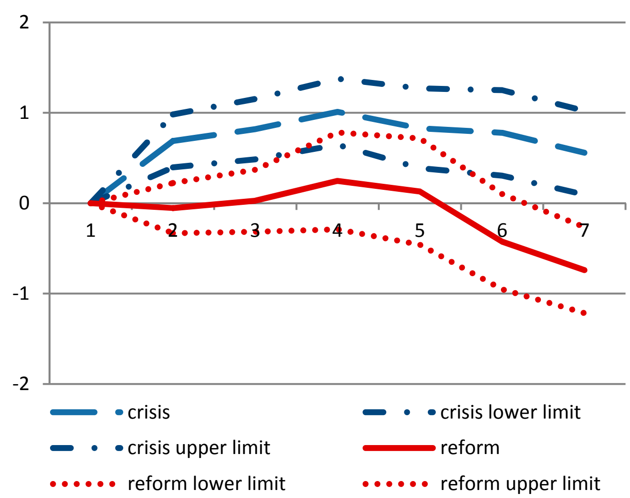


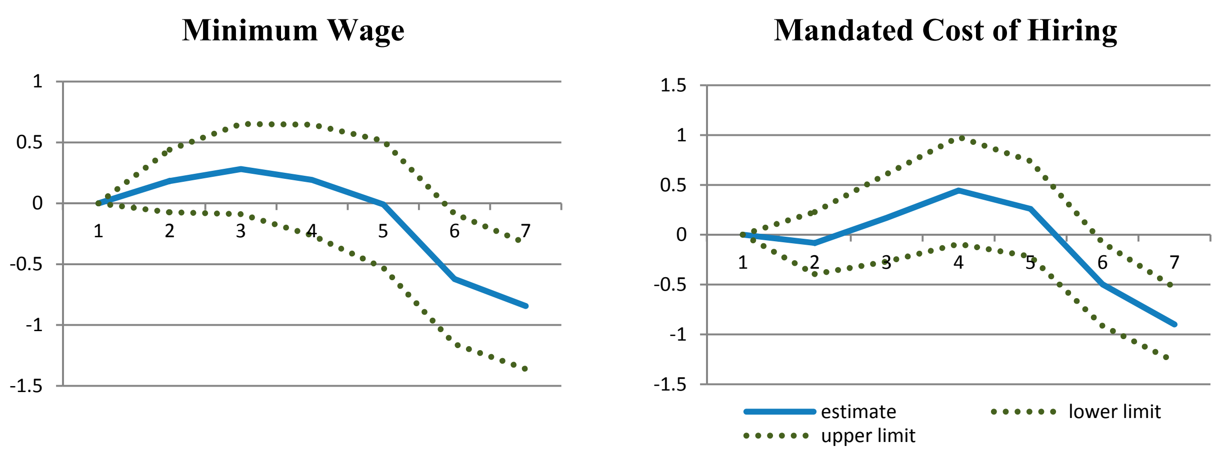
Figure 8. The effects of Reforms vs. Crises on unemployment-OLS



Note: Solid line represents the estimated IRF; dotted lines represent 95 percent confidence bands.

Regarding youth and long-term unemployment, the medium-term estimates for the large-scale changes associated with the composite flexibility indicator (Table 6) are statistically significant for youth unemployment but insignificant for long-term unemployment. Among the sub-indicators, we find that changes in minimum wages and mandated cost of hiring have a statistically significant impact on youth unemployment (Figure 9), while changes in hiring and firing regulation and conscription have a statistically significant medium-term impact on long-term unemployment. Overall, the results seem to suggest that labor market policies aimed at lowering hiring costs and reducing rigidity in hiring and firing regulation may have large and significant impact on unemployment outcomes.

Figure 9. The Effects of Labor Market Policies on Unemployment-OLS
(In percentage points)



Note: Solid line represents the estimated IRF; dotted lines represent 95 percent confidence bands.

