Department of Economics

Essays on Labor and Development Economics

Edwin Antonio Goñi-Pacchioni

Thesis submitted for assessment with a view to obtaining the degree of Doctor of Economics of the European University Institute

Florence, June 2011

EUROPEAN UNIVERSITY INSTITUTE
Department of Economics

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To my family. To those who were, are and will be there.

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Abstract

This work comprises four essays in two related areas: labor and development economics. On the labor side, two essays study (i) the effects of sizeable policy reforms over labor informality and (ii) the relation between productivity and wages in a context of substantial informality and high turnover rates. On the development side, two essays provide a comparison between developed and developing countries in the following aspects: (iii) the degree of complementarity of production factors and their capacity to translate R&D investments into economic growth and (iv) the effects of fiscal redistribution over income inequality.

Within the context of Latin America - the most income-unequal and labor-informal region in the world - this work intends to augment the understanding of the behavior, dynamics, interactions and contributions of productive factors (labor and innovative capital) and the effects that policies aimed at formalizing labor, innovating capital or redistributing factors retributions may have. The study applies recent measurement techniques and exploits rich novel datasets which combined with reformulated models help us to propose alternative appealing explanations.

Lessons learnt from these four essays suggest that (i) job dynamics play a fundamental role in the success (or failure) of policies aimed at promoting labor formality. Against the conventional wisdom, we contend that reductions in hiring rather than increases in separation rates are the main determinants of informality increases following protectionist policies. (ii) Job dynamics also play a differentiating role in the determination of wage-productivity elasticities and income risk (with new hires reacting more than incumbents). (iii) Yet, returns of labor and physical capital are constant across countries and periods regardless the stage of development whereas they exhibit an inverted U shape for technological capital (this is, highest returns observed for mid developed cases). (iv) Comparable private returns of productive factors are mirrored in comparable market income inequality measures observed across some developed and developing regions. However, while in Europe fiscal redistribution helps to achieve better distributed disposable income, in Latin America fiscal redistribution has meager or even countervailing effects.

Foreword

Essays on Labor and Development Economics is the result of a long journey dated back to late nineties when I started doing applied economic research.

Experience gathered doing applied research in the academic and public sectors in Peru and in the headquarters of The World Bank shaped my academic and professional interests, technical and topic wise. Among others, these interests comprise the study of labor markets and economic development from an applied perspective, exploiting novel techniques and rich empirical sources.

A reason to pursue this approach is that far-from-orthodox economic outcomes in emerging and less developed countries (LDC) challenge continuously the prescriptions of canonical theoretical models. Many of such unconventional outcomes were difficult to address inside and outside the scope of the mainstream of the literature -conceived as a set of theories and models aimed at explaining a world endowed with particular *givens*[1]. This was the case until recent emergence of very rich sources of information or novel technical methods developed to better exploit such an information. Nowadays to think outside the box of the conventional mainstream with some grounds stronger than just theoretical assumptions is becoming more likely. To this end, empirical sources are proving to be a nontrivial complementary input to produce knowledge and understanding of several economic phenomena, especially in the insufficiently explored world of LDC, or even to revisit and reformulate canonical models that have failed to show a correspondence between predictions of their essential components and its actual observed values in the vastly explored developed world.

This work is aimed at contributing to the understanding of some relevant issues in economies that cannot necessarily be seen through the lens of the canonical models provided by the mainstream but through the lens of contesting means, namely rich novel data and recent measurement techniques which combined with reformulated models help us to propose alternative appealing explanations to such issues. As it happens with those endorsing more conventional and theoretical approaches, our intention is to see what we can learn from the outcomes of our approach eventhough the caveats and limitations that every method comprises.

The result of this research is presented in four essays (chapters):

Chapter 1 presents the results of exploiting dense rotational labor panel data sets in order to understand the effects of radical policy reforms[2] on the dynamics across employment and unemployment states[3] into a market with high labor informality and within the context of a Search and Matching model. Our findings suggest meager effects of Trade Reforms but noticeable effects of Constitutional Reforms in the size and volatility of the flows of different labor sectors. Against the conventional

[1] Initial conditions such as the rule of law, institutional framework, market imperfections, initial quantities of factor endowments, initial qualities of factor endowments, etc. which can influence dramatically the feasibility of the assumptions imposed in most of the canonical models when generally applied to LDC.

[2] Namely, Constitutional Reforms comprising dramatic changes in labor regulation and Trade Reforms.

[3] And consequently on the size of different labor sectors.

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wisdom, we contend that reductions in hiring rather than increases in separation rates are the main determinants of these effects.

Chapter 2 extends the analysis of Chapter 1. This chapter explores the role of *marginal workers*[4] in the explanation of the the poor correspondence found between volatilities of productivity and wages[5]. Consistent with very recent contributions regarding the US economy we find that wage and productivity volatilities are more strongly related for those moving between distinct employment (or unemployment) categories than for those remaining in their same jobs. Our work adds to the existing literature by emphasizing the asymmetric reactions according to worker's mobility (new hires, sector movers, stayers) and formality (formal, informal) groups in both, the determination of wage-productivity elasticities and the income volatility process[6].

Chapter 3 presents the results of applying Instrumental Variable Varying Coefficient techniques in order to measure the contribution of R&D investments to the process of economic growth. Our findings contend the existence of a non linear relation between such a contribution and the level of economic development. This contribution departs from the conventional approach that treats homogeneously to all the observations of a pool of countries comprising developed and less developed economies. By using an advanced semiparametrical approach, our contribution also confirms and extends existing parametrical attempts conveying the message of nonlinearity in the contribution of R&D investments into economic growth of developed economies.

Chapter 4 takes a more aggregate perspective to study the implications of heterogeneity between observed units in the application of economic policy. This chapter presents the results of combining and exploiting national accounts information and government income and expenditure data of European and Latin American economies to see the differences of the redistributional impact of fiscal policy between these regions. Our findings show that while in Europe Fiscal redistribution helps to improve income equality, in Latin America fiscal redistribution seems not to work or worse, to have countervailing effects.

All of these results may be appealing for applied economists or policy makers interested in the functioning of LDC. The first two chapters shed some light to the understanding of how policies intended at first to improve quality conditions for workers in labor markets with non-negligible informal sectors may end up encouraging such informal labor arrangements. The asymmetric response of different type of flows (hiring, job to job, separation) in different sectors (formal, informal) play an important role in the determination of the final size of those sectors. Asymmetric reactions are also found in the way in which wages are set across sectors[7] or in how productivity shocks impact into workers' income risk according to their formality status. Chapter 3 extends our attention beyond the study of the labor factor and it analyzes the contribution of other productive factors to the process of economic growth. Departing from the microeconometric approach employed in the first two chapters, Chapter 3 operates at a more aggregate level and, using cross country multiperiod data, it stresses the importance of the asymmetric response of economic growth to R&D investments according to the development level of each observation. In spite of this apparent heterogenous composition of the factors or of the observations under study, Chapter 4 shows that compared to more developed countries, a sample of LDC (specifically some Latin American economies) fares similar high levels of market-income inequality. The difference between inequality of incomes at disposal of European and Latin American households relies in a great extent on the effectiveness of the redistributive role

[4]Those who do not stay in the same work category between two periods of analysis, say for example new hires or job movers.

[5]When measured at aggregate levels.

[6]Especially in its transitory component.

[7]Wages of newly hired or job movers show higher productivity elasticities than those of non-movers.

of the Government through fiscal policy (namely taxes and transfers). Such effectiveness is noticeable in Europe and still poor in Latin America where disposable-income inequality hovers around much higher levels than in the European region.

In this sense, these essays provide provocative findings that - I hope - will contribute to the discussion about both, the implications of policies aimed at promoting development in distinct dimensions (regulation of labor markets, investment in innovation and fiscal redistribution) and about some of the remaining challenges left for research striving to reconcile such findings with the theoretical mainstream.

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Chapter 1

Trade Liberalization, Labor Reforms and Formal-Informal Employment Dynamics [1]

Mariano Bosch[2]
Edwin A. Goñi-Pacchioni[3]
William Maloney[4]

Abstract

This paper studies gross worker flows to explain the rising informality in Brazilian metropolitan labor markets from 1983-2002. In particular, we examine the impact of trade and constitutional reforms, (that include increased firing costs, tighter restrictions on overtime work, and fewer restrictions on union activity) occurring during the period. We find aggregate sectoral movements to be driven largely by changes in the hiring rates which, in turn are driven largely by the constitutional reforms. Trade liberalization accounts for roughly 1% of the increase in informality, while the constitutional reforms account for 40-50%.

1.1 Introduction

A growing literature explores the insights that labor flows can shed about how regulations and institutions affect the functioning of labor markets. On the theoretical side, Bertola & Rogerson (1997); Alvarez & Veracierto (1999); Mortensen & Pissarides (2003); Pries & Rogerson (2005) have all analyzed the impact of policy reforms on labor market flows in a search and matching context. These modeling efforts offer well-defined predictions of gross labor flows and hence a disaggregated view of the processes underlying observed trends in unemployment stocks. For the developing world, the impact of regulations on what is perhaps the distinguishing characteristic of poor country labor markets, the large unregulated or informal sector, has recently been explored by Kugler (2004); Zenou (2008); Albrech *et al.* (2009).

[1] We are very grateful to Francisco Carneiro, Marcello Estevao, Gustavo Gonzaga Lauro Ramos, Jose Guilherme Reis and Gabriel Ulyssea, to participants in the NBER workshop on informality, October 2006 in Bogota, and to those at the University of Michigan conference on Labor Markets in Developing and Transition Economies, May 2007, in particular Gary Fields, Ann Harrison, Ravi Kanbur, and Jan Svejnar, for helpful advice and reality testing of ideas. We are grateful to Lauro Ramos for providing the old PME dataset for 2002. All conclusions are, of course, our own.

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On the empirical side, the advanced country literature has looked at the impact of employment protections on worker and job flows (see, for example, Messina & Vallanti 2007, for Europe, Kugler & Pica 2008, for Italy). Although data is less easily available[5], developing countries often offer more extreme policy experiments. Indeed, the evolution of the Brazilian labor market from 1990 to 2000 offers an especially dramatic experiment. Across this single decade, the share of the metropolitan area work force unprotected by labor legislation and thereby classified as "informal" rose an astronomical ten percentage points. Dramatic outcomes often spring from dramatic innovations and Brazil offers several significant policy changes across the period. The end of the 1980s saw a far reaching trade reform, and the establishment of a new constitution in 1988 that had substantial impacts on labor costs and flexibility. In particular, it increased overtime costs, raised substantially the penalty for firing workers, and importantly relaxed restrictions on union activity.

However, the empirical work to date has been surprisingly indeterminate. Looking at separations from formal sector work, Paes de Barros & Corseuil (2004) found, unexpectedly, no impact of the very large rise in firing costs. Looking at the impact of the reduction in trade protection, Menezes-Filho & Muendler (2007) find mixed results on outflows from formal employment and into informality, depending on specification, while Goldberg & Pavcnik (2003)'s work finds no impact on the size of the informal sector.

We revisit this experiment, analyzing the impact of a set of trade and labor reforms. We argue that, conceptually, the effects of these policies on the overall level of informality work through both relative informal/formal inflows and outflows and that the overall impacts can be ambiguous. This underscores the need to look at the full set of adjustments when evaluating the impact of reform. We then estimate the impact on the overall level of informality as well as the relative flows using a detailed and extensive rotating panel data set. Taking advantage of the differential impact of reforms across industries, we find little compelling evidence that trade reform was the prominent or even statistically significant factor. All three labor related dimensions of the constitutional reform however, appear more important. In all cases, the effect comes more through lower formal job finding rates as opposed to the separations that Paes de Barros & Corseuil (2004)investigated. We estimate that around 40% of the trend in informality can be explained by changes related to the constitutional reform while changes in trade can explain no more than 1% of the trend.

1.2 Data and Context

The period from the late 1980s to the first half of the 1990s was a turbulent one, comprising a persistent hyperinflation and six major stabilization plans designed to control it, a constitutional change, and several other reforms including a dramatic reduction in barriers to trade. Across the whole period Brazil experienced the 1990 crisis and slowdowns in 1999 and 2001 with corresponding recoveries.

We draw on the Monthly Employment Survey (Pesquisa Mensal de Emprego, hereafter PME[6]) that conducts extensive monthly household interviews in 6 of the major metropolitan regions (São Paulo, Rio de Janeiro, Belo Horizonte, Porto Alegre, Recife and Salvador) and covers roughly 25% of the national labor market. The questionnaire is extensive in its coverage of participation in the labor market, wages, hours worked, benefits received, and other variables that are traditionally found in

[5]Boeri & Burda (1996) found a limited impact on job matches of active labor market policies in the Czech Republic during the transition from socialism. Hopenhayn (2004) found that the introduction of fixed term contracts and of special trial period provisions in Argentina led to higher separations from formal employment. Kugler (1999, 2004) found the reduction of firing costs in Colombia led to greater exit rates in and out of unemployment as well as a reduction in unemployment

[6]For descriptions of the methodology of the Pesquisa Mensal de Emprego, see Sedlacek *et al.* (1990); IBGE (1991); Oliveira (1999).

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such employment surveys. We also draw on the National Household Survey (Pesquisa Nacional por Amostra de Domicilios or PNAD) for selected cross checks. The PNAD covers the entire country, but lacks any panel dimension and hence is not suitable for the study of gross labor flows.

The PMEs structure as a rotating panel allows us to create time series of gross labor market flows. It tracks workers across four consecutive months, then drops them from the sample for 8 months, and then reintroduces them for another 4 months. Each month one fourth of the sample is substituted with a new panel. After 12 months the initial sample is re-interviewed. Over a period of two years, three different panels of households are surveyed, and the process starts again with three new panels. Regrettably, the PME was drastically modified in 2002 and it is not possible to reconcile the new and old definitions for unemployment and job sectors[7]. Hence, our analysis begins in 1983 and stops at 2002.

We follow the literature in dividing employed workers into three sectors[8]. The formal salaried (F) are those public and private employees whose contract is registered or signed (*asinada*) in his/her work-card (*carteira de trabalho*) as dictated by Brazilian law. This registration entitles the worker to labor rights and benefits including 30 days of paid holiday per year, contributions for social security, the right to request unemployment benefit in case of dismissal, monetary compensation if dismissed without a fair cause, and maternity and paternity paid leave among others benefits. The informal salaried (I) are those employees whose work card has not been signed (*sem carteira* or without *carteira*). Finally, the informal self employed (S.E.) are workers who are not employees and hence are not covered by the benefits afforded by a signed work card. Ideally, we might also employ a definition of informality based on firm size as well, focusing on establishments of fewer than 5-10 as informal employees. However, the PME until 2002 does not tabulate this information[9]. That said, Henley *et al.* (2009) find that there is a close correspondence between access to protections, our measure of informality (employment registration), as well as size.

Figure 1.1 plots the share of informal employment (comprising both informal workers and self-employed) over total employment from 1983-2002. The share of informal employment remained relatively constant around 35% of the work force during the 1980's despite major macroeconomic shocks. However, as has been documented by numerous previous studies (See for example, Ramos & Reis 1997; Ramos 2002; The-World-Bank & IPEA 2002; Ramos & Brito 2003; Goldberg & Pavcnik 2003; Veras 2004; Ulyssea 2005; Ramos & Ferreira 2005b,a the share begins a major secular upturn at the beginning of the 1990s that levels off at 45%, 10 percentage points above its level at the beginning of the 1990s.

These movements in formal sector size are necessarily a function of inflows and outflows into each sector relative to the other. The next two panels of figure 1 present the evolution of these two series that compactly and completely capture the relevant dynamics: inflows into informal employment relative to those into formal; and outflows from informal employment relative to outflows from formal employment. We calculate relative inflows for each year the number of workers transiting into

[7]The unemployment rate jumps from 8% to 14% after the change in methodology of the PME.

[8]There is broad consensus in the literature on the definition of informality in the Brazilian literature. A comprehensive survey of work studying the size and evolution of the Brazilian informal sector in the labor market can be found in Ulyssea (2005) and a summary of stylized facts of the eighties and nineties is detailed in Ramos & Reis 1997; Ramos 2002; Ramos & Brito 2003; Veras 2004; Ramos & Ferreira 2005b,a

[9]The ILO for a period defined informality as consisting of all own-account workers (but excluding administrative workers, professionals and technicians), unpaid family workers, and employers and employees working in establishments with less than 5. In fact, Bosch & Maloney (2006) find that in Mexico, the criteria of small firm size and ours of lack of registration are similar in motivation conceptually and lead to a great deal of overlap. 75% of informal workers are found in firms of 10 or fewer workers. Since owners of firms or self-employed are not obliged to pay social security contributions for themselves, we in fact consider them as informal self-employed with no social security contributions (and hence without the benefits that are perceived by salaried workers holding a *carteira*).

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an informal sector job (from unemployment, out of the labor force or formal jobs) relative to those transiting to a formal sector job (from unemployment, out of the labor force or an informal jobs). Analogously, we calculate the relative outflows of informal jobs relative to formal jobs. It is clear that relative inflows into informality (formality) were strongly countercyclical (procyclical) until the beginning of the 1990s. However after 1992, the relationship breaks down with relative accessions into formality no longer tracking the economic recovery of the next five years. Relative outflows from informality (panel 3) also show a secular decline across the entire period.

Table 1.1 suggests that only a small part of this is due to changes in economic structure. Consistent with Figure 1.1 , there are virtually no changes in either the share of total employment by sector or the sectoral degree of informality in the 1983-1988 period. The 1990s, by contrast, saw a fall in the share of tradables (manufactures) of 10 percentage points, a phenomenon labeled the tertiarization of the Brazilian economy. However, the impact of this reallocation on informality is dwarfed by the intra-sectoral evolution: formality decreased within 28 of the 30 sectors in the table, falling 16% in manufacturing overall, and reaching 23% in some sectors[10]. This is broadly consistent with Ramos (2002); Goldberg & Pavcnik (2003) who find that the vast majority, of the increase in the informal employment, in the latter study eighty-eight percent, arises from movement of workers from formal to informal jobs within industries*within* industries[11]. Hence, the source of the secular rise in informality is largely working through the within composition of subsectors of workers, formal and informal and our modeling strategy reflects this.

1.2.1 Decomposing gross flows

We can generalize the preceding discussion at the industry level. In practice transitions into an informal job in a given industry may occur not only from other employment sectors, but also from other jobs within the same industry or from other industries. The law of motion of the number of informal workers in a given industry, m is given by

$$\dot{n}_{i,m} = u\times f_i^m + n_{f,m}\times d_m^m + \sum_{j\neq m}(n_{i,j}\times j_j^m + n_{f,m}\times d_j^m) - (s_i^m + g_m^m + \sum_{j\neq m}(j_m^j + g_m^j))n_{i,m} \quad (1.1)$$

Equation (1.1) indicates that the change in the total number of informal jobs in industry m is determined by four sets of flows. First, unemployed and out of the labor force workers (u) find informal jobs in industry m at rate f_i^m. Second, within industry m, formal jobs, $n_{f,m}$, are downgraded into informal jobs at a rate d_m^m. Third, from other industries informal ($n_{i,j}$) and formal ($n_{f,j}$) workers transit to informal jobs in industry m at rate j_j^m, and d_j^m respectively. Finally, the last term of equation (1.9) represents the outflow of informal jobs from industry m. This outflow may occur towards unemployment (and out of the labor force) at rate s_i^m, towards other informal jobs within industry m, g_m^m, and other jobs (formal and informal) in and industry different than m, $\sum_{j\neq m}(j_m^j + g_m^j)$). The analogous law of motion for formal jobs in industry m is:

$$\dot{n}_{f,m} = u\times f_f^m + n_{i,m}\times g_m^m + \sum_{j\neq m}(n_{f,j}\times t_j^m + n_{i,j}\times g_j^m) - (s_f^m + d_m^m + \sum_{j\neq m}(t_m^j + d_m^j))n_{f,m} \quad (1.2)$$

Where in this case t_j^m and t_m^j denote the job to job transition rates of formal workers between industries j and m respectively. The steady state relative sizes of the formal to informal sectors in

[10]The share of formal employment only increases considerably among domestic workers. This probably happened because the Union of Women Domestic Employees (UWDE) in Brazil - which originally was established as an association - won the status of a trade union in 1989. In 1992, UWDE became an affiliate of the Central Workers Union (CUT), which considerably increased the number of its members Ulku (2005)

[11]Similar results are reported by Bosch & Maloney (2006) for the Mexican case.

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Table 1.1: Changes in Informality by sector 1983-1988 and 1988-2000

Code	Sector	Informality Share				Employment Share			
		1983	2002	Change 1983-1988	Change 1988-2002	1983	2002	Change 1983-1988	Change 1988-2002
	All sectors	**0.36**	**0.45**	**-1.37**	**10.22**	**1.00**	**1.00**		
	Manufacturing	**0.14**	**0.28**	**0.09**	**14.16**	**0.32**	**0.22**	**0.96**	**-10.40**
400	Nonmetallic Mineral Goods	0.16	0.23	-0.79	6.99	0.02	0.01	-0.17	-0.54
401	Metallic Mineral Goods	0.12	0.28	-0.57	16.27	0.04	0.03	0.34	-1.25
402	Machinery and Equipment	0.06	0.22	0.66	16.19	0.02	0.01	0.09	-1.22
403	Electrical and Electronic Equipment and Components	0.05	0.26	0.14	20.82	0.02	0.01	0.13	-0.86
404	Vehicle and Vehicle Parts	0.04	0.09	-0.11	5.18	0.03	0.02	0.35	-1.27
405	Wood Sawing, Wood Products and Furniture	0.41	0.62	3.27	18.25	0.02	0.01	-0.07	-0.37
406	Paper Manufacturing, Publishing and Printing	0.14	0.29	0.79	14.76	0.02	0.02	-0.09	-0.37
407	Rubber	0.06	0.13	-0.32	7.77	0.00	0.00	0.06	-0.21
408	Footwear and Leather and Hide Products	0.19	0.19	0.50	-0.87	0.02	0.02	-0.03	-0.72
409	Non petrochemical Chemicals	0.06	0.13	-0.61	7.40	0.01	0.01	0.19	-0.84
410	Petroleum Refining and Petrochemical	0.04	0.09	-0.46	5.71	0.01	0.00	0.07	-0.19
411	Pharmaceutical Products, Perfumes and Detergents	0.13	0.19	-0.35	6.52	0.01	0.01	-0.05	0.02
412	Plastics	0.07	0.17	-0.49	10.32	0.01	0.01	0.16	-0.20
413	Textiles	0.10	0.23	1.19	12.24	0.02	0.01	0.04	-1.20
415	Apparel and apparel accessories	0.26	0.48	2.24	18.96	0.03	0.02	0.08	-1.02
416	Food	0.17	0.38	0.69	19.39	0.03	0.03	-0.12	0.10
417	Beverages	0.05	0.12	-0.55	7.71	0.01	0.00	-0.04	-0.28
	Services	**0.46**	**0.49**	**-1.62**	**5.07**	**0.68**	**0.78**	**-0.96**	**10.40**
700	Distribution of Water	0.05	0.23	0.07	17.63	0.02	0.01	-0.18	-0.60
701	Banking and Insurance	0.04	0.20	1.84	14.33	0.06	0.03	-0.78	-2.46
702	Transportation	0.36	0.48	-4.30	15.86	0.09	0.09	-0.43	0.89
703	Postal Services, Phones	0.03	0.16	0.73	11.81	0.01	0.02	-0.05	0.61
704	Lodging	0.58	0.57	-1.05	0.55	0.05	0.07	0.60	0.86
705	Repairs	0.76	0.77	-4.01	4.68	0.04	0.05	0.44	0.10
706	Clothing	0.92	0.88	0.04	-4.81	0.05	0.05	-0.57	0.39
707	Domestic workers	0.66	0.51	-4.03	-11.42	0.18	0.18	-1.56	1.69
708	Artistic, Radio	0.40	0.65	2.51	22.98	0.01	0.02	0.11	0.82
709	Technical	0.42	0.59	-0.41	18.11	0.04	0.07	0.58	3.08
710	Auxiliary	0.31	0.39	3.37	3.94	0.03	0.06	0.58	2.67
711	Social Services	0.22	0.42	0.53	19.67	0.04	0.04	-0.04	-0.29
712	Doctors	0.16	0.23	0.91	5.52	0.05	0.08	0.34	2.63

Note: The table shows the changes in employment and informality shares by sector.

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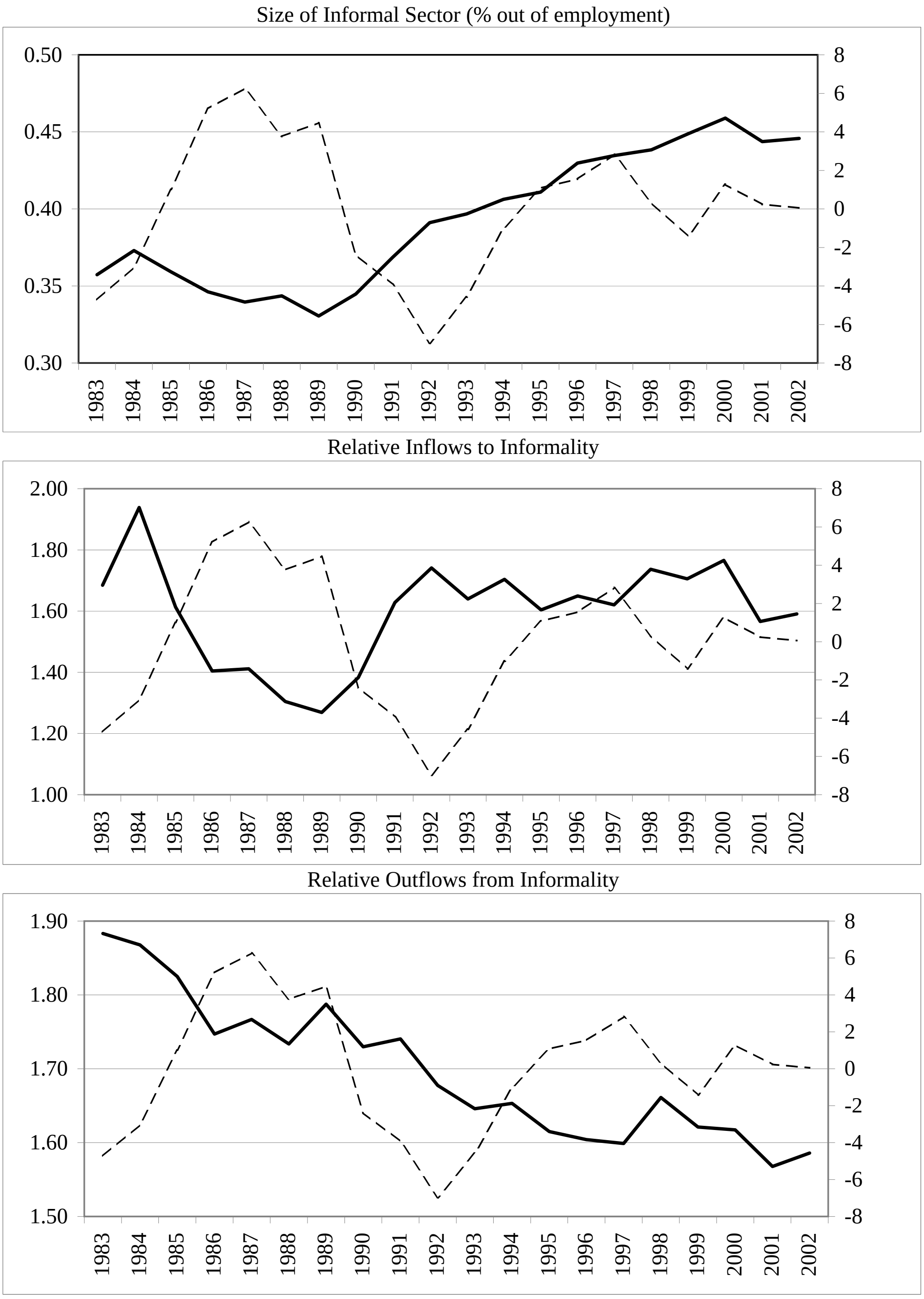
Figure 1.1: Size and Relative Informality Flows: Brazil 1983-2002



Note: The dash line represents the HP detrended Brazilian log GDP (right axis). The solid lines represent the values of the size of the informal sector, the relative inflows and the relative outflows to and from informality for the overall country.

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industry m can be written using equations Equations (1.1) and (1.2) as

$$\gamma_m^{ss} = \frac{RI_{i,m}}{RO_{i,m}} \tag{1.3}$$

Where $RI_{i,m}$ and $RO_{i,m}$ represent the relative inflows and outflows of informal workers in industry m. In steady state, these are given by,

$$RI_{i,m} = \frac{u \times f_i^m + n_{f,m} \times d_m^m + \sum_{j \neq m}(n_{i,j} \times j_j^i + n_{f,m} \times d_j^m)}{u \times f_f^m + n_{i,m} \times g_m^m + \sum_{j \neq m}(n_{f,j} \times t_j^m + n_{i,j} \times g_j^m)} \tag{1.4}$$

$$RO_{i,m} = \frac{s_i^m + g_m^m + \sum_{j \neq m}(j_m^j + g_m^j)}{s_f^m + d_m^m + \sum_{j \neq m}(t_m^j + d_m^j)}$$

Using the panel structure of the PME and equations (1.4) we can compute the steady state values of the share of informal employment of industry m.

$$i_m^{ss} = \frac{\gamma_m^{ss}}{1 + \gamma_m^{ss}} \tag{1.5}$$

These steady state values are remarkably similar to the actual series. Figure 2a shows the scatter plot of the actual share of informal workers and its steady state derived from equation (1.5) by industry for all years from 1983 to 2002. Virtually all data points lie close to the 45 degree line. Furthermore, we aggregate across sectors to show how the evolution of the steady state share of informal workers tracks the actual share over time. This is shown in figure 2b. The message is the same. Because the magnitude of flows is relatively high the steady state value of the share of informal employment is a very good approximation to the actual series.

In addition, this exercise allows us to decompose the changes in the share of informal employment by industry into changes in the relative inflows and the relative outflows. In particular, the growth rate of γ_m^{ss} can be decomposed into changes in $RI_{i,m}$ and $RO_{i,m}$.

$$\frac{\dot{\gamma}_m^{ss}}{\gamma_m^{ss}} = \frac{\dot{RI}_{i,m}}{RI_{i,m}} + \frac{\dot{RO}_{i,m}}{RO_{i,m}} \tag{1.6}$$

Table 1.2 shows this decomposition for the period 1983-1988 and 1988-2000 for all the industries in our sample. On average around 76% of the changes in the share of informal employment were due to the increased entry into informality relative to inflows into formal employment. The remaining 24% was due to changes in the relative outflows from informality. The rest of the article examines how the trade and labor reforms affected these flows and hence the aggregate rise in informality.

1.3 Policy Innovations

We explore two major policy shocks that were at play in Brazil across the period and which are likely to have had a major role on these flows, and hence in the reallocation of the work force from formal into informal employment: the opening of the Brazilian economy to foreign trade, and the 1988 constitutional reforms. Appendix 3.9 presents an illustrative model in the matching context that offers some structure for thinking through these effects in a matching context. For the most

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Table 1.2: Contributions of the Relative Inflows and Outflows into Steady State Informality Rate

Code	Sector	Relative Inflows		Relative Outflows		Steady State		Contribution of RI	Contribution of RO
		1988	2002	1988	2002	1988	2002	1988-2002	1988-2002
	All sectors	**1.22**	**1.59**	**1.75**	**1.58**	**0.41**	**0.50**	**0.76**	**0.24**
	Manufacturing								
400	Nonmetallic Mineral Goods	0.44	0.65	2.48	2.02	0.15	0.24	0.72	0.28
401	Metallic Mineral Goods	0.26	0.56	1.87	1.46	0.12	0.28	0.83	0.17
402	Machinery and Equipment	0.13	0.36	2.07	1.29	0.06	0.22	0.82	0.18
403	Electrical and Electronic Equipment and Components	0.15	0.60	2.75	1.57	0.05	0.28	0.88	0.12
404	Vehicle and Vehicle Parts	0.10	0.23	2.87	2.05	0.03	0.10	0.81	0.19
405	Wood Sawing, Wood Products and Furniture	1.18	2.23	1.46	1.39	0.45	0.62	0.95	0.05
406	Paper Manufacturing, Publishing and Printing	0.48	0.86	3.10	2.42	0.14	0.26	0.78	0.22
407	Rubber	0.14	0.30	2.68	2.43	0.05	0.11	0.93	0.07
408	Footwear and Leather and Hide Products	0.83	0.84	3.65	3.61	0.19	0.19	0.49	0.51
409	Non petrochemical Chemicals	0.15	0.29	2.58	1.80	0.05	0.14	0.75	0.25
410	Petroleum Refining and Petrochemical	0.06	0.19	1.87	1.70	0.03	0.10	0.96	0.04
411	Pharmaceutical Products, Perfumes and Detergents	0.29	0.61	2.41	2.22	0.11	0.21	0.93	0.07
412	Plastics	0.16	0.41	2.46	1.75	0.06	0.19	0.84	0.16
413	Textiles	0.37	0.62	3.01	2.19	0.11	0.22	0.71	0.29
415	Apparel and apparel accessories	0.77	1.63	2.00	1.67	0.28	0.49	0.87	0.13
416	Food	0.50	1.24	2.72	1.95	0.16	0.39	0.84	0.16
417	Beverages	0.17	0.18	3.50	1.81	0.05	0.09	0.11	0.89
	Services								
700	Distribution of Water	0.31	0.58	4.78	1.72	0.06	0.25	0.57	0.43
701	Banking and Insurance	0.28	0.69	4.69	2.67	0.06	0.20	0.77	0.23
702	Transportation	0.21	0.43	1.90	1.77	0.10	0.20	0.94	0.06
703	Postal Services, Phones	0.87	1.55	5.70	2.48	0.13	0.38	0.58	0.42
704	Lodging	2.50	2.58	2.05	1.92	0.55	0.57	0.31	0.69
705	Repairs	2.24	3.01	0.89	0.90	0.71	0.77	1.03	-0.03
706	Clothing	8.22	6.61	0.83	0.95	0.91	0.87	0.57	0.43
707	Domestic workers	2.13	1.90	1.30	1.88	0.62	0.50	0.19	0.81
708	Artistic, Radio	1.69	2.59	1.94	1.46	0.47	0.64	0.68	0.32
709	Technical	0.86	1.59	1.25	1.14	0.41	0.58	0.91	0.09
710	Auxiliary	0.62	0.75	1.17	1.21	0.35	0.38	1.17	-0.17
711	Social Services	0.61	1.16	2.26	1.67	0.21	0.41	0.78	0.22
712	Doctors	0.48	0.60	2.22	2.01	0.18	0.23	0.72	0.28

Note: The table shows the contributions of relative inflows and outflows in the steady state informality level by industry.