B. Endogeneity

Estimates of the impact of labor market policies could be biased because of endogeneity of the reforms. In particular, while potential reverse causality is addressed by estimating changes in unemployment in the years that *follow* a large change in labor market institutions, it could be still the case that unobserved factors influencing the dynamics of unemployment over time could affect the probability of a change in labor market institutions. In addition, given that a worsening in economic conditions (i.e., an increase in unemployment) is likely to increase the probability of change in labor market policies, and assuming that β_k is negative, the OLS estimates of equation (3) are likely to be biased towards zero and to underestimate the effect of labor market policy on unemployment.

To address the endogeneity issue, we propose two strategies, which are both based on a Probit estimation of determinants of labor market policy. The first strategy consists of re-estimating equation (3) by instrumenting the dummy R for those variables that we have found to be significantly correlated with the probability of changes in labor market institutions. The second approach consists of decomposing the effect of labor market policies on unemployment in two parts: the expected effect owing to all the variables which are affecting the probability of a change in labor market institutions, and the residual effect which can be thought of as the effect of the labor market policy itself:

$$R_{i,t} = R_{i,t}^{\mathrm{E}} + \omega_{i,t} \tag{4}$$

where $R_{i,`t}^{\mathrm{E}}$ represents the predicted probability of changes in labor market institutions, and $\omega_{i,t}$ is the error term of the Probit model. In particular, this decomposition distinguishes between the effect of an increase in the probability of a change in labor market institutions and the effect resulting from the change itself. As we estimate the probability of reform using the Probit model, this strategy is similar but not identical to directly adding to equation (3) all the variables used to predict the reform.

Determinants of labor market policies

The literature on the determinants of reforms has suggested that the probability of reforms depends on several macroeconomic (initial labor market and output growth conditions, economic size, trade openness, exchange rate regime, fiscal conditions) and political variables (ideology of the executive, the degree of decentralization and fractionalization in the political power, political stability, election cycles) (Alesina, 1987; Alesina and Drazen, 1991; Curkierman and Tommasi, 1998; Saint-Paul, 2004; and Duval, 2008).

Table 5 shows our estimation results for several Random Effect Probit models. In our baseline specification (column 1), we find that the pre-existing level of the labor market institutions index plays an important role in determining the probability of implementing a change in labor market institutions. Specifically, we find that the higher the quality of the existing labor market institutions, the less likely a country is to implement such a reform. Presumably, an economy with already flexible labor markets would not be in need of implementing further changes in labor market institutions, and would therefore be less likely to go through such an episode.

A favorable economic situation (as measured by the GDP growth gap with respect to a five-year moving average, *gap_growth*) is also found to decrease the probability of a change in policy occurring, although to a lesser extent. Intuitively, a country would display less appetite for changes in labor market institutions (in the presence of eventual political costs) when the economy's growth rate is above average. By contrast, other macroeconomic and demographic factors were not found to have a statistically significant effect. Interestingly, lagged unemployment rates do not seem to affect the probability of the implementation of large-scale changes, which implicitly suggests that reverse causality from unemployment to labor market policies may not be an issue.

Among the political variables included in our estimations, we find that an increase in the degree of decentralization (e.g. going from a presidential system to one in which the president is elected by the assembly) is the variable that plays the most important role in increasing the probability of changes in labor market institutions. This is in line with the findings of Dabrowski and Gortat (2002) that the more politically liberal countries implemented relatively fast market reforms, while the more authoritative and centralized regimes tended to preserve economic institutions that favored the former communist oligarchy. Our findings contrast, however, with the results of Alesina et al., (2006) who show that strong executives (in unified governments and presidential systems) are more likely to swiftly implement fiscal and inflation stabilization programs than their less centralized counterparts.

We also find empirical evidence that the length of time that the chief executive's party has been in power has a negative effect on the probability of changes in labor market institutions. This variable can be interpreted as a measure of the dominance or preponderance that the official party has in the domestic politics of a country. As argued in the political science literature, dominant parties tend to promote reforms that protect their dominance, and to avoid the ones that tarnish it. For instance, Southall (1997) argues that dominant parties favor policies that increase their popularity at the expense of other, deeper, economic and industrial reforms. In light of electoral and other political costs associated with reforms that promote labor market flexibility, it would be natural for a dominant party to postpone such reforms, as evidenced by our results.