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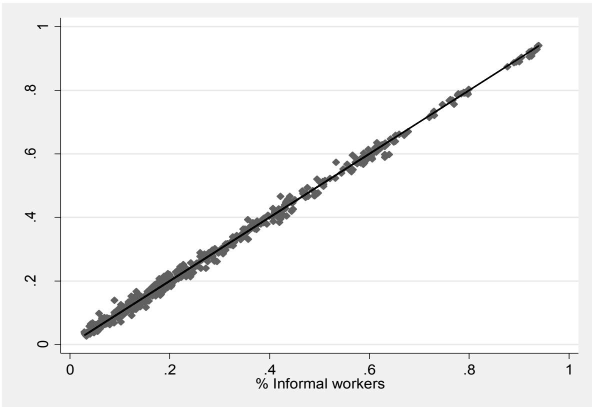
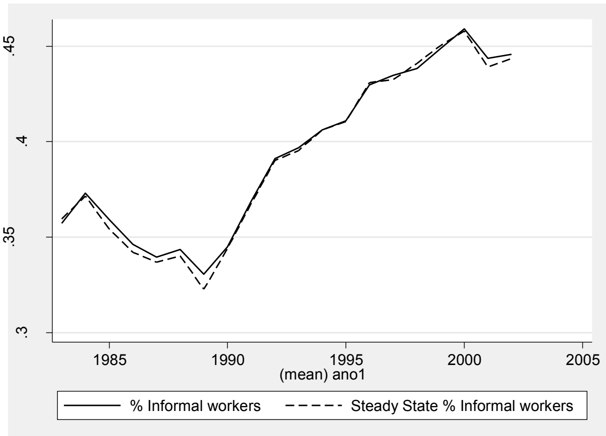
Figure 2a: Steady State Informality by Industry



Notes: The figure shows the actual % of Informal workers by sectors and its steady state valued calculated from inflows and outflows of worker according to equation () for all years. The solid line is a 45 degree line.

Figure 2b : Steady State Informality 1983-2002



Notes: The figure shows the evolution of the actual series of the % of informal workers and its steady state valued calculated from inflows and outflows of workers. The steady state value is computed by calculating the steady state of each sector year cell according to equation (5) and averaging across years using employment weights.

part, however, the effects are straightforward heuristically and the model mostly helps in showing whether we may expect an impact on relative flows in or out of formality (or both) and where net effects may be ambiguous.

It is worth highlighting that conceptually, these effects can occur in a context of a competitive or integrated labor market where, at the margin, workers are indifferent between formal and informal sectors. Hence the model in Appendix 3.9 works more in a context of informality as discussed by

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Maloney 1999; Bosch & Maloney 2010 among others where the market is not necessarily segmented and hence informality is not intrinsically inferior to formal employment. In very simple terms, any policy innovation that causes formal firms to see lower productivity per worker implicitly "shifts in" the demand for formal labor, but does not imply segmentation, since the marginal worker is indifferent between working in the formal or the informal sector. Similarly, changes in labor regulations may lead to a shifting up of the labor supply curve in such a way that less formal labor is hired, but, again, there is no segmentation. Clearly, this approach does not exclude segmentation emerging from the reforms as well.

1.3.1 Trade Liberalization

Far reaching trade reform began in the mid 1980s but intensified around 1990. Figure 1.2 plots two variables measuring the degree of trade protection of the Brazilian economy during the period; Muendler (2004)'s import penetration ratio, and Kume *et al.* (2003)'s real effective trade tariffs rates[12]. The trade opening translated into a dramatic reduction to one third of the level of effective protection (from 1988 to 2002) and to a doubling of imports penetration rates (during the same period).

As Goldberg & Pavcnik (2003) argue, the fact that Brazil had not participated in the tariff-reducing GATT rounds prior to the trade reforms implies that the usual concerns about the endogeneity of trade policy changes and political economy of protection are attenuated. As they argue, the government's objective when reducing tariff rates across industries was to achieve the relatively uniform tariff rate negotiated with the WTO and hence policy makers were accordingly less concerned with catering to special lobby interests. This is supported by figure 1.3. We compare the tariff levels of 1988 against those in the year 2000. We confirm Goldberg & Pavcnik (2003) findings that tariff declines in each industry are proportional to the industry's pre-reform tariff level in 1988.

The impact of trade liberalization is theoretically ambiguous. On the one hand, lower barriers increase the competition that an industry exposed to trade faces. The lost formal profitability both reduces hires into, and increases separations from, the formal sector. Both effects work in the same direction of a reduction in the share of formal employment. However, reducing tariffs and quotas also permits greater access to imported capital goods and other intermediate inputs that may increase productivity, or improve the quality of output that may enhance competitiveness relative to imports or in export markets. This implies exactly the opposite effects.

To date, the most thorough test of the hypothesis of a relationship between trade liberalization and informality was undertaken by Goldberg & Pavcnik (2003) who, exploiting sectoral variation in protection across time, found no relationship with the share of informality in Brazil, and a modest relationship in Colombia. More recently, Soares (2005); Menezes-Filho & Muendler (2007) find a significant effects of trade liberalization on the labor market although the impact on the size of the informal sector is sensitive to the inclusion of fixed effects.

We follow Goldberg & Pavcnik (2003) in exploiting the inter-industry variation of the impacts of trade liberalization over time using the Muendler (2004) and Kume *et al.* (2003) proxies. The variation is large across sectors especially since we expand our coverage beyond the manufacturing sectors and include non-tradable sectors services which, in principle, are less directly affected by the opening of the economy. Panel (a) in figure 1.4 plots the log change in the share of informal workers by industry against the log change in tariffs and import penetration from 1988 to 2000, suggesting an unconditional correlation between the change in informality with changes in effective tariffs and virtually zero correlation with changes in import penetration.

[12]Effective protection is preferred to nominal tariffs as before 1988 non-tariff barriers implied that most tariffs were redundant, That is the tariffs exceeded the differential between internal and external prices (see Hay 2001; Kume *et al.* 2003)

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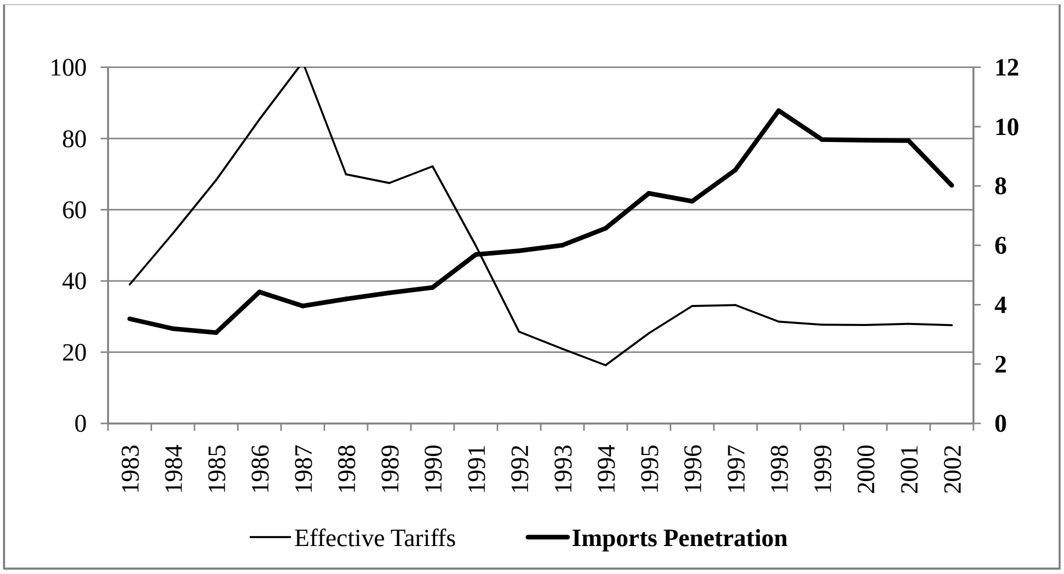
Figure 1.2: Effective Tariff Protection and Import Penetration



Note: Figures correspond to weighted averages of all industrial sectors (with weights given by the size of each industrial sector). Tariffs are obtained from Kume *et al.* (2003) for 1987-1998; Pinheiro & de Almeida (1994) for 1983-1986. Imports penetration corresponds to weighted imports/consumption by industry and are obtained from Muendler (2004) for 1987-1999; Pinheiro & de Almeida (1994) for 1983-1986; Nassif & Pimentel (2004) for 1999-2002.

1.3.2 Constitutional changes

The 1988 Constitutional changes had important implications for the labor code in several areas that theory predicts could lead to increasing informality. First, there was a generalized increase in labor costs and reduction in formal employer flexibility. Maximum working hours per week were reduced from 48 to 44, overtime remuneration was increased from 1.2 to 1.5 times the normal wage rate; vacation pay was raised from one to 4/3 of the monthly wage, and maternity leave increased from 90 to 120 days[13]. Second, the power of organized labor was expanded. Unions were no longer required to be registered and approved by the Ministry of Labor; decisions to strike were now left purely to union discretion, and the required advance notification to the employer cut from five to two days; and strikes in certain strategic sectors were no longer banned. Finally, firing costs were raised. The penalty levied on employers for unjustified dismissal, a category encompassing most legitimate separations for economic reasons in the US, increased by four times from 10% to 40% of the accumulated separation account (FGTS, *Fundo de Garantia por Tempo de Servico*). These are private funds into which the employer by law must contribute, every month, the equivalent of 8% of the employee's monthly wage. The accumulated value is thus a function of tenure and the average wage of the worker over that tenure. Workers only have access to the fund if dismissed, but on dismissal, they receive the entire fund, plus a penalty in proportion to the accumulated fund in the job from which they are being dismissed (See Paes de Barros & Corseuil (2004) for more detail).

[13]Paes de Barros & Corseuil (2004) among others also note that the maximum continuous work day was reduced from 8 to 6 hours although the exact meaning of this is unclear given that 8 hours remains the standard work day.

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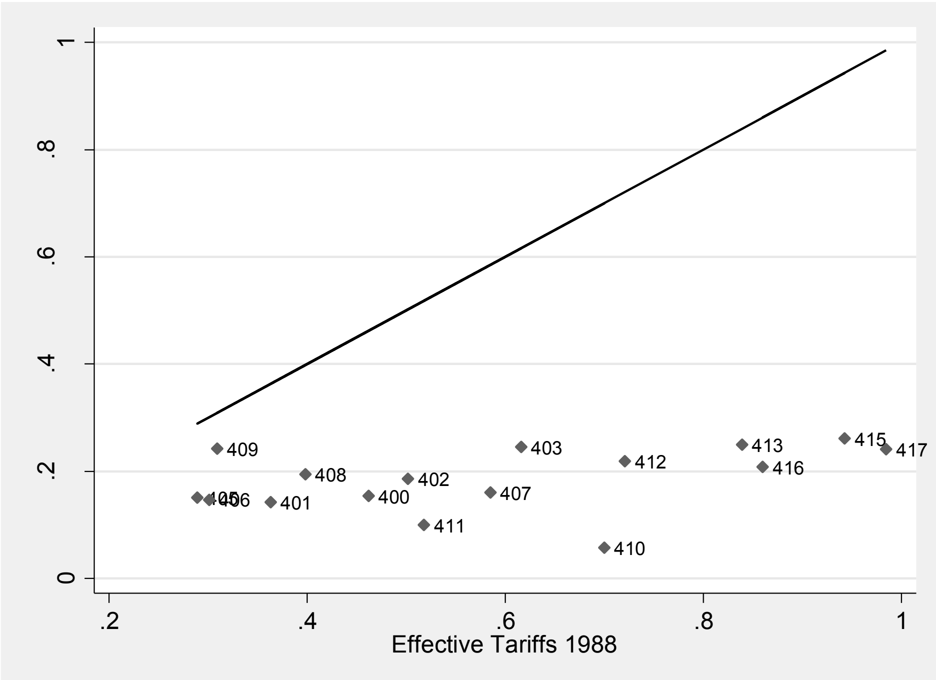
Figure 1.3: Effective Tariff 1988 and 2000 by Industry



Note: The figure shows the Effective Tariffs in 1988 vs Tariffs in 2000. The solid line is a 45 degree line. Tariffs are obtained from Kume *et al.* (2003) for 1987-1998; Pinheiro & de Almeida (1994) for 1983-1986.

To date, the most comprehensive work relating these changes to the functioning of the labor market was undertaken by Paes de Barros & Corseuil (2004) who find that separation rates decreased after the constitutional changes for short employment spells and increased for longer spells, but find inconclusive results about impacts on flows into informality from the formal sector. However, again, matching models such as that sketched in Appendix 3.9 suggest that several of these reforms would lead to a reduction in hiring (job finding) rates as opposed to the separations that Paes de Barros & Corseuil (2004) study.

We find very strong impacts through this second channel as well. The constitutional reform was implemented one off and simultaneously for all sectors. This implies that we must rely entirely on the cross industry variation in the impact of this shock for identification.
Overtime Pay:
An increase in overtime pay raises labor costs and this leads both to reduced formal hiring relative to informal hiring, and an increase in the relative separation from formal jobs (see Appendix 3.9). Both forces imply a reduction in the share of formal employment. Further, we argue that the impact will be greater in those industries where the use of overtime (prior to the reform) was greater. Hence, we expect that industries with a higher share of their working hours above the post constitutional maximum hours a week would see the greatest impact.

It is important to note a potential countervailing *numeraire* effect: Though total hours worked by formal workers may fall, the fewer hours that each employee may legally work implies that the number of workers may actually rise. Determining the net effect requires knowledge of the number of overtime hours, the cost imposed by the overtime legislation, and especially the elasticity of for-

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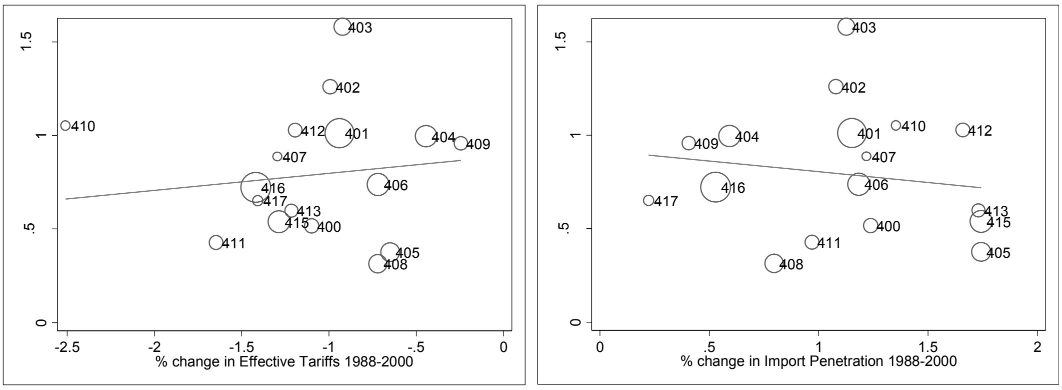
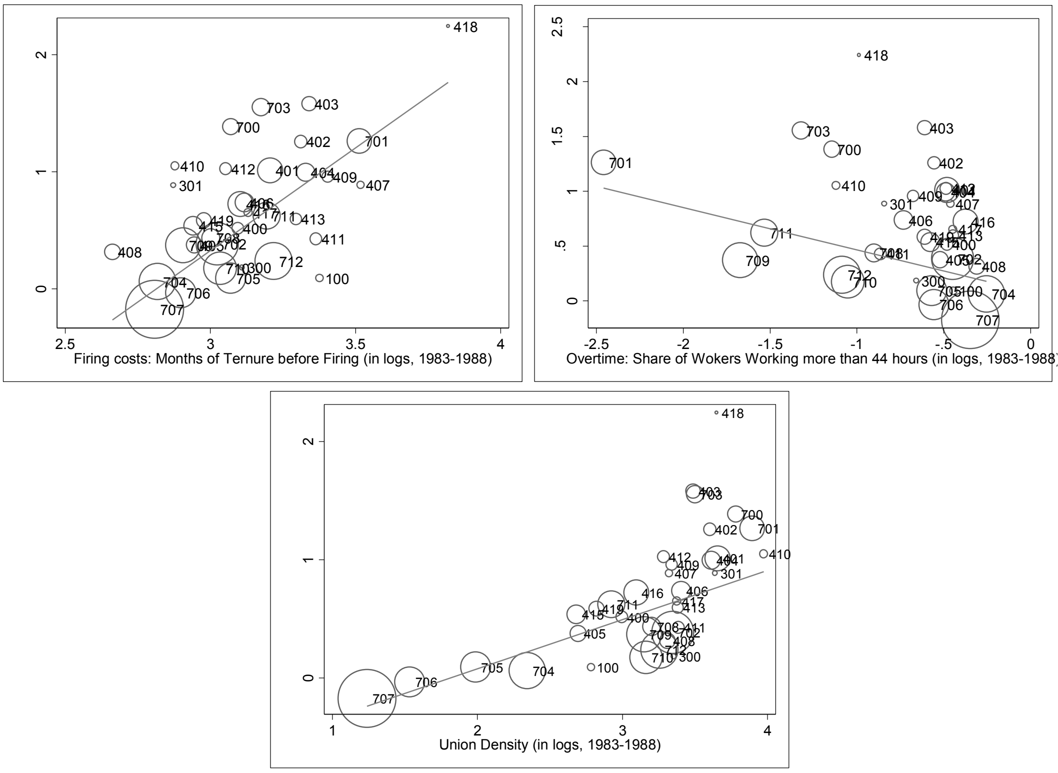
Figure 1.4: Unconditioned Correlations

Panel (a): Trade Variables Vs. Size of the Informal Sector



Panel (b): Constitutional Change Variables Vs. Size of the Informal Sector



Note: The graphs plots he changes in the share of informal workers from 1988-2000 against our 5 policy variables. Bubbles' sizes reflect the relative size of the particular sector (in terms of workers). See Table 1.1 for sector definitions.