Other political variables which feature prominently in the empirical literature—such as election cycles, political ideology, government fractionalization, measures of political stability, and the presence of a constitutional limit on the number of years the executive can serve before new elections—have been tested but proved to be statistically insignificant.

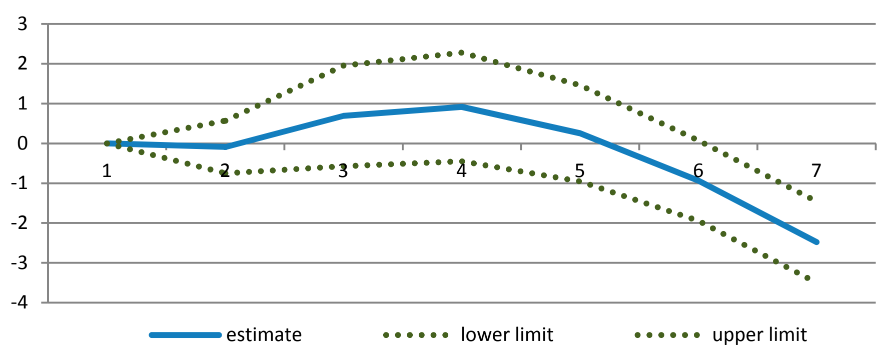
Results

First, we instrument labor market policies in equation (3) with the variables found to be significantly correlated with the probability of changes in labor market institutions, namely (i) the pre-existing level of the labor market institution index, (ii) the degree of decentralization in the executive power, (iii) the length of time the chief executive's party has been in power, and (iv) a measure of the output gap. The results obtained using this approach are reported in Figure 9, and confirm the significant effect of labor market policies on unemployment over the medium term (the IRFs are statistically significant after five years).

The results also suggest a larger effect than that shown for the baseline, which is consistent with the fact that OLS estimates of β_k may be biased towards zero. In particular, the IV estimates suggest that labor market policies may lead to a reduction in unemployment of about 2.5 percentage points over the medium term. This result is robust to different specifications and estimates of the probability of changes in labor market institutions (Table 6).

In our second approach, we decompose the medium-term effect of *expected* and *non-expected* change in labor market institutions as in equation (4). Table 7 shows that the anticipated effect of changes in policy is considerably larger than the unpredicted component. However, the difference between the effect of expected and non-expected changes in labor market institutions is not statistically significant in most of the cases. Interestingly, the coefficient associated with the unanticipated effect is statistically significant in all specifications, and its magnitude is close to that obtained for the baseline regression.

Figure 10. The Effects of Labor Market Policies on Unemployment-IV
(In percentage points)



Note: Solid line represents the estimated IRF; dotted lines represent 95 percent confidence bands.

C. Gradual Changes in Labor Market Institutions

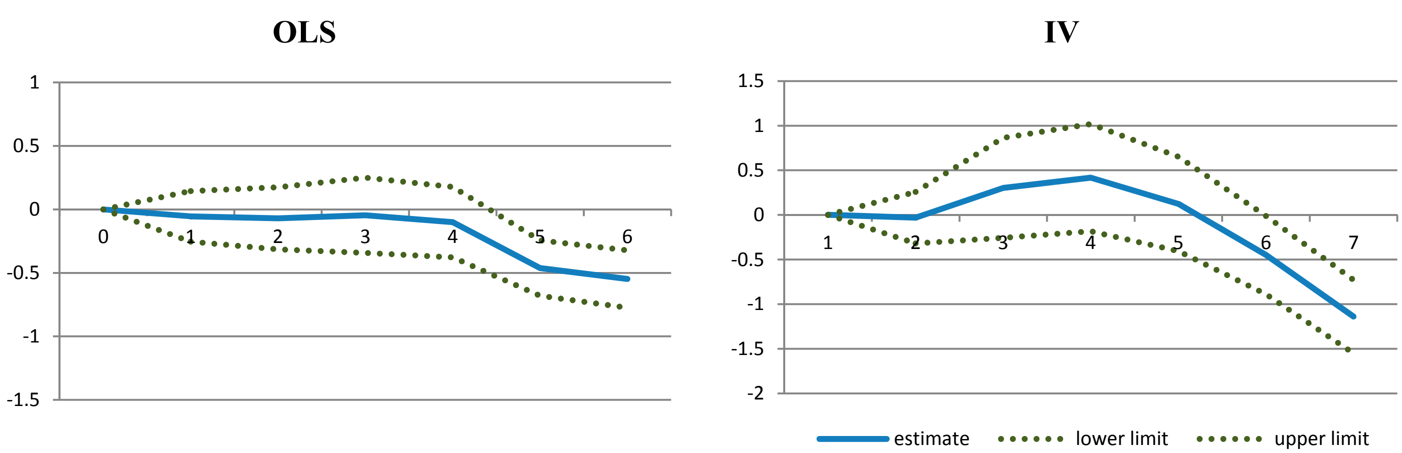
Although the use of a large threshold has the advantage of limiting measurement error in the identification of substantial policy changes, and of reducing the possibility of reverse causality, it may also prevent identification of those change in labor market institutions that are spread over a longer period. In practice, however, gradual reforms are much more difficult to identify; they also tend to occur less frequently (Duval, 2008). As noted by Duval (2008), "certain reforms start small before getting big over the years and therefore cannot be associated with any specific year. One example is the slow but quasi continuous decline in tax wedges in Denmark between the late 1980s and the mid-1990s or in Finland since the mid-1990s."

To address this issue, and to check the robustness of our results, we re-estimate equation (3), replacing the dummy variable associated with the occurrence of large-scale changes ($R_{i,t}$) with the change in the labor market flexibility indicator ($\Delta L_{i,t}$):

$$U_{i,t+k} - U_{i,t} = \alpha_i^k + \sum_{j=1}^{l} \gamma_j^k \, \Delta U_{i,t-j} + \beta_k \Delta L_{i,t} + \delta' \boldsymbol{X}_{i,t} + \varepsilon_{i,t}^k \tag{5}$$

By estimating equation (5) we can quantify the effect of gradual changes in labor market flexibility on unemployment. The first panel of Figure 11 presents the IRFs obtained when equation (5) is estimated using OLS. The results are in line with those obtained in Figure 6, and suggest that an increase of one point in the labor market flexibility indicator decreases unemployment by about 0.5 percent over the medium term. The effect tends to be larger when changes in the labor market flexibility composite indicator are instrumented for those variables that we have found to be significantly correlated with the probability of reforms. In particular, the IRFs reported in the second panel of Figure 11 suggest that an increase of one point in the labor market flexibility composite indicator decreases unemployment by about 1.2 percent over the medium term.

Figure 11. The Effects of Gradual Labor Market Policies on Unemployment
(In percentage points)



Note: Solid line represents the estimated IRF; dotted lines represent 95 percent confidence bands.

VI. Conclusions

Using a novel panel data set for a large number of countries, this paper shows that the impact of financial crises on unemployment largely depends on the flexibility of labor market institutions. Impulse responses show that in countries with more flexible labor markets, the impact of financial crises is sharper but short-lived. Conversely, in countries with more rigid labor markets, the effect of financial crises appears to be initially more subdued, but highly persistent. The effects are more pronounced for youth unemployment in the short term, perhaps underscoring their higher vulnerability as well as declining labor market participation in the medium term. The impact on long-term unemployment was found to be very pronounced in the medium term, highlighting potential hysteresis effects of financial crises on unemployment.

The paper also provides some evidence that large upfront, or gradual but significant comprehensive labor market policies may reduce unemployment, albeit only in the medium term. The positive impact of labor market policies is particularly pronounced for the young. The adoption of such labor market policies is less likely in more centralized political regimes that are long-lasting, and when institutions are already more flexible and economic conditions relatively good.

The results, however, raise the issue of the design and possible sequence of such reforms. Employment protection should be designed in such a way as to internalize social costs and not inhibit job creation and labor reallocation, to also improve the quality of employment and to minimize possible negative short-term effects, not investigated in this paper, on inequality and job destruction. Micro- and macro-studies on OECD countries over the past decade show that it is important to protect workers, rather than jobs, by coupling unemployment benefits with pressure on the unemployed to take jobs and measures to help them (Blanchard, 2006). Moreover, artificial restrictions on individual employment contracts should also be avoided.