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mal/informal labor demand, two out of three of which we do not know.

We use the pre-reform share of workers above the 44 hour limit as a proxy for the incidence of the reform in a given industry. Here again we find substantial variation in the data. For instance, while the non-metallic mineral goods industry was employing 60% of their workers above 44 hours, technical services were only employing 18% in 1988. These overtime workers were mostly formal (between 80% and 90%). After the constitutional reform, there was a substantial reduction in above 44 hours week in all industries indicating the strong reallocation effect of the policy. Panel (b) of Figure 1.4 plots the pre-reform share of share of workers working above 44 hours against changes in informality. It suggests that, unconditionally, the *numeraire* effect dominates, with industries with more pre-reform overtime showing a reduction in informal share.

Firing Costs:

As has been observed in the literature, increases in firing costs reduce both entry and exit and hence have ambiguous effects on formal employment (see for example,Kugler 2004). Raising the costs of firing a worker reduces formal sector. However, it also increases the cost of formal separations thereby decreasing the relative outflows from formal jobs.

Since the penalty firms have to pay upon dismissal is proportional to the wage of worker and the time the worker has spent in the firm, we are able to exploit the variation of the increase in F_m across industries. Translated into the parameters of the model this is

$$F_m = \rho T_m \bar{w_{f,m}} \epsilon \tag{1.7}$$

where ϵ is the share of the gross wage that gets accumulated into the FTGS (8%), $\bar{w_{f,m}}$ is the average formal wage in industry m, ρ is the penalty imposed by the government for unjustified dismissal and T_m is the average years of tenure of formal workers in the industry. As noted, the constitutional reform engineered a fourfold increase in ρ from 10% to 40%. This increase applied equally across sectors. Since the relevant firing cost must be standardize by the productivity of the worker being dismissed and hence wages, we exploit the only remaining source of variation across industries, the pre-reform variation in average turnover, T_m. We expect that industries with higher turnover will face higher penalties for dismissal and hence are more likely to be affected by the change in the penalty fee, ρ. In particular, we use the average pre-reform tenure of fired formal workers at the industry level as our source of variation. Intuitively, we argue that those industries that, on average, fire workers at with longer tenure, after the constitutional reform, faced significantly higher prices to do so.

Consistent with the discussion above, panel (b) of Figure 1.4 suggests an unconditional positive relationship between average pre-reform tenure of formal workers, and the change in the informal work force between 1988 and 2002.

Unions:

Finally, the degree of unionization may capture how the increased union power enhanced the bargaining position of workers and changed the incentives for firms to hire (and dismiss) formal workers. This can be shown to have a similar effect to an increase in overtime pay: lower formal hiring, higher formal firing and overall, a lower overall formal sector. We exploit the pre-constitutional variation in the unionization rate as a proxy for how the changes in the treatment of unions differentially affected sectors. This varies from 15% in some service sectors to 40% in heavy manufacturing. Panel (b) of Figure 1.4 suggests that unconditionally higher union density is correlated with higher levels of

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informality.

1.4 Estimation

We investigate the relationship between informality and our policy variables by estimating

$$Y_{jt} = \alpha_t + \alpha_j + TRADE_{jt}\beta_{TRADE} + D \times CC_j\beta_{D\times CC} + u_{jt} \tag{1.8}$$

Where Y_{jt} represents one of three dependent variables relating to industry j: the share of informal workers in total labor force, which we include as a reference, and the two variables of interest, relative informal inflows, or relative informal outflows, as defined in Equation (1.4).

The scalars α_t and α_j represents year and industry fixed effects, respectively, $TRADE_{jt}$ is a vector containing both the log of effective tariffs and the log imports penetration as defined above. Though, in theory, the two trade variables are imperfect proxies for the same phenomenon, since we are interested in capturing as much explanatory power from the trade liberalization that might be correlated with the constitutional variables, we include them both.

The effect of the constitution is captured by the vector CC_j which contains the log of unionization, the log of average firing cost by industry (proxied by pre-reform average tenure of fired formal workers), and the log of average overtime costs by industry (proxied by the pre-reform share of workers working more than 44 hours). Each variable is computed yearly for each of the 30 sectors (industrial and non industrial) from 1983-2002 based on the PME. D is a dummy variable with zeros up to 1988 (included) and ones thereafter.

Table 1.3 reports unit root tests for the various series. The Levin, Lin and Chu, and the Britung tests impose a uniform AR1 process across all panels while the Im Pesaran and Shin, ADF and Philips Peron tests allow different panels to exhibit different dynamics. The effective tariff is borderline I(0) and import penetration appears I(1) in the constrained tests but I(0) at the 10% level where the panels are allowed independent dynamics. For our reference dependent variable, the unconditional sector size, we cannot reject the presence of unit roots. Further, we find no evidence of cointegration with our explanatory variables so the reference specifications with these variables are under some suspicion of spurious regression. However, with the exception of the PP test (the least appropriate to our sample), for the two dependent variables on which our inference and simulations are based, the relative outflows and inflows series, are I(0)[14]. Algebraically, it is straightforward to show of the two I(0) series interaction allows them to generate the I(1) properties of the aggregate series. Though our analysis is focused on these flows, we report the results for the size variable for reference purposes, fully cognizant of the unreliability of the standard errors.

Our estimation strategy is twofold. As a first cut, we begin with a simple static specification. The first three columns of Table 1.4 present OLS estimations and suggest that both the trade variables and our proxies of constitutional reform played a role in increasing the share of informality by increasing relative inflows into informality (import penetration, firing costs, unionization) and decreasing the relative outflows of informality (import penetration and unionization). Consistent with the unconditioned correlations plotted in Figure 1.4 sectors with higher shares of overtime workers show smaller increases in informality after 1988.

However, these estimates are subject to Betrand & Mullainathan (2004)'s critique of the validity of the standard errors in situations where the observations across time on either side of the discon-

[14]The PP test is best suited to longer t panels and hence is the least appropriate here.

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Table 1.3: Unit Root Tests

	Sector Size	Relative Inflows	Relative Outflows	Effective Tariff	Imports Penetration	GDP
	Null: Unit root (assumes common unit root process)					
Levin, Lin & Chu t	0.11	0.01	0.00	0.07	0.29	0.00
Breitung t-stat	0.16	0.02	0.00	0.00	0.87	0.37
	Null: Unit root (assumes individual unit root process)					
Im, Pesaran and Shin W-stat	0.16	0.00	0.00	0.05	0.07	0.00
ADF - Fisher Chi-square	0.26	0.02	0.00	0.05	0.07	0.00
PP - Fisher Chi-square	0.98	0.17	0.00	0.18	0.28	0.00

Note: The table shows the p-values of the respective tests.

tinuity cannot be taken as independent. Since our constitutional variables are not continuous, but depend on the cross sectional variation across the discontinuity at 1988, this is a concern. We pursue two different approaches to correct for this. First, closely following Betrand & Mullainathan (2004) we abstract from the time dimension of the data and average the observations pre-1988 and post 1988. The results are presented in the second three columns of Table 1.4. As expected, the standard errors increase substantially, although though the point estimates are not substantially altered compared to the OLS results. We now find no significant impact of trade liberalization in the shaping of informality trends during the 1990s, consistent with Goldberg & Pavcnik (2003). In contrast, we still find significant effects of the constitutional variables albeit with somewhat lower levels of significance. Both a rise in firing costs and unionization increase the share of informality. While both firing costs and unionization increase relative informal hiring, only the prevalence of unions seem to have generated lower (higher) relative outflows from informality (formality). Restrictions on overtime, continue to lower the share of the workforce in informality (albeit only marginally significantly) through increased relative informal outflows.

The second approach allows for dynamics in the model by introducing the lagged dependent variable which is our preferred specification for several reasons. First, it allows for a more realistic modeling of the adjustment process. Second, it increases the number of observations usable in the regression while controlling for serial correlation across observations that would bias the standard errors. Third, it more efficiently uses the information from the continuous trade protection which is lost in the previous specification. We estimate a GMM system estimation model using internal instruments for the lagged depending variable. In particular, following Arellano & Bover (1995) we use lagged levels of the dependent variable dated $t-2$ and earlier as instruments for the equations in first-differences and lagged first-differences of the dependent variable as instruments for equations in levels.

The dynamic specifications are shown in columns (7) to (9) of table 1.4. The very significant and large coefficient on the lagged dependent variable suggests the importance of its inclusion. Further the test statistics for second order serial correlation reject the presence of serial correlation in the differenced residual in all three cases increasing our confidence in the reliability of the standard errors and the validity of the internal instruments. Overall, the results are broadly the same as those in columns (4) to (6). The short run coefficients confirm the roles of the firing costs and unionization variables. We now find that the overtime variable enters as a factor that increases (decreases) inflows

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Table 1.4: Effects of the Trade Liberalization and Constitutional Reforms. 1983-2002

	Baseline Specification			Ignoring Time Dimension			Dynamic Structure		
	Size	Relative Inflows	Relative Outflows	Size	Relative Inflows	Relative Outflows	Size	Relative Inflows	Relative Outflows
Trade Liberalization									
Tariffs	-0.030	-0.017	-0.002	-0.035	0.028	0.013	-0.012**	-0.008	0.001
	(0.019)	(0.019)	(0.011)	(0.121)	(0.095)	(0.052)	(0.005)	(0.005)	(0.005)
Import Penetration	0.110***	0.103***	-0.055***	0.056	0.092	-0.019	-0.008	-0.005	0.003
	(0.022)	(0.020)	(0.013)	(0.100)	(0.101)	(0.041)	(0.012)	(0.008)	(0.006)
Constitutional Change									
Firing Costs	0.478***	0.340***	-0.084	0.532**	0.374**	-0.098	0.064**	0.059*	-0.000
	(0.096)	(0.081)	(0.068)	(0.210)	(0.172)	(0.125)	(0.031)	(0.033)	(0.040)
Unions	0.191***	0.182***	-0.155***	0.195***	0.188***	-0.157***	0.018**	0.020**	-0.013
	(0.025)	(0.023)	(0.019)	(0.040)	(0.031)	(0.037)	(0.008)	(0.008)	(0.010)
Overtime	-0.221***	-0.063*	0.161***	-0.198*	-0.036	0.154***	-0.003	0.015*	0.008
	(0.044)	(0.035)	(0.030)	(0.097)	(0.059)	(0.050)	(0.013)	(0.008)	(0.018)
Lagged Dependent Variable							0.824***	0.683***	0.597***
							(0.047)	(0.044)	(0.030)
Long Run Effects									
Tariffs							-0.068**	-0.025	0.002
							[0.035]	[0.016]	[0.011]
Import Penetration							-0.045	-0.016	0.007
							[0.070]	[0.024]	[0.014]
Constitutional Change									
Firing Costs							0.362**	0.185*	-0.001
							[0.178]	[0.105]	[0.097]
Unions							0.101**	0.063**	-0.032
							[0.051]	[0.026]	[0.026]
Overtime							-0.018	0.047*	0.021
							[0.077]	[0.027]	[0.044]
First Order (p-value)							0	0	0
Second Order (p-value)							0.712	0.611	0.15
R2	0.644	0.664	0.497	0.870	0.895	0.826	0.936	0.873	0.763
Observations	600	600	600	60	60	60	540	540	540

Notes:
* significant at 10%; ** significant at 5%; *** significant at 1%
Sector size corresponds to the share of informal (I) workers (both informal salaried and self-employed) in an specific industry. Relative Inflows and outflows correspond to the new informal entries and exist into and from a particular industry relative to formal (entries and exits) according to equation (19). All pooled by year. Imports penetration corresponds to weighted imports/consumption by industry Firing costs corresponds to a dummy variable (active since 1989) interacted with the pre-treatment tenure (log of average months 1983-1988) of workers fired in the specific industrial sector. Unions correspond to a dummy variable (active since 1989) interacted with union enrollment - understood as % of unionized workers in the specific industrial sector (log of average 1986 and 1988). Overtime corresponds to a dummy variable (active since 1989) interacted with the proportion of workers working more than 44 hours in the specific industrial sector (log of average 1983-1987). The Baseline specification includes all years and industries (number of observations 600: 30 industries x 15 years). Ignoring time dimensions estimations takes average of all variables before and after the constitutional reform (number of observations 60: 30 industries x 2 time periods). Standard errors are clustered at the industry level. The Dynamics specification shows a GMM system estimator using lagged levels of the dependent variable dated $t-2$ and earlier as instruments for the equations in first-differences and lagged first-differences of the dependent variable as instruments for equations in levels, as suggested by Arellano & Bover (1995). In all dynamic estimations the Sargan Test cannot reject the exogeneity of the internal instruments.

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to informality (formality) and the counterintuitive finding on relative outflows disappears. This suggests that previous results were potentially capturing serial correlation and hence were inconsistent. We now do find effects of the tariff variable on the size of the informal sector. However, again, we are not entirely confident of these standard errors due to the non-stationarity of the series and because the estimates from the two stationary flows series suggests no effect. The bottom panel of the table presents the long run coefficients and the joint measures of their significance. The results remain largely unchanged although the magnitudes are only a third to one half compared to the previous exercise

In all, we find very little evidence that trade liberalization played a major role in shaping informality trends in the 90s. However, the constitutional variables appear consistently significant in our regressions. Importantly, as in Paes de Barros & Corseuil (2004), we find no impact of higher firing costs on relative outflows. However, we do find an increase in relative inflows into informality consistent with a fall in formal hiring suggesting that this heretofore unexamined hiring channel is important.

Robustness checks

Tables 1.5 reports a series of additional robustness checks of our main results for our preferred dynamic specification. First, we revisit Goldberg & Pavcnik (2003) concern that changes in the composition of the workforce are partly driving the trend. For instance, if formalization is related to educational attainment of the labor force, the observed rise in average years of education of workers rose across this period from 6.23 to 8.68 may change the propensity of workers to become formal. Though we do not postulate a reason why such shifts should be correlated with our covariates, we repeat the analysis controlling for shifts in observable worker characteristics by including a set of gender and education dummies as well as age and age squared. None of these controls enter significantly. Overall, the results stand.

Second, ideally, we would have a control of firms unaffected by both reforms to ensure that the variation in our variables is not picking up other effects, most notably the aggregate movements in the economy that might have affected high exposure and low exposure industries differently. To approximate this, we rerun our exercise over an arguably similar period, the 1982-83 recessions and the subsequent expansion. In this case we set the dummies underlying the proxies for CC equal to 0 for the period prior to 1985, a time in the cycle similar to that where the constitution was approved, and 1 for the three years after. The second panel of both tables shows that none of the variables in the specification emerge significantly with the exception of the overtime variable which enters weakly in the suspect aggregate size regression, Our proxies do not appear to be picking up any systematic correlations with periods of high or low growth.