Table 1. Summary Statistics for Labor Market Outcomes and Flexibility Indicators

	Obs.	Mean	Std. Dev.	Min	Max
Labor market outcomes					
Unemployment	2826	8.9	5.9	0.0	37.3
Long-term unemployment	984	33.8	18.3	0.5	84.9
Youth unemployment	1669	17.6	10.5	0.7	70.9
Labor market flexibility					
Composite index	1214	5.9	1.5	1.8	9.5
Minimum wage	1135	6.2	2.7	0.0	10.0
Hiring and firing regulations	1056	4.7	1.5	1.0	8.8
Centralized collective bargaining	1124	6.4	1.5	1.8	9.5
Mandated cost of hiring	1166	6.9	2.0	1.9	10.0
Mandated cost of worker dismissal	927	5.8	3.1	0.0	10.0
Conscription	1656	5.9	4.3	0.0	10.0

Source: Fraser Institute's Economic Freedom of the World (EFW) database.

Table 2. Correlation Matrix of Labor Market Outcomes and Flexibility Indicators

	U	YU	LU	L	M	H	C	MCH	MCW	CO
U	1									
YU	0.51***	1								
LU	0.90***	0.56***	1							
L	-0.11***	-0.22***	-0.13***	1						
M	-0.22***	-0.19***	-0.15***	0.64***	1					
H	-0.21***	-0.24***	-0.21***	0.44***	0.26***	1				
C	0.01	-0.03	0.00	0.46***	0.17***	0.54***	1			
MCH	-0.02	-0.31***	-0.02	0.63***	0.43***	0.28**	0.30***	1		
MCW	0.01	0.02	-0.04	0.47***	0.13***	0.19**	0.061	-0.02	1	
CO	0.01	-0.11**	-0.04	0.70***	0.20***	-0.09***	0.03	0.32***	0.06*	1

Note: U=unemployment; YU= youth unemployment; LU=long-term unemployment; L=composite labor market flexibility index; M= minimum wage; H=hiring and firing regulation; C=centralized collective bargaining; MCH=mandated cost of hiring; MCW=mandated cost of work dismissal; CO= conscription.
*, **,*** denote significance at 10percent, 5 percent, and 1 percent, respectively.

Table 3. Short- and Medium-Term Effects of Financial Crises on Unemployment: Flexible vs. Rigid Labor Markets

	L	M	H	C	MCH	MCW	CO
				Initial effect			
Crisis	1.314	1.392	1.315	1.834	1.503	2.116	1.254
	(4.43)***	(4.86)***	(4.33)***	(3.49)***	(4.47)***	(6.04)	(3.36)***
Crisis*Flexibility	0.395	0.078	0.300	0.300	0.160	0.149	-0.028
indicator	(2.26)**	(0.69)	(1.60)*	(1.60)*	(1.22)	(1.60)*	(-0.46)
				Medium term effect			
Crisis	-0.260	0.019	-0.043	-0.805	-0.394	-1.110	0.179
	(-0.50)	(0.02)	(-0.10)	(-1.47)	(-0.61)	(-3.12)***	(0.32)
Crisis*Flexibility	-0.564	-0.201	-0.441	-0.441	-0.354	-0.402	-0.324
indicator	(-2.14)**	(-0.58)	(-1.87)*	(-1.87)*	(-1.41)	(-6.81)***	(-3.16)***

Note: L=composite labor market flexibility index; M= minimum wage; H=hiring and firing regulation; C=centralized collective bargaining; MCH=mandated cost of hiring; MCW=mandated cost of work dismissal; CO= conscription. T-statistics based on robust clustered standard errors in parenthesis. *,**,*** denote significance at 10 percent, 5 percent, and 1 percent, respectively.

Table 4. Medium-Term Effect of Labor Market Policies—OLS

	L	M	H	C	MCH	MCW	CO
				Unemployment			
Reform $_t$	-0.740	-0.844	0.516	-0.387	-0.900	0.040	-0.220
	(-3.05)***	(-4.07)***	(1.38)	(-0.88)	(-5.52)***	(0.12)	(-0.28)
				Youth Unemployment			
Reform $_t$	-1.908	-1.983	0.349	-1.207	-1.921	n.a.	-1.627
	(-3.64)***	(-4.64)***	(0.38)	(-0.62)	(-4.53)***		(-0.56)
				Long-term unemployment			
Reform $_t$	-0.621	-0.073	-7.283	-1.728	-0.410	n.a.	-6.225
	(-0.54)	(-0.07)	(-4.02)***	(-0.71)	(-0.30)		(-1.88)*

Note: L=composite labor market flexibility index; M= minimum wage; H=hiring and firing regulation; C=centralized collective bargaining; MCH=mandated cost of hiring; MCW=mandated cost of work dismissal; CO= conscription. T-statistics based on robust clustered standard errors in parenthesis. *,**,*** denote significance at 10 percent, 5 percent, and 1 percent, respectively.

Table 5. Probability of Large-Scale Changes in Labor Market Institutions

	(I)	(II)	(III)	(IV)	(V)	(VI)	(VII)
l.area5b	-0.241	-0.253	-0.240	-0.210	-0.238	-0.243	-0.266
	(-4.22)***	(-3.93)***	(-4.20)***	(-4.16)***	(-4.42)***	(-4.25)***	(-4.01)***
gap_growth	-0.046	-0.065	-0.046	-0.046	-0.049	-0.046	-0.064
	(-2.20)**	(-2.40)**	(-2.20)**	(-2.19)**	(-2.50)**	(-2.20)**	(-2.37)**
system	0.268	0.244	0.269	0.279	0.281	0.269	0.269
	(2.89)***	(2.33)**	(2.87)***	(2.90)***	(3.25)***	(2.91)***	(2.70)***
prtyin	-0.013	-0.011	-0.013	-0.013	-0.001	-0.013	-0.009
	(-1.81)*	(-1.51)	(-1.81)*	(-1.84)*	(-1.13)	(-1.82)*	(-1.25)
l.lur	-	-0.012	-	-	-	-	-0.010
		(-0.78)					(-0.63)
lnpop	-	-	0.006	-	-	-	-0.023
			(0.12)				(-0.35)
lnopenk	-	-	-	-0.025	--	-	-0.101
				(-0.14)			(-0.44)
lncg	-	-	-	-	-0.050	-	0.014
					(-0.24)		(0.05)
d.l.cg	-	-	-	-	-	0.054	0.288
						(0.72)	(1.74)*
Pseudo R^2	0.09	0.09	0.09	0.09	0.08	0.09	0.10
N	667	510	667	667	667	667	510

Note: Z-statistics based on robust clustered standard errors in parenthesis. *,**,*** denote significance at 10 percent, 5 percent, and 1 percent, respectively.

Table 6. Medium-term Effect of Labor Market Policies-IV Robustness Checks

	(I)	(II)	(III)	(IV)	(V)	(VI)	(VII)
Policy $_t$	-2.478 (-4.76)***	-2.478 (-4.76)***	-1.991 (-3.92)***	-2.456 (-4.81)***	-2.512 (-4.81)***	-2.358 (-4.71)***	-1.879 (-3.84)***
Controls							
$U_{i,t}$	-1.546 (-10.23)***	-1.546 (-10.23)***	-1.553 (-10.33)***	-1.546 (-10.24)***	-1.546 (-10.22)***	-1.548 (-10.25)***	-1.554 (-10.35)***
$\Delta U_{i,t-1}$	0.462 (3.50)***	0.462 (3.50)***	0.459 (3.47)***	0.462 (3.50)***	0.462 (3.50)***	0.461 (3.49)***	0.458 (3.47)***
$\Delta U_{i,t-2}$	0.347 (3.34)***	0.347 (3.34)***	0.349 (3.41)***	0.347 (3.34)***	0.347 (3.33)***	0.348 (3.36)***	0.349 (3.43)***
Crisis dummies	0.591 (1.38)	0.591 (1.38)	0.591 (1.38)	0.591 (1.38)	0.591 (1.38)	0.591 (1.38)	0.590 (1.38)
R^2	0.85	0.84	0.85	0.85	0.84	0.85	0.85
N	380	380	380	380	380	380	380

Note: T-statistics based on robust clustered standard errors in parenthesis. *,**,*** denote significance at 10 percent, 5 percent, and 1 percent, respectively. Each column corresponds to the different Probit estimates reported in Table 4.