Third, thus far, in the interest of explaining the overall increase in informality we have included the entire sector, both informal salaried workers and the self employed. The third panel of Table 1.5 removes the self-employed from the sample focusing only on the salaried sector. We obtain very similar results for the informal salaried alone in both orders of magnitude of the coefficients and degree of significance.

Another confounding factor determining hiring/firing decisions at the sector level is the evolution of sector productivity. However, for two reasons we do not include a proxy in our main specifications. First, several of the reforms are thought to have impacts through productivity and hence, we do not want to short circuit those effects by controlling for them. Second, the relevant measure of formal sector productivity should be complemented by the equivalent in the informal sector data which does not exist. Hence the inclusion of an overall sectoral output variable should be seen as a rough proxy for shared cyclical fluctuations that may be correlated with our proxies. The results of

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Table 1.5: Robustness Checks: Dynamic Structure

	Controlling by Human Capital			Placebo Law pre 1988			Only Informal Salaried			Including Sectorial Output			Unweighted		
	Size	Relative Inflows	Relative Outflows	Size	Relative Inflows	Relative Outflows	Size	Relative Inflows	Relative Outflows	Size	Relative Inflows	Relative Outflows	Size	Relative Inflows	Relative Outflows
Trade Liberalization															
Tariffs	-0.013**	-0.008	0.002	0.010	0.003	0.001	-0.007	-0.005	0.001	-0.010**	-0.008	-0.000	-0.012**	-0.008	0.001
	(0.006)	(0.005)	(0.005)	(0.012)	(0.016)	(0.013)	(0.005)	(0.006)	(0.005)	(0.005)	(0.006)	(0.005)	(0.005)	(0.005)	(0.005)
Import Penetration	-0.006	-0.003	-0.001	-0.061*	-0.020	-0.000	-0.007	-0.008	-0.001	-0.011	-0.005	0.003	-0.008	-0.005	0.003
	(0.012)	(0.010)	(0.005)	(0.035)	(0.014)	(0.011)	(0.010)	(0.006)	(0.005)	(0.013)	(0.008)	(0.006)	(0.012)	(0.008)	(0.006)
Constitutional Change															
Firing Costs	0.065**	0.062**	0.001	-0.007	0.001	0.002	0.040	0.024	-0.005	0.057*	0.059*	0.001	0.064**	0.059*	-0.000
	(0.031)	(0.030)	(0.028)	(0.005)	(0.002)	(0.001)	(0.030)	(0.024)	(0.034)	(0.034)	(0.034)	(0.044)	(0.031)	(0.033)	(0.040)
Unions	0.015	0.015**	-0.008	-0.052	-0.035	-0.020	0.023**	0.014**	-0.006	0.016*	0.020**	-0.012	0.018**	0.020**	-0.013
	(0.009)	(0.007)	(0.012)	(0.050)	(0.025)	(0.023)	(0.011)	(0.007)	(0.012)	(0.009)	(0.008)	(0.011)	(0.008)	(0.008)	(0.010)
Overtime	0.004	0.020*	0.002	0.027	-0.017	0.043*	-0.002	0.004	0.010	0.000	0.015*	0.007	-0.003	0.015*	0.008
	(0.018)	(0.011)	(0.022)	(0.051)	(0.016)	(0.022)	(0.012)	(0.007)	(0.015)	(0.017)	(0.009)	(0.019)	(0.013)	(0.008)	(0.018)
										-0.052	-0.007	0.029			
Sectoral Output										(0.037)	(0.035)	(0.021)			
Lagged Dependent Variable	0.825***	0.688***	0.612***	0.656***	0.749***	0.521***	0.773***	0.613***	0.584***	0.814***	0.683***	0.593***	0.824***	0.683***	0.597***
	(0.047)	(0.043)	(0.030)	(0.091)	(0.087)	(0.109)	(0.055)	(0.047)	(0.038)	(0.046)	(0.044)	(0.030)	(0.047)	(0.044)	(0.030)
Human Capital															
Primary Education	-0.016	-0.009	0.013												
	(0.014)	(0.019)	(0.016)												
Secondary Education	-0.005	0.002	-0.007												
	(0.007)	(0.009)	(0.006)												
Age	0.303	0.271	-0.356												
	(0.222)	(0.344)	(0.230)												
Age2	-0.131	-0.164	0.146												
	(0.110)	(0.164)	(0.094)												
First Order (p-value)	0.00	0.00	0	0.00	0.00	0	0.00	0.00	0	0.00	0.00	0	0.00	0.00	0
Second Order (p-value)	0.75	0.60	0.096	-	-	-	0.678	0.108	0.091	0.734	0.612	0.102	0.717	0.611	0.08
Observations	540	540	540	90	90	90	540	540	540	540	540	540	540	540	540
R2				0.961	0.934	0.892	0.925	0.853	0.753	0.936	0.874	0.763	0.912	0.820	0.659
				0.942	0.903	0.842	0.917	0.838	0.728	0.929	0.860	0.738	0.903	0.801	0.624

Notes: * significant at 10%; ** significant at 5%; *** significant at 1%. In all estimations the Sargan Test cannot reject the exogeneity of the internal instruments. See also Notes in Tables 1.3 and 1.4.

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including a sectoral output variable are shown in column four. The results do not change. Finally, we rerun the regression without weighting by sector size. Again, the results do not change.

It is possible that sectors with less formal flexibility in the face of trade liberalization may have recurred more to informal employment. Alternatively, those sectors that were most affected by constitutional reforms would have found themselves affected more if they faced high or increasing competition from abroad. In Table 1.6, we interact each one of our three constitutional variables with the two trade variables. We find statistically significant interactions between firing costs and tariffs, and between overtime pay and import penetration for both the overall size of the sector and for relative inflows. A consistent effect is found on the relative outflows side that firms more affected by overtime legislation increased their relative outflows with trade liberalization. Hence, there is some evidence that the two reforms taken together exacerbated the impact on informality that either would have had separately or if they had been designed to be complementary. However, as we show below quantitatively, trade reforms can explain little of the overall change in informality.

In sum, the impact of the constitutional variables appears robust to a variety of specification changes.

1.5 Explaining the Rise in Informality

We can now use our preferred estimates in Table 1.4 to ask how much each of the two sets of reforms analyzed in this paper may explain the increase in informal employment. We generate the counterfactual increases in the share of informal employment, relative informal inflows and relative formal outflows had imports and tariffs remained at its level in 1988 (in the absence of trade) and, second, had the constitution not being approved (in the absence of constitution).

Table 1.7 shows this exercise for our preferred dynamic specification[15]. The first row shows the actual trends from 1988 to 2002 for the size, relative inflows and relative outflows from informality. The share of informality over total employment increased by 0.90 percentage points a year, while the relative inflows into informality increased by 2.6 percentage points a year and the relative outflows from informality decreased by 3.4 percentage points a year. The next rows present the counterfactual changes in these three variables accounting for the effects of the reforms. We find that, had the trade reforms not taken place the relative inflows into informality would have increased by only 0.8% less, and relative outflows by 0.23%. By contrast, the constitutional reforms can explain up to 52% of the relative increase in inflows into informality although as expected, substantially less of relative outflows. The coefficients on the size variables suggest that trade can explain 1.2% and the constitution 43% of the increase. However, given our concerns about the non stationarity of this variable, we also present the simulated impacts on the steady state levels of informality using the counterfactual predictions of the relative flows and their contributions to changes in the level of informality from Table 1.2. We find that around 39% of the changes in the steady state level of informality can be attributed to the constitutional reform and less than 1% to trade. The changes occur, essentially, through changes in the hiring patterns of firms. The results show that trade can actually explain very little of any of the dynamics we observe in the data. The effect of the constitutional reforms again, on within sector informality, is sizeable.

Two sources of healthy skepticism have emerged about these results. First, Kucera & Roncolato (2008) argue that both real wages and unionization experienced falls across the period of magnitudes that potentially dwarf the cost implications of the constitutional reforms. On the first count, when the correct deflator is used (see Corseuil & Foguel 2002), real wages average roughly the same

[15] Very similar quantitative results are obtained if instead we use the estimates of Tables 1.5 and 1.6

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Table 1.6: Trade and Constitutional Interactions

	Dynamic		
	Size	Relative Inflows	Relative Outflows
Trade Liberalization			
Tariffs	0.005	-0.017	-0.002
	(0.010)	(0.016)	(0.017)
Import Penetration	-0.029	-0.012	0.008
	(0.018)	(0.011)	(0.013)
Constitutional Change			
Firing Costs	0.009	0.044	0.034
	(0.037)	(0.040)	(0.051)
Unions	0.018**	0.020**	-0.014
	(0.008)	(0.008)	(0.011)
Overtime	-0.019	0.011	0.014
	(0.014)	(0.011)	(0.017)
Interactions			
Firing Costs X Tariffs	-0.032**	-0.032*	-0.002
	(0.014)	(0.018)	(0.012)
Unions X Tariffs	0.003	0.010	-0.014
	(0.018)	(0.023)	(0.014)
Overtime X Tariffs	0.066	0.074	0.052**
	(0.050)	(0.052)	(0.024)
Firing Costs X Imp. Pent.	-0.025	-0.033	-0.020
	(0.024)	(0.029)	(0.019)
Unions X Imp. Pent.	-0.006	0.006	0.009
	(0.029)	(0.035)	(0.018)
Overtime X Imp. Pent.	0.116***	0.085*	0.017
	(0.038)	(0.046)	(0.026)
Lagged Dependent Variable	0.789***	0.678***	0.847***
	(0.049)	(0.044)	(0.061)
First Order (p-value)	0.00	0.00	0.00
Second Order (p-value)	0.99	0.60	0.15
Observations	540	540	540
R2	0.646	0.69	0.473

Notes: * significant at 10%; ** significant at 5%; *** significant at 1%. In all estimations the Sargan Test cannot reject the exogeneity of the internal instruments. See also Notes in Tables 1.3 and 1.4.

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Table 1.7: Actual and Latent Trends

	Size	Relative Inflows	Relative Outflows	Explained Changes in SS
Actual	0.909***	2.596***	-3.467***	
	(0.092)	(0.831)	(0.190)	
Latent (in the absence of Trade)	0.898***	2.575***	-3.459***	
	(0.090)	(0.827)	(0.192)	
Explained by changes in Trade	1.2%	0.8%	0.23%	0.65%
Latent (in the absence of Constitution)	0.516**	1.250	-3.447***	
	(0.177)	(0.715)	(0.290)	
Explained by changes in the Constitution	43%	52%	0.6%	39.15%

Notes: * significant at 10%; ** significant at 5%; *** significant at 1%.
The table shows the trends in the size, relative inflows and relative outflows of the informal sector. The actual coefficients are obtained from regressing the appropriate variable on a linear trend for the period 1988 to 2002. The latent trends are obtained by subtracting from the actual data the effect of either trade of constitutional variables using the dynamic specification in Table 1.4 and then regressing the resulting series on a linear trend. The last column shows the explained changes in the steady state level of informality using the results from relative inflows and relative outflows and the estimates from Table 1.2 on the contributions of each flow to the changes in the steady state.

level in the pre- and post-reform periods[16]. On the second count, our calculations with the PNAD confirm the findings from others that union density indeed fell across the period (see, for example, Arbache (2004)) and this presumably led to a decrease of indeterminate magnitude in the bite of the reforms. That said, our estimates rely on the variance of union density across sectors and this was preserved: the overall decline occurred remarkably uniformly leading to a correlation of the 1984 and 1999 sectoral densities of .93 and a rank correlation of .92. Hence, the pre-reform values of unionization that we use are well-correlated with ongoing cross-sectoral variation in union power, and are unaffected by the reforms which is precisely what we need for consistent estimation. Hence neither concern weakens our results.

The second concern is that 2000s witnessed a significant recovery of formality despite the persistence of the reforms above and, in fact, a sustained rise in the minimum wage. Though falling outside of the coverage of our panel and analysis, this evolution merits a brief comment.

First, the central fact about this period is the sustained expansion of roughly 5% after 2004. Informality is highly responsive to the business cycle and hence, raw comparisons of its evolution across periods might be misleading. Bosch *et al.* (2009) show the cyclical formal employment elasticity of 0.5 with respect to the cyclical component of output and we may expect the long run elasticity to be somewhat larger. Since this elasticity is effectively conditional on extant labor distortions, this implies a secular increase in the formal sector of at least 2.5% per year.

Second, kernel density plots (available on request) confirm that, indeed, the minimum wage became increasingly binding from 2002 to 2010. However, there are reasons to believe that this is consistent

[16]The official CPI implies a dramatic increase in real wages across the post reform period, a phenomenon discussed by Chamon (1998). However, Corseuil & Foguel (2002) argue that the official CPI series badly understates inflation in 1994 due to mishandling the coexistence of the Real and the Cruzeiro in that year. When their deflator is used and correctly centered in the series, the real wage, while showing some volatility, is roughly equal in the pre and post reform periods.

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Table 1.8: Measures of Informality by Geographical Division

Table 8a: Employment Shares by Geographical Division

Sector	Survey	Year	Urban	Rural	Metropolitan	Non-Metropolitan
Formal	PNAD	1981	57.37	18.51	65.57	38.92
	PME	1983			63.25	
Informal	PNAD	1981	42.63	81.49	34.43	61.08
	PME	1983			36.75	
Formal	PNAD	1990	54.57	26.68	61.04	42.16
	PME	1990			61.34	
Informal	PNAD	1990	45.43	73.32	38.96	57.84
	PME	1990			38.66	
Formal	PNAD	2001	46.78	20.00	51.52	38.95
	PME	2001			52.96	
Informal	PNAD	2001	53.22	80.00	48.48	61.05
	PME	2001			47.04	

Table 8b: Relative Growth Rates of Population by Geographical Division

	Survey	Year	Urban	Rural	Metropolitan	Non-Metropolitan
Millions of Inhabitants	PNAD	1981	55.12	19.56	24.81	49.87
	PNAD	1990	70.20	22.52	30.07	62.66
	PNAD	2001	102.80	18.21	39.87	81.14
Growth (in %)		1990/1981	27.36	15.17	21.19	25.64
		2001/1990	46.45	-19.16	32.61	29.51

Table 8c: Relative Growth Rates of Employed Labor Force by Geographical Division

	Survey	Year	Urban	Rural	Metropolitan	Non-Metropolitan
Millions of Workers	PNAD	1981	29.53	8.70	13.78	24.45
	PNAD	1990	40.87	11.05	17.79	34.13
	PNAD	2001	57.26	8.93	22.22	43.97
Growth (in %)		1990/1981	38.38	27.09	29.10	39.60
		2001/1990	40.11	-19.22	24.86	28.84

Notes: The figures consider to individuals above 15 years of age only. PNAD has national coverage. PME covers 6 major metropolitan areas only, but has a panel dimension. Panel (a) shows that both surveys yield similar employment shares in metropolitan areas. Panel (b) shows that, while there is clearly rural/urban migration this does not translate to substantial non-metropolitan to metropolitan shifts. Panel (c) is the counterpart of (b) for the employed labor force.

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with the observed rise in formality. First, as has been found in other studies[17], the minimum wage is often more binding in the informal sector than the formal sector, leaving some ambiguity about the degree to which relative wages have moved against formal employment for low wage workers. Second, Lemos (2009) finds no impact of the minimum wage on employment, although Neumark *et al.* (2004), more pessimistically, find that an increase that binds an additional 10% of the workforce reduces employment of household heads by 1.56 percentage points, with no or positive impact on non-household heads[18]. The rise from 13% to 17% bound from 2002 to 2010 would therefore lead to a maximum of 2.5 percentage points (assuming that half of the formal employed are household heads) decrease in formality, a fraction of our minimum estimate of the rise caused by growth.

Other possible explanations

In the simulations above, the trade variables with substantial variation explain around 1% of the secular movements in informality. The constitutional reform is able to explain around 40%. Ideally, we might have more time series variation that could concretely rule out other possible phenomena not related to labor market legislation. We briefly review two possible candidates.

First, along with the Constitutional reforms affecting labor markets were initiatives changing the nature of health system implemented in the early 1990s that granted universal access to health services[19]. Carneiro & Henley (2003) suggest that uncovered employment may have risen because employees and employers collude to avoid costly contributions to a social protection system that is perceived to be inappropriate, inefficient and poor value for the money[20]. In principle, then, a universalization of health care de-linked from the labor market may have changed the cost benefit analysis of being enrolled in, and hence contributing to, formal sector benefits programs. In the end, they conclude that this is unlikely, not only because public health services continued to be thought of as substantially worse than the formal sector product[21], but also because the effective supply of these services was available even for non contributors several years before the reforms took place: As early as 1981 roughly 49% of self employed and 59% of the informal salaried compared to 48% of the formal salaried report that they received attention from a public health provider[22]. Further, little progress had been made on implementing the measures contemplated in the 1991 Social Security Reform.

[17]This ”*efeito farol*” or lighthouse effect where norms in the informal sector are set in the formal has been documented earlier in the Brazilian case by Souza & Baltar (1979), Neri *et al.* (2000), and Latin America more generally by Maloney & Nunez (2001).

[18]Second, much of rise represents catch up from declines across the post reform period (see Lemos 2009). We find that the share of workers at the minimum wage in 2010 is 17%, up from 8% in 1998, but not so far above the 15% in the immediate post reform (1990-1994) period. Hence, viewed as an extension of our previous analysis, the net change in the minimum wage from the beginning of the reforms to the present is relatively small.

[19]Among the changes contemplated in the Social Security System Reform of 1991 (which comprises pensions, health, and social aid), health related amendments are the only candidates to be considered as possibly determinants. Although pensions reforms loosened the requirements to perceive a pension (age for eligibility and required years of services were lowered) and increased the benefits of recipients (see De Carvalho (2002) for a summary of the characteristics of the Brazilian security system before and after the reform), two reasons reduce its suitability to explain the composition and dynamics of the labor market: first, benefits are computed as a function of documented past earnings over the cumulated time of services except for those perceiving the minimum pensions hence in any of those cases there is no incentive for workers to move between formality or informality because of potential gains in switching due to pensions; second, the reforms should have exerted more effects over the elder population close to retirement which is not the critic mass driving the size and dynamics of the labor sectors.

[20]Their estimates suggest that the earnings premium needed in the marketplace to compensate covered workers for having to make social security contributions varies between 7.5% and 12.2% of the mean uncovered hourly wage.

[21]The public system acts as a floor, available to all but used primarily by the lower classes (Jack, 2000). Although evaluation of standards for minimum quality in infrastructure, human resources, ethical, technical and scientific procedures in hospitals have been implemented, these practices are far from being universal in the services network, (Pan-American-Health-Organization, 2005).

[22]PNAD 1981. Non urban dwellers excluded.

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Second, there was an increase in the magnitude of flows from the rural to the urban areas across the 1990s that, in principle, were it all directed toward the informal sector, might explain part of the rise[23]. Table 1.8 suggests that this is not the case. Panel (a) shows that the PNAD survey, which covers the entire population, yields very similar formal and informal sector shares when the sample is restricted to the metropolitan regions covered by the PME[24]. This cross check suggests that our two surveys are telling very consistent stories. The PNAD also shows that the rural sector is far less formal and hence the idea that migration to the city might lead to urban informalization is plausible[25]. However, panels (b) and (c) show that, while there was substantial rural to urban migration across the period, this did not map into a large shift in metropolitan-non metropolitan shares of population or of the work force. For the 1990-2001 period, the employed rural work force contracted by 19% (2 million workers) while the urban growth rate rose 40%. However growth rate of the metropolitan and non-metropolitan areas was a very similar 25% and 29% respectively reflecting far small reallocations. Even under the extreme and unsupportable assumption that those 2 million rural workers who moved to urban areas were informal and had migrated exclusively to metropolitan areas (which are a subset of urban areas), they could explain at most half of the 4 million observed increase of the informal. Further, this would imply no increase in the number of informal workers in the non-metropolitan areas which actually rose from 20 to 27 million as can be inferred from panels (a) and (c). In short, most of the rise in informality must be due to reallocation of labor within geographical sub regions.

1.6 Conclusions

This paper has sought to explain the dramatic rise in the size of the informal sector over the decade of the 1990s by studying the impact of trade and constitutional reforms on gross labor flows. We establish that trade liberalization played a relatively small part in this increase, but find suggestive evidence that several dimensions of the constitutional reform, in particular, regulations relating to firing costs, overtime, and union power, explain much more. Both effects work mostly through the reduction in hiring rates, rather than separation rates that have been investigated most in the literature to date. Overall, the findings confirm the importance of labor legislation to firms' decisions to create new formal sector jobs in Brazil.

1.7 Appendix

1.8 An Illustrative Model

As a way of organizing thinking on the impacts on gross flows of the innovations discussed above, we build a highly stylized search and matching model of interactions between the formal and informal sector. We assume that firms can hire workers under two production technologies. The first is suitable for relatively low productivity workers and is difficult to monitor by the government. Getting the most out of intrinsically more productive workers requires, however, a technology that is easier to monitor by the government and hence the firm must comply with all regulations and make

[23] See Ramos & Ferreira (2005b,a) for a comprehensive description of the regional patterns of the Brazilian workforce.

[24] A Metropolitan Region is a group of specific limiting municipalities. By Constitutional prerogative, such a group is defined by each specific State of the country with the purpose of improving the planning and the execution of public functions of common interest. As a point of reference, according to PNAD, about 97% of the population and of the employed workforce older than 15 years living in Metropolitan Regions is classified as Urban.

[25] Dwellers are classified either as Urban or Rural according to the geographical location of their residences. A residence is classified as Urban if it is located on a city (municipal level), on a village (distrital level) or in isolated urban areas. A residence is classified as Rural if it is located outside the aforementioned locations.

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the worker formal. We might think of subcontracting jobs that are able to be done at home, or that are simply not very visible compared to work that involves specialized machinery and a fixed location or plant. The model focuses on the hiring and firing decision of firms, as they are confronted with high and lower productivity applicants and their decision at the margin between choosing formal or informal labor.

Let V be the present discounted value (PDV) for a firm of the expected profit from posting a vacancy. The total number of matches between firms and workers, m, is given by the matching technology $m = m(u, v)$, where u and v represent the number of unemployed workers and vacancies respectively. Firms find workers at an average rate $q(\theta) = m/v$, where θ is the vacancy to unemployment ratio (v/u). Similarly, workers find firms at an average rate of $\theta q(\theta) = m/u$

We assume that once the firm and the worker have met, the firm observes the true idiosyncratic productivity of the match, x, which is drawn randomly from a known c.d.f. $G(x)$. Given x, the firm decides to hire the worker formally or informally. Here we greatly simplify the decision by assuming that only formal workers can take advantage of the idiosyncratic productivity of the match. If the firm hires the worker formally the productivity of the match is given by $p_f x$ where p_f is an overall productivity parameter for formal jobs. If, instead the firm uses an informal contract the match produces p_i, independently of the value of x. This mechanism generates that, consistent with the data, the most productive matches give rise to formal contracts. $J_f(x)$ and J_i represent the PDV for the firm of occupied formal and informal jobs respectively. It is straightforward, then, to show that there is a reservation productivity x_f that will make the firm indifferent between hiring the worker formally or informally. Hence, we can write the flow value of a vacant job for the firm as

$$rV = -c + q(\theta)\left[\int_{x_f}^{x^{max}} (J_f(x) - V)dG(x) + G(x_f)(J_i - V)\right] \tag{1.9}$$

where r is the interest rate, c is the instantaneous cost of keeping the vacancy open. Similarly, the value for the firm of occupied formal and informal jobs can be expressed as

$$rJ_i = p_i - w_i - \lambda_i(J_i - V) \tag{1.10}$$

$$rJ_f(x) = p_f x - w_f - \eta h + \int_{x_f}^{x^{max}} (J_f(s) - J_f(x))dG(s) - \lambda_f G(x_d)(J_f(x) + F - V) \tag{1.11}$$

The production technology of informal workers is very simple. A worker hired produces p_i in an exchange for w_i. Exogenous shocks arrive to informal jobs at rate λ_i, at which point the match is destroyed.

Formal firms have to abide by labor codes, so that, on top of the wage w_f, they have to pay an overtime premium η, in case they have to work over the legal maximum hours per week. This excess in hours is represented by h, which is considered exogenous in the model. The last two terms of equation (1.11) capture the continuation value of the job. We assume that ongoing formal jobs are subject to idiosyncratic productivity shocks a la Mortensen & Pissarides (1994) that modify the productivity of the match. These shocks arrive at rate λ_f. However, in this case the jobs are not automatically destroyed. Upon the arrival of a shock, a new value for x is drawn from $G(x)$. The matches are only destroyed if the new productivity value renders the match unprofitable. Let this idiosyncratic productivity be x_d. If the new productivity is above x_d the match persists. If is below x_d the firm destroys the relationship paying firing costs F.

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The workers problem is similar of that of the firm. The present discounted value of unemployment, U, can be written as

$$rU = b + \theta q(\theta)\left[\int_{x_f}^{x^{max}} (W_f(x) - U)dG(x) + G(x_f)(W_i - U)\right] \tag{1.12}$$

where b is the flow of income when unemployed and $\theta q(\theta) = m(u,v)/u$ is the rate at which workers meet firms and $W_f(x)$ and W_i are the present discounted value for the worker of a formal and informal job respectively. Again, the contract will be formal if the idiosyncratic productivity is above x_f.

Once the contract is signed the worker enjoys $w_f(x)$ or w_i depending on whether the contract is formal or informal, until the job is destroyed endogenously for formal workers or exogenously for informal ones[26].

$$rW_i = w_i + \lambda_i(U - W_i) \tag{1.13}$$

$$rW_f(x) = w_f(x) + \int_{x_d}^{x^{max}} (W_f(s) - W_f(x))dG(s) - \lambda_f G(x_d)(W_f - U) \tag{1.14}$$

As is standard in the literature, wages in this model maximize the joint surplus and determine the following sharing rules for formal and informal jobs[27].

$$J_f(x) - V + F = \frac{1-\beta}{\beta}(W_f(x) - U) \tag{1.15}$$

$$J_i = \frac{1-\beta}{\beta}(W_i - U)$$

where β is the workers bargaining power.[28] Using this sharing rule and equations (1.9) to (1.14) we can obtain the two wage equations for formal and informal jobs respectively,

$$w_f(x) = (1-\beta)b + \beta p_f(x + \theta c + rF - \eta h) \tag{1.16}$$

$$w_i = (1-\beta)b + \beta p_i(1 + \theta c)$$

In this framework the equilibrium of the model must satisfy three conditions that determine our three endogenous variables; how many vacancies to post (which would determine θ) when to hire a formal worker (x_f) and when to fire formal worker (x_d). As usual in search and matching models, the free entry condition determines that there cannot be any profitable opportunities from vacancy posting, hence $V = 0$. Second, optimal hiring of formal workers must satisfy that the firm is indifferent between hiring the marginal worker formally or informally, that is $J_f(x_f) = J_i$. Finally, optimal

[26] Note, that neither overtime pay, ηh, or firing costs, F, accrue to the value of a formal job for the worker. This is due to the nature of wage negotiation in this framework. Since firm and worker share the surplus of the match any transfer between the two parties will not have an impact in equilibrium. This is standard in the literature, see Pissarides (2000).

[27] It can be argued that the initial bargaining rule should not consider firing cost in the threat point of the firm since they are still not operational. This would give rise to two different wage equations for formal workers. This variation leaves the results qualitatively unchanged.

[28] One could argue that the bargaining power of formal and informal workers, β, is different. This has no consequences in the model.

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firing of workers must satisfy that the reservation productivity x_d makes the formal job unprofitable. This happens when $J_f(x_d) + F = 0$.

Using equations (1.9) to (1.16) and our three equilibrium conditions we obtain the three equilibrium equations in the model:

Free entry

$$\frac{c}{q(\theta)} = (1-\beta)\left[\int_{x_f}^{x^{max}} (\frac{p_f\,(x-x_d)}{(r+\lambda_f)} - \frac{F}{(1-\beta)})dG(s) + G(x_f)(\frac{p_i - b - \frac{\beta}{(1-\beta)}\theta c}{r+\lambda_i})\right] \quad (1.17)$$

Optimal Hiring

$$p_f x_f = p_f x_d + \frac{(r+\lambda_f)}{(1-\beta)}F + \frac{(r+\lambda_f)}{(r+\lambda_i)}[p_i - b - \frac{\beta}{(1-\beta)}\theta c] \quad (1.18)$$

Optimal Firing

$$p_f x_d = b + \frac{\beta}{(1-\beta)}\theta c + \eta h - rF - \frac{\lambda_f}{(r+\lambda_f)}\int_{xd}^{x^{max}} p_f\,(s-x_d)\,dG(s) \quad (1.19)$$

It is straight forward to show that the equilibrium in this model exists and it is unique (see Appendix 1.8.1 for details).

With knowledge of θ, x_f and x_d we can derive the evolution stock of workers. Formal employment, $n+f$, is determined by the law of motion

$$\dot{n_f} = \theta q(\theta)(1-G(x_f))u - \lambda_f G(x_d)n_f \quad (1.20)$$

Similarly, informal employment follows the law of motion

$$\dot{n_i} = \theta q(\theta)G(x_f)u - \lambda_i n_i \quad (1.21)$$

Normalizing the labor force to one, unemployment is given by,

$$u = 1 - n_f - n_i \quad (1.22)$$

The share of informal employment, is given by $i = \frac{n_i}{n_f+n_i}$ which can also be written as $i = \frac{\gamma}{1+\gamma}$ where $\gamma = \frac{n_i}{n_f}$. From, equations (1.20) and (1.21) we obtain the steady state value of γ as,

$$\gamma = \frac{G(x_f)}{1-G(x_f)}\frac{\lambda_f G(x_d)}{\lambda_i} \quad (1.23)$$

Equation (1.23) states that the share of informal employment in the model is a composite of the relative inflows into informality $\frac{G(x_f)}{1-G(x_f)}$, and the relative outflows from informality $\frac{\lambda_i}{\lambda_f G(x_d)}$. These, along with the sectoral shares, are the dependent variables whose movements we seek to explain in this study.

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1.8.1 Equilibrium and Effect of Policies

The equilibrium of the model is determined by free entry, optimal hiring and optimal hiring conditions. These determine the three equilibrium equations in our model (equations (1.17) to (1.19)).

It is straightforward to show the existence and uniqueness of the equilibrium in this model. Higher θ, increases the left hand side of the free entry condition since more vacancies per unemployed increases waiting time for firms and hence the expected cost of posting a vacancy. Furthermore, it lowers the right hand side this same equation since the formal separation threshold in equilibrium (x_d) depends positively on θ. Note that in equilibrium, the formal/informal threshold, x_f, does not alter the expected profits from posting a vacancy due to in the virtue of the envelope condition it satisfies $J_f(x_f) = J_i$. Hence, there is a unique value of θ that satisfies equation the free entry condition.

Effects of policies

Trade liberalization: We model trade liberalization simply as a change in the relative productivity of the formal sector relative to the informal. On the one hand, lower barriers increase the competition that an industry, and reduces the wedge between formal and informal productivity (formal sector rents) within an industry (for a given p_i, p_f decreases). This shifts hiring towards informal labor (increase in x_f) and increases the threshold of separations in the formal sector (increase in x_d). Both effects generate a reduction in the share of formal employment formal sector. However, in a contrary effect, reducing tariffs and quotas also permits greater access to imported capital goods and other intermediate inputs that may increase relative formal sector productivity.

Firing Costs: Raising the costs of firing a worker enters into the overall cost calculation in a manner similar to that of the overtime pay and shifts hiring from formal to informal employment (higher x_f). However, in this case the increased cost of formal separation decreases the relative outflows from formal jobs (lower x_d). As has been observed in the literature this implies that increases in firing costs, by reducing both entry and exit, have ambiguous effects on formal employment (See Kugler 2004).

Total differentiation of equations (1.17) to (1.19) allows us to unambiguously sign the effect of firing costs on the three endogenous variables of the model as,

$$\frac{d\theta}{dF} = \frac{-(1-\beta)\theta q(\theta)\left[\frac{[1-G(x_f)]}{(r+\lambda_f)}\left(\frac{(r+\lambda_f)\lambda_f G(x_d)}{r+\lambda_f G(x_d)}\right)\right]}{c\xi(\theta) + (1-\beta)\theta q(\theta)\frac{[1-G(x_f)]}{(r+\lambda_f)}\frac{\partial p_f x_d}{\partial \theta} + G(x_f)\frac{\beta}{r+\lambda_i(1-\beta)}c} < 0$$

$$\frac{d(p_f x_d)}{dF} = -\frac{r(r+\lambda_f)}{[r+\lambda_f G(x_d)]} + \frac{(r+\lambda_f)\beta c}{[r+\lambda_f G(x_d)]\,(1-\beta)}\frac{d\theta}{dF} < 0$$

$$\frac{d(p_f x_f)}{dF} = \frac{d(p_f x_d)}{dF} + \frac{(r+\lambda_f)}{(1-\beta)} - \frac{(r+\lambda_f)}{(r+\lambda_i)}\frac{\beta c}{(1-\beta)}\frac{d\theta}{dF} > 0$$

where $\xi(\theta) = -\frac{\partial q(\theta)}{\partial \theta}\frac{\theta}{q(\theta)} > 0$

Overtime Pay: Our model suggests that that an increase in the overtime pay (η) will reduce demand for formal labor. This translates into a reduced formal hiring relative to informal hiring, higher x_f, and an increase in the relative separation from formal jobs, higher x_d. Both forces imply a reduction in the share of formal employment. Further, we argue that the impact will be greater in those industries where the use of overtime (prior to the reform) was greater, greater h in the model. Hence,

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we expect that industries with a higher share of their working hours above the post constitutional maximum hours a week would see the greatest impact.