Table 7. Medium-Term Effect of Labor Market Policies—Expected vs. Non-expected

	(I)	(II)	(III)	(IV)	(V)	(VI)	(VII)
Policy $_t$							
Anticipated	-3.795	-3.652	-3.792	-3.806	-3.806	-3.753	-3.480
	(-2.26)**	(-2.31)**	(-2.26)**	(-2.27)**	(-2.27)**	(-2.25)**	(-2.38)**
Non-anticipated	-0.721	-0.710	-0.721	-0.721	-0.721	-0.720	-0.694
	(-2.90)***	(-2.83)***	(-2.90)***	(-2.90)***	(-2.90)***	(-2.89)***	(-2.77)***
Controls							
$U_{i,t}$	-1.576	-1.581	-1.576	-1.576	-1.576	-1.576	-1.580
	(-6.93)***	(-6.91)***	(-6.93)***	(-6.93)***	(-6.93)***	(-6.92)***	(-6.89)***
$\Delta U_{i,t-1}$	0.427	0.438	0.427	0.427	0.427	0.428	0.442
	(2.33)**	(2.37)**	(2.33)**	(2.33)**	(2.33)**	(2.33)**	(2.38)**
$\Delta U_{i,t-2}$	0.349	0.348	0.349	0.349	0.349	0.350	0.354
	(2.63)***	(2.61)**	(2.63)***	(2.64)***	(2.64)***	(2.64)***	(2.67)***
Crisis dummies	0.636	0.649	0.637	0.637	0.637	0.637	0.657
	(1.12)	(1.15)	(1.12)	(1.12)	(1.12)	(1.12)	(1.16)
R^2	0.87	0.87	0.87	0.87	0.87	0.87	0.87
N	380	380	380	380	380	380	380

Note: T-statistics based on robust clustered standard errors in parenthesis. *,**,*** denote significance at 10 percent, 5 percent, and 1 percent, respectively. Each column corresponds to the different Probit estimates reported in Table 4.

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ANNEX—DATA

The dependent and control variables included in the analysis belong to one of several categories, namely:

Unemployment
- Unemployment rate (*lur*, from WEO): Percentage of the total labor force that is currently unemployed.
- Youth unemployment rate (*unempyouth*, from WDI): Percentage of the total labor force of ages 15 to 24 that is currently unemployed.
- Long-term unemployment (*unemplong*, from WDI): Fraction (in percent) of the unemployment rate that is of long-term.

Macroeconomic variables
- GDP per capita (*rgdpl*, from WEO): Purchasing power parity (PPP) converted GDP per capita (with the Laspeyres methodology), derived from growth rates of private consumption, government expenditures, and investment at 2005 constant prices.
- Demand pressure (*gap_growth_n*): Gap in the current real GDP per capita growth with respect to a moving average of n years, centered at the current period.
- Government size (*lncg*, from PWT): (log) Government consumption share of PPP converted GDP per capita at current prices, in percent.
- Openness (*lnopenk*, from PWT): (log) Openness at 2005 constant prices, in percent.

Demographic variables
- Population size (*lnpop*, from PWT): (log) Total population (in thousands).
- Urbanization (*lnurbpop*, from WDI): (log) Urban population, as percent of total population.
- Density (*lnpopdens*, from WDI): (log) Population density, measured by the number of people per square kilometer of land area.

Political institutions
- Executive system (*system*): Assigns values of 0 if the system is "presidential", 1 if there is an Assembly-elected President, and 2 if the system is "parliamentary".
- Length of party of executive chief in office (*prtyin*): Number of years that the party of the executive chief is in the office.

Financial crisis
- Financial crisis indicator (*crisis*): This dummy variable assigns a value of 1 to years in which a country was going through a financial crisis according to Laeven and Valencia (2010), and 0 otherwise.