As before, total differentiation with respect to η also gives unambiguous results,

$$\frac{d\theta}{d\eta} = \frac{-(1-\beta)\theta q(\theta)\left[\frac{[1-G(x_f)]}{(r+\lambda_f)}\frac{\partial(p_f x_d)}{\partial\eta}\right]}{c\xi(\theta) + (1-\beta)\theta q(\theta)\frac{[1-G(x_f)]}{(r+\lambda_f)}\frac{\partial p_f x_d}{\partial\theta} + G(x_f)\frac{\beta c}{r+\lambda_i(1-\beta)}} < 0$$

$$\frac{d(p_f x_d)}{d\eta} = \frac{h}{\left[\frac{r+\lambda_f G(x_d)}{(r+\lambda_f)}\right]}\frac{[c\xi(\theta) + G(x_f)\frac{\beta c}{r+\lambda_i(1-\beta)}]}{c\xi(\theta) + (1-\beta)\theta q(\theta)\frac{[1-G(x_f)]}{(r+\lambda_f)}\frac{\partial p_f x_d}{\partial\theta} + G(x_f)\frac{\beta c}{r+\lambda_i(1-\beta)}} > 0$$

$$\frac{d(p_f x_f)}{d\eta} = \frac{d(p_f x_d)}{d\eta} - \frac{\beta c}{(1-\beta)}\frac{d\theta}{d\eta} > 0$$

It is important to note that while our model captures the depressive effect of increased costs on formal labor demand there may a countervailing *numeraire* effect: Though total hours worked by formal workers may fall, the fewer hours that each employee may legally work implies that the number of workers may actually rise. Determining the net effect requires knowledge of the number of overtime hours, the cost imposed by the overtime legislation, and especially the elasticity of formal/informal labor demand, two out of three of which we do not know.

Unions: Finally, the degree of unionization may capture how the increased union power enhanced the bargaining position of workers and changed the incentives for firms to hire (and dismiss) formal workers. This effect is captured in the model by the parameter β. We can show that an increase in β, under some regularity conditions (that the elasticity of the matching function with respect to unemployment is equal to β), has a similar effect to an increase in overtime pay: lower formal hiring, higher formal firing and overall, a lower overall formal sector.

Total differentiation with respect to β gives,

$$\frac{d\theta}{d\beta} = -\frac{\frac{\theta}{1-\beta} + (1-\beta)\frac{\theta q(\theta)}{c}\left[\frac{[1-G(x_f)]\theta c}{(1-\beta)^2[r+\lambda_f G(x_d))]} + \frac{1}{(1-\beta)^2} + G(x_f)\frac{\frac{c\theta}{(1-\beta)^2}}{r+\lambda_i}\right]}{\xi(\theta) + (1-\beta)\frac{\theta q(\theta)}{c}[\frac{(r+\lambda_f)\beta c}{\lambda_f G(x_d))(1-\beta)} + G(x_f)\frac{\frac{\beta}{(1-\beta)}}{r+\lambda_i}]} < 0$$

$$\frac{d(p_f x_d)}{d\beta} = \xi(\theta)\frac{\partial(p_f x_d)}{\partial\beta} - \frac{\theta}{1-\beta}\frac{\partial(p_f x_d)}{\partial\theta} = \frac{\xi(\theta)-\beta}{(1-\beta)^2\frac{r+\lambda_f G(x_d))}{(r+\lambda_f)}\left[\xi(\theta) + (1-\beta)\frac{\theta q(\theta)}{c}[\frac{(r+\lambda_f)\beta c}{\lambda_f G(x_d))(1-\beta)} + G(x_f)\frac{\frac{\beta c}{(1-\beta)}}{r+\lambda_i}\right]}$$

$$\frac{d(p_f x_f)}{d\beta} = \frac{d(p_f x_d)}{d\beta} + \frac{(r+\lambda_f)}{(1-\beta)^2}F - \frac{(r+\lambda_f)}{(r+\lambda_i)}[\frac{c\theta}{(1-\beta)^2} + \frac{\beta}{(1-\beta)}\frac{d\theta}{d\beta}]$$

In this case, we cannot unambiguously sign the effect of unions on the hiring and firing margin. However it can be shown that assuming $\xi(\theta) = \beta$, as it is standard in the literature, we can prove that marginal changes in β do not change the firing margin but shift the hiring margin towards hiring informal workers.

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Chapter 2

Formality, Labor Productivity, Wage Setting and Income Risk[1]

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Abstract

One possible explanation for the modest correlation observed between wages and productivity in most of the existing related literature is that such lack of correspondence holds for workers in ongoing work contracts -the group with bigger participation in the labor workforce- but not for workers starting new labor relationships. Recent studies have shown that wages for this last group correlates to aggregate productivity with elasticities close to one. Our paper adds to this tradition with some distinctive aspects. First, it revisits the hypothesis of rigid wages (understood as barely responsive or non-responsive to productivity shocks) for job stayers and flexible wages for new hires (or job movers) in a LDC context, for the biggest Latin American economy. Second, it exploits finer measures of productivity coming from firm surveys rather than just traditional aggregates such as GDP per unit of aggregate labor. Third, it expands the conventional approach from the formal sector to the informal one. Fourth, it attempts for a variance decomposition of earnings for workers grouped according to several criteria (mobility, formality) as well as for a measure of the contribution of productivity shocks to income volatility. Our findings are consistent with those observed in the US economy and suggest that in the Brazilian manufacturing sector, wages elasticities are higher for newcomers than for incumbents and that such elasticities are close to one for new hires in the formal sector. We also find that transitory income shocks are dominant in workers earnings volatility with productivity shocks being fully insured by firms for formal workers and not for the informal ones.

2.1 Introduction

Recent contributions to the unemployment volatility puzzle literature have focused in the theoretical role of the marginal worker in the process of wage setting and in the empirical correlation between

[1]This study has benefitted from invaluable discussions with Mariano Bosch, William Maloney, Alan Manning and Massimiliano Marcellino. We thank to Marc-Andreas Muendler for sharing his data on productivity, to Barbara Petrongolo and LSE Labor-Seminar participants for helpful insights and to Luigi Guiso, Salvador Ortigueira and Moritz Meyer for their comments during the EUI 2010 Third Year Forum. The usual disclaimers apply.

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this wage and broad productivity measures. Before these contributions, the mainstream of the literature suggested poor empirical correlation between aggregate measures of wages and aggregate measures of productivity. Indeed, the traditional view that wages are sticky over the cycle[3] was commonly derived from aggregate time-series[4]. This convention has been recast recently. Significant efforts in the labor market literature have been devoted to appeal for some mechanism[5] able to improve the performance of labor market models with search frictions in order to match the business cycle facts that are found in the data (Haefke *et al.*, 2008; Pissarides, 2009; Carlsson *et al.*, 2009; Gertler & Trigari, 2009; Costain & Reiter, 2008; Menzio, 2005; Rudanko, 2008; Farmer, 2006; Moen & Rosen, 2006; Blanchard & Gali, 2008; Hall & Milgrom, 2008; Shimer, 2009). Most of these contributions stress the importance of marginal workers (or equivalently, workers transiting from unemployment into employment in the flows of job creation or, simply, the new hires) in the wage bargaining process. For instance, Pissarides (2009) shows that the job creation condition that drives the volatility of the job finding rate depends on the wage bargain in new jobs. Moreover, he claims that time-series or panel studies on the cyclical volatility of wages show considerable stickiness, but this evidence is dominated by wages in ongoing jobs and it is not relevant for job creation in the search and matching model. He also claims that an examination of panel data evidence on the volatility of wages in new jobs shows that volatility is about the same as in the Nash wage equation of the canonical search and matching model. In a related vein, Haefke *et al.* (2008) find that US data is consistent with the conventional argument that wages are rigid but just in ongoing jobs. More interesting, they also find that this is not longer the case when focusing on wages of newly hired workers or new matches. In fact, such wages, unlike the aggregate ones, are volatile and respond one-to-one to changes in labor productivity. We build mainly upon these two contributions.

Our study adds to this literature and makes a number of contributions. First, it shows consistency between the findings for the US market and those observed in the biggest Latin American labor market, where wages paid to newly hired formal workers (or to job movers) in the manufacturing sector are strongly correlated to the firms productivity of that sector. Second, it exploits finer measures of productivity -and hence it exploits cross sectional industrial variation- coming from firm surveys rather than just traditional aggregates such as global GDP per aggregate unit of labor. Third, it expands the scope of the seminal contributions by introducing the analysis of the informal labor sector which roughly represents half of the Brazilian labor force Bosch *et al.* (2009) and therefore merits attention. In addition, we decompose the wage volatility observed in different sectors (formal, informal, stayers, movers) into their permanent and transitory parts and we attempt for a quantitative approximation of the contribution of the productivity volatility to the determination of these components.

The remainder of this document is organized as follows. Section 2.2 describes the source of information used in this study, Section 2.3 explains the relevance of new hires and job to job transitions and its role to understand the wage and productivity volatility puzzle, Section 2.4 provides a brief summary of the Brazilian institutional framework for wage settings and a description of the estimation technique, Section 2.5 reports the empirical results, Section2.6 explains the methodology for the permanent and transitory decomposition and delivers the results, Section 4.4 concludes.

[3] Keynes-Tarshis-Dunlop controversy.
[4] Series which in turn pick up the dominating effect of less-volatile incumbents' earnings.
[5] usually consisting of wage rigidities

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2.2 Data

2.2.1 Labor allocation and wages

We draw from the Monthly Employment Survey (*Pesquisa Mensal de Emprego*, hereafter PME[6]) that conducts extensive monthly household interviews in 6 of the major metropolitan regions[7] covering roughly 25% of the national labor market. The questionnaire is extensive in its coverage of participation in the labor market, wages, hours worked, etc. that are traditionally found in such employment surveys. The PME is structured as a rotating panel, tracking each household during four consecutive months and then dropping them from the sample for 8 months, then reintroducing them for another 4 months. Each month one fourth of the sample is substituted with a new panel. Thus, after 4 months the whole initial sample has been rotated, after 8 months a third different sample is being surveyed, and after 12 months the initial sample is interviewed. Over a period of two years, three different panels of households are surveyed, and the process starts again with three new panels. Regrettably, the PME was drastically modified in 2002 and it is not possible to reconcile the new and old definitions for unemployment and job sectors. Hence, analysis done with this data can at best begin in 1983 and stop at 2002.

From this dataset we can also identify the formality or informality of the labor force. There is broad consensus in the literature on the definition of informality from a labor market perspective both in the mainstream and Brazilian literature. A comprehensive survey of the literature studying the size and evolution of the Brazilian informal sector in the labor market can be found in Ulyssea (2005) and a summary of stylized facts of this sector covering eighties and nineties is detailed in Ramos & Reis (1997); Ramos (2002); Ramos & Brito (2003); Veras (2004); Ramos & Ferreira (2005b,a); Bosch *et al.* (2007, 2009). We follow this literature in definition by dividing employed workers into three sectors: formal salaried (F)-public employees and workers whose contract is not registered in his/her work-card or *carteira de trabalho*[8] that entitle the worker to labor rights and benefits; informal salaried (I), without *carteira*; and informal self employed (S.E.). Ideally, following the ILO we would distinguish by firm size as well, focusing on establishments of small number of employees as informal employees, however the PME does not tabulate this information and hence, we rely purely on the basis of lack of signed *carteira*- as the critical distinguishing characteristic[9].

We match workers observed in four consecutive monthly datasets. This allows us to identify newly hired workers as those workers that were not employed for at least one of the three months before we observe their wages. In a similar fashion, we can also trace job to job movements (defined as displacements from one labor sector or industrial sector to another between periods). Non-mover workers are defined as those who are not classified as new hires or job to job movers. In addition, we have information on worker characteristics (gender, age, education, race, marital status, etc.), industry, occupation and labor sector so we can control for these covariates.

[6]For descriptions of the methodology of the PME, see Sedlacek *et al.* (1990); IBGE (1991); Oliveira (1999).

[7]São Paulo, Rio de Janeiro, Belo Horizonte, Porto Alegre, Recife and Salvador.

[8]According to the Brazilian legislation, registered workers are the ones whose labor contract is registered on their work-card. This registration entitles them to several wage and non-wage benefits such as 30 days of paid holiday per year, contribution for social security, right to request unemployment benefit in case of dismissal, monetary compensation if dismissed without a fair cause, maternity and paternity paid leave and so on.

[9]The ILO defines informality as consisting of all own-account workers (but excluding administrative workers, professionals and technicians), unpaid family workers, and employers and employees working in establishments with less than 5. In fact, Bosch & Maloney (2006) find that in Mexico, the ILO's criteria of small firm size and ours of lack of registration are similar in motivation conceptually and lead to a great deal of overlap. 75% of informal workers are found in firms of 10 or fewer workers. Since owners of firms or self-employed are not obliged to pay social security contributions for themselves, we in fact consider them as informal self-employed with no social security contributions (and hence without the benefits that are perceived by salaried workers holding a *carteira*).

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2.2.2 Firms Productivity

Our data of firms[10] comprises a random three-firm aggregate of the Annual Manufacturing Survey (*Pesquisa Industrial Anual* or *PIA*)[11]. This is because IBGE's (the Brazilian Statistical Bureau) publication rules allow data from PIA to be withdrawn in the form of tabulations of cells having at least three firms. The random three-firm aggregates are formed within a continuous sample presence pattern after excluding discontinuous spells[12]. A unit of observation is drawn then from each Industrial Sector [13], calendar year [14] and location (metropolitan São Paulo[15] or rural). If the total number of firms within such a group (sector-year-location) is not divisible by three, a single four- or five-firm group is formed.

In this dataset, firms output is defined as sales (excluding resales) plus inventory buildup. We also dispose of employment grouped in blue and white collar work (separately used for productivity estimation). Log total factor productivity comes from an estimation using the Olley-Pakes method at the individual firm level[16]. Briefly, if the conventional Total Factor Productivity (TFP) approximation by Griliches & Mairesse (1990) is given by:

$$lnTFP_{i,t}^{GM} = ln\frac{Y_{i,t} - M_{i,t}}{L_{i,t}} - \alpha ln\frac{K_{i,t} + S_{i,t}}{L_{i,t}}$$

where $Y_{i,t}$ represents the levels of output Y of firm i in period t, M represents the level of intermediate inputs, L represents the level of labor, K the level of equipment and α is a exogenously set parameter [17] and if the OLS estimate of the same object is given by:

$$lnTFP_{i,t}^{OLS} = y_{i,t} - (\hat{\beta}_K k_{i,t} + \hat{\beta}_S s_{i,t} + \hat{\beta}_M m_{i,t} + \hat{\beta}_L l_{i,t})$$

where variables not stated in capital letters represent the natural logarithm of the corresponding value in levels, then the TFP results from the Olley and Pakes estimation procedure takes this last expression and extends it by allowing to the managers of the firms to do an efficiency-choice adjustment.

Once log TFP is computed, log labor productivity is inferred by substracting the capital stock measures, weighted with their production-function coefficients, from log TFP at the firm level.

It is important to mention that we got access to both, the estimates of the firm-level log TFP series and log labor productivity series, but not to the intermediate inputs necessary for its computation. Nevertheless, provided that we got access to the measures of output and of labor, we also compute the ratio output/labor for each three-firm aggregate as an additional proxy for productivity.

Finally, we draw from *Instituto de Pesquisa Economica Aplicada* (IPEA) series of aggregate industrial productivity (ratio of real industrial production index and workers in the industrial sector) and of aggregate real GDP.

[10] We thank to Marc-Andreas Muendler for sharing the data of firm productivity.

[11] For a detailed description of data construction, see Menezes-Filho *et al.* (2008).

[12] As explained by Menezes-Filho *et al.* (2008), exclusion of discontinuous spells means that firms are retained starting with their first observation in the sample and through all subsequent years until a year of absence occurs for a firm. From the first year of absence on, the firm is kept out of the sample. This way, discontinuous spells are effectively precluded. The idea behind forming aggregates only within continuous sample presence patterns was to lump similar firms regarding entry and exit.

[13] Nivel 50 Sectors consist of 31 manufacturing sectors corresponding roughly to the two digit ISIC3 sectors

[14] The dataset covers the period 1990-1998. The calendar year 1991 is missing from the PIA sample because a federal austerity program suspended the survey in 1991.

[15] São Paulo state hosts roughly half of Brazil's manufacturing value added during the sample period.

[16] See Muendler (2004) for a detailed description of the method

[17] Usually $1/3$

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2.3 New hires, job stayers and job movers

Yet this study is mainly focused into the analysis of wages, it is still closely related to the behavior of allocations. This is not only by virtue of obvious equilibrium conditions [18] but also because the classification of workers according to their mobility group (new hires, job to job movers, incumbents) implicitly carries a process of allocations and reallocations within and between sectors.

Moreover, and due to the differentiated patterns one finds in the flows fueling the expansion (or contraction) of those sectors, it is possible to distinguish the relative importance of creation flows (new hires) in the determination of the size of the employment sector and not only in wages as this study is aimed at showing.

For instance, Bosch & Maloney (2007) explain "Shimer (2005a) -studying the US labor market- finds that there are substantial fluctuations in unemployed workers job finding probability at business cycle frequencies, while employed workers separation probability is comparatively acyclic. Such an observation is particularly true in the last two decades, period in which the separation probability has steadily declined despite two spikes in the unemployment rate. Thus, Shimer found that virtually all of the increase in unemployment and decrease in employment during the 1991 and 2001 recessions was a consequence of a reduction in the job finding probability."[19] Shimer (2005a) concludes that if one wants to understand fluctuations in unemployment, one must understand fluctuations in the transition rate from unemployment to employment, the *outs of unemployment*[20].

In this context, Figure 2.1 -borrowed from Bosch *et al.* (2007)-, depicts cyclical movements of Brazilian gross worker flows. As documented in Bosch *et al.* (2007), two top panels of Figure 2.1 show creation and destruction flows (transition probabilities) in the Brazilian labor market. In particular, destruction flows are understood as those taking workers from each of the employment sectors into unemployment (*ins of unemployment*). For all sectors, as found for the US by Blanchard & Diamond (1991); Hall (1995), flows into unemployment are clearly countercyclical and dramatically so during the 1983 and 1999 crisis. More interesting, the distinctive behavior of formal and informal figures become apparent already in the analysis of the flows: formal separations are relatively invariant while the informal show the largest volatility in separations. Similarly, creation flows (top panel) suggest a mirroring asymmetry: the job finding rate in the formal sector is highly procyclical and very volatile. However, the job finding rate in the informal sector although noisy is reasonably constant, including during the crisis. Finally, the bottom panel of Figure 2.1 shows the probabilities of transiting between formality and informality and it suggests pro-cyclical patterns of job allocation across all sectors of employment with movements that are highly correlated within pairs of bilateral flows.

More interesting and stemming from these findings is the observation that Bosch *et al.* (2007) make by adapting Shimer (2005a)'s simulation exercise[21]. They observe that it is the job creation the most influential flow in the determination of the sector size. This conclusion is also validated by parametrical results coming from experiments of different nature conducted by Bosch *et al.* (2009).

[18] Allocations are endogenously driven by workers compensations and exogenously affected by productivity shocks in the distinct formal/informal sectors (see Bosch *et al.* (2009)).

[19] Related empirical contributions focusing in Latin American markets can be found in Bosch & Maloney (2007) for Mexico and Paes de Barros & Corseuil (2004); Bosch *et al.* (2007) for Brazil.

[20] This conclusion was striking as it is in direct opposition to the conventional wisdom, built around research byDarby *et al.* (1985, 1986); Blanchard & Diamond (1991); Davis & Haltiwanger (1990); Darby *et al.* (1992), that recessions are periods characterized primarily by high job loss rates and opened a vast cascade of research stressing the importance of transition probabilities and worker flows.

[21] The exercise consists into simulating the size of the employment sector at steady state when it is reconstructed upon the combination of all the worker flows and checking the response of the size after shutting down one by one each flow.

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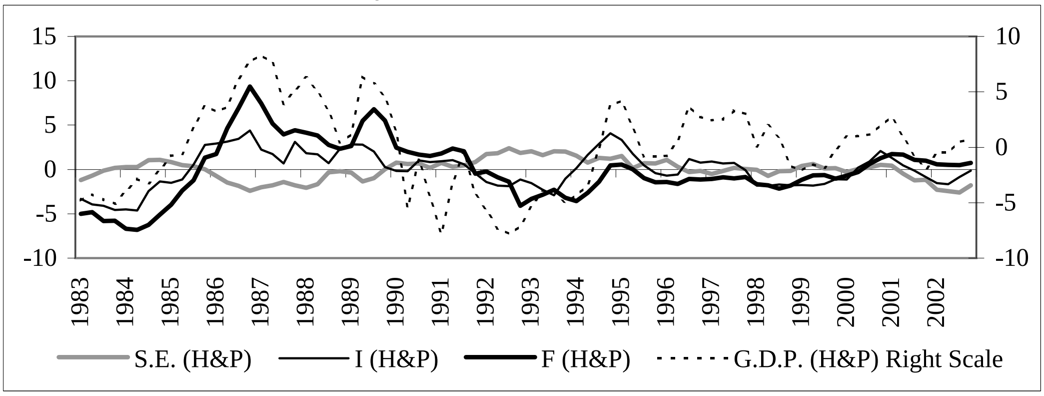
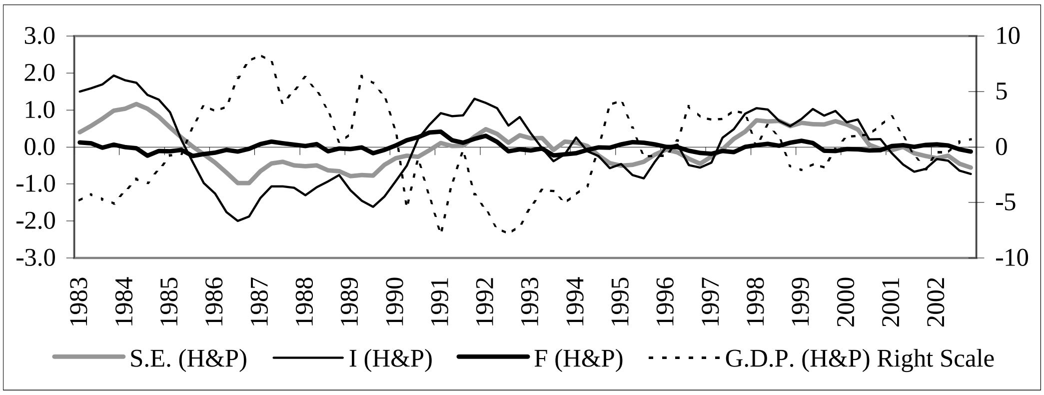
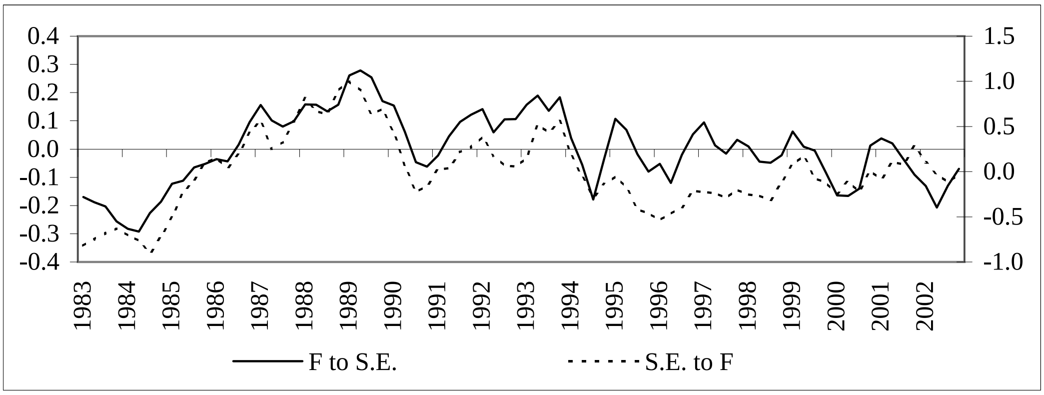
Figure 2.1: Gross Worker Flows



Creation



Destruction



Job to Job

Note: Transitions rates are inferred from a Markov transition matrix for each period using monthly data from the Monthly Labor Survey (PME) from Jan 1983 to Dec 2002 following the procedure by Geweke *et al.* (1986) as it is explained in Bosch *et al.* (2007). Computations are based on 10,000 Monte Carlo replications. The series have been averaged per quarter, de-trended using a HP filter with smoothing parameter 10^5 and smoothed using a 3 quarter moving average smoother. Depicted points correspond to the middle period of these rolling windows. Source: Bosch *et al.* (2007)

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All of these studies are focused into gross workers flows and labor reallocation, stressing the importance of job creation and job to job transitions but paying little or none attention on their effects over the behavior of wages. In this sense, this study is aimed at filling part of this gap by exploiting the particular dynamics of workers flowing from various employment categories (unemployment, formal salaried and informal salaried) as a distinctive characteristic to understand the differences in the volatility of their wages. Thus, we put special attention to the role of new hires (outs of unemployment) and job-to-job movers in the wage-productivity elasticity determination.

2.4 Wages and productivity

The asymmetric response of entering and outgoing flows of unemployment described in the previous section can help to explain -both theoretically and empirically- differentiated effects of productivity shocks to unemployment volatility[22]. Such an asymmetry can also help to distinguish the links between productivity volatility and wage volatility for workers with different mobility[23] and belonging to distinct labor sectors[24]. The theoretical challenge still needs to be addressed as part of what is known as the unemployment volatility puzzle[25] (Shimer, 2005b; Pissarides, 2009) and its implications over wage determination. This study contributes to this debate with an empirical exercise for the case of Brazil and leaves the theoretical work for future research.

To understand the context in which this exercise operates, we start by providing some institutional background.

2.4.1 Institutional Background

Brazilian wages -and in general, prices- were not free to be set according to real fundamentals during the immediate precedent years to those we analyze. In contrast, they were induced by considerable unionized bargaining power and then by binding nominal pressures. Carneiro (1999) documents and explains that, prior to 1964, collective wage adjustments were defined as part of a collective labor contract between workers and employers unions. Between the 1964 military coup and 1985, collective bargaining was replaced by a centralized state wage indexation which provided for a state minimum wage to be annually indexed based on the average real wage of the previous two years. However by the end of the 1970s, Brazilian inflation was growing rapidly and this provoked an upsurge in strike activities in the most organized sectors of the economy which in turn provoked (i) an immediate reduction in the period between salary adjustments from 12 months, until 1979, to six months thereafter; and (ii) the return of direct negotiations between unions and employers, which opened the door for wage adjustments above the past rate of inflation. The re-emergence of democracy in the 1980s combined with labor dissatisfaction with the state wage policy led to the development of a new unionism[26]. In the late 1980s and early 1990s, control of inflation became much more problematic in Brazil. One consequence of the generally unsuccessful counter-inflationary

[22]Related to this, Pissarides (2009) claims that the recent literature has either ignored the inflow rate when studying the cyclical dynamics of unemployment or treated shocks to the inflow rate as one of the exogenous forces driving changes in the outflow rate. But because more low-productivity jobs are destroyed in recession, at least some part of job separations is driven by endogenous decisions in response to aggregate productivity shocks. This entails a significant difference in the understanding of ins of unemployment (and its volatility) as if all job destruction were driven by exogenous separation shocks, the jobs destroyed in recession would be a random draw from the productivity distribution.

[23]job movers, new hires and job stayers

[24]Formal, Informal

[25]The observation that the response of unemployment to cyclical productivity shocks is bigger than implied by the canonical Search and Matching model).

[26]Trade unions slowly increased their role in the wage determination process and experienced an enhancement in their bargaining ability at both the regional and industrial levels. In terms of degree of bargaining centralization, an intermediate level of collective bargaining had emerged.

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policy shocks was an increased defensiveness on the part of labor unions, keen to protect real wage levels in the face of deflationary shocks. This contributed to the breakdown in coordination and synchronization of wage determination, and the development of a structure for collective bargaining much more akin to that prevailing in western Europe.

After the structural reforms implemented during early nineties and especially after the *Plan Real* (1994), Brazil entered into a phase of recovery and competitiveness in which fear to float and indexation were not longer the prevailing conditions and in which price settings (including wages) were subject to less distortions. Nonetheless, unionization still played a role. For instance, Carneiro (1999) finds that wages in the Brazilian industrial sector respond mainly to insider effects[27]. According to him, this is indicative that the Brazilian wage-setting structure is far from conforming to the traditional description of competitive models and appears rather conducive to rent sharing. In a similar vein, Carneiro & Henley (1998, 2001) examine wage determination in Brazilian manufacturing during the 1980s and early 1990s and show that the reduction in state regulation of collective bargaining led to the development of a system of wage determination which is increasingly characterized by rent sharing and insider trade union bargaining power. Even with by-now-outdated analytical tools they concluded that "real wages appear increasingly inflexible with respect to movements in open unemployment, with a large informal sector disciplining formal sector wage bargaining and cushioning the impact of broader labor market conditions."

Our work complements this analysis from a refreshed perspective by using updated estimation techniques (together with data of higher quality) and referring to a more suitable theoretical framework (stemming from the wage-productivity volatility puzzle and from the unemployment volatility puzzle).

2.4.2 Estimation Methodology

In essence, we exploit the dynamics of labor allocation in order to see how closely related the volatility of wages bargained between firms and new hired workers (or job to job movers) and the volatility observed in firms' productivity[28] are. Given that the volatility of wages can be driven by other factors besides factor productivity retribution (namely specific worker's characteristics) our analysis distinguishes between conditional and unconditional wages.

In this context, heterogeneity among workers can arise at least in two dimensions. In the individual dimension, heterogeneity exists because workers have different characteristics. In the aggregate dimension, heterogeneity can exist because wages for different groups are bargained under different schemes along the business cycle. For instance, newly hired workers signing formal contracts will bargain differently than those without signing a contract or than those with already long tenures. At the same time, workers with a formal status might have bargaining prerogatives that informal workers lack (or maybe informal workers might renegotiate their salaries more frequently). In addition to this, and given that our attention is mostly focused on the new hires and on the job movers, an additional source of heterogeneity bias stems from the fact that newly hired workers may not be representative of the whole labor force (see Figure 2.2[29]) or even worse that the composition of newly

[27] Insider variables are usually represented by different measures of productivity, union density, relative prices, and financial factors, such as liquidity and the firms ability to pay. Outsider variables, on the other hand, usually include alternative wages and measures of the state of the outside labor market, such as the level of unemployment.

[28] In particular, labor productivity.

[29] How important are non-stayer workers (those for who we contend that productivity measures should be better related)? Figure 2.2 sheds some light about this. It shows that sector non-stayers workers constitute a nontrivial group, more noticeable among informal workers (a half of them being sector movers and a third being new hires). It also shows that the composition of the labor force by transition groups is not constant across employment groups and across time with decreasing trends among incumbents in favor of increasing shares of job-to-job movers.

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hired workers varies over the business cycle, as is also pointed out by Haefke *et al.* (2008).

Such sources of heterogeneity would generate a bias in the estimate of wage cyclicality[30]. We follow Haefke *et al.* (2008)'s approach to take into account individual heterogeneity[31] and we cope with aggregate heterogeneity partly by distinguishing among mobility groups and labor sectors and partly by analyzing the wages after controlling by characteristics. Thus the wage w_{it}^j of an individual worker i of the group j at time t, depends in part on worker i individual characteristics and in part on a residual that may or may not depend on aggregate labor market conditions.

$$\ln w_{it}^j = \ln \hat{w}_{st}^j + x_{it}^{\prime j} \beta_t^j + \psi_{it}^j$$

where, x_{it}^j is a vector of individual characteristics (education, working experience, and their squared values), $ln\hat{w}_{st}^j$ is a vector of s industries fixed effects and ψ_{it}^j is the residual wage that is orthogonal to those characteristics. In other words, to obtain composition-bias corrected wages, we regress log wages on observable worker characteristics and take the average non stochastic component (fixed effects by industry) non attributable to workers characteristics. We call Conditioned and Unconditioned to the composition-bias corrected and uncorrected specifications respectively.

Using data described in Section 2.2, we construct monthly time series for the wage index $ln\hat{w}_{st}^j$. We do this for the j groups[32] and start comparing the cyclical properties of each series with those observed in aggregate productivity measures. Figures 2.3 and 2.4 depict the cyclical relation between productivity and unconditional (Figure 2.3) and conditional (Figure 2.4) wages for the aggregate manufacturing sector. A number of interesting patterns can be observed. First, unconditional figures are slightly countercyclical before 1994 and very acyclical afterwards[33] whereas the conditional series look quite procyclical during the whole period. Second, informal compensations are more volatile, especially in the conditional case. Third, new hires and job-to-job movers earnings are more volatile among formal workers while new-hires and incumbents earnings are more volatile among the informal ones.

Our interest remains, however, in how wages covary with *firms* annual productivity variables (and not just with *aggregate* manufacturing productivity series). Since the two main sources of information providing wages and firms productivity are not the same[34] we attempt to preserve as much cross sectional and intertemporal variation as possible. The best it can be done is to collapse information coming from both sources at yearly frequency and at industrial level of aggregation. Thus we aggregate monthly data on wages of workers at specific industrial sectors into yearly averages for the *j* different subgroups of workers. Then we collapse the yearly firms productivity values into industrial averages. An informal glimpse is provided by Figure 2.5 where we observe that firms' productivity distribution shares a significant portion of its domain with the distribution of incumbents' unconditional earnings. That is not the case for unconditional wages of new hires or job-to-job movers. After conditioning, the distribution of productivity shares a more significant part of its domain with that of the distribution of new hires earnings.

[30] Solon *et al.* (1994) claim that composition is an important source of cyclical aggregate wage variation.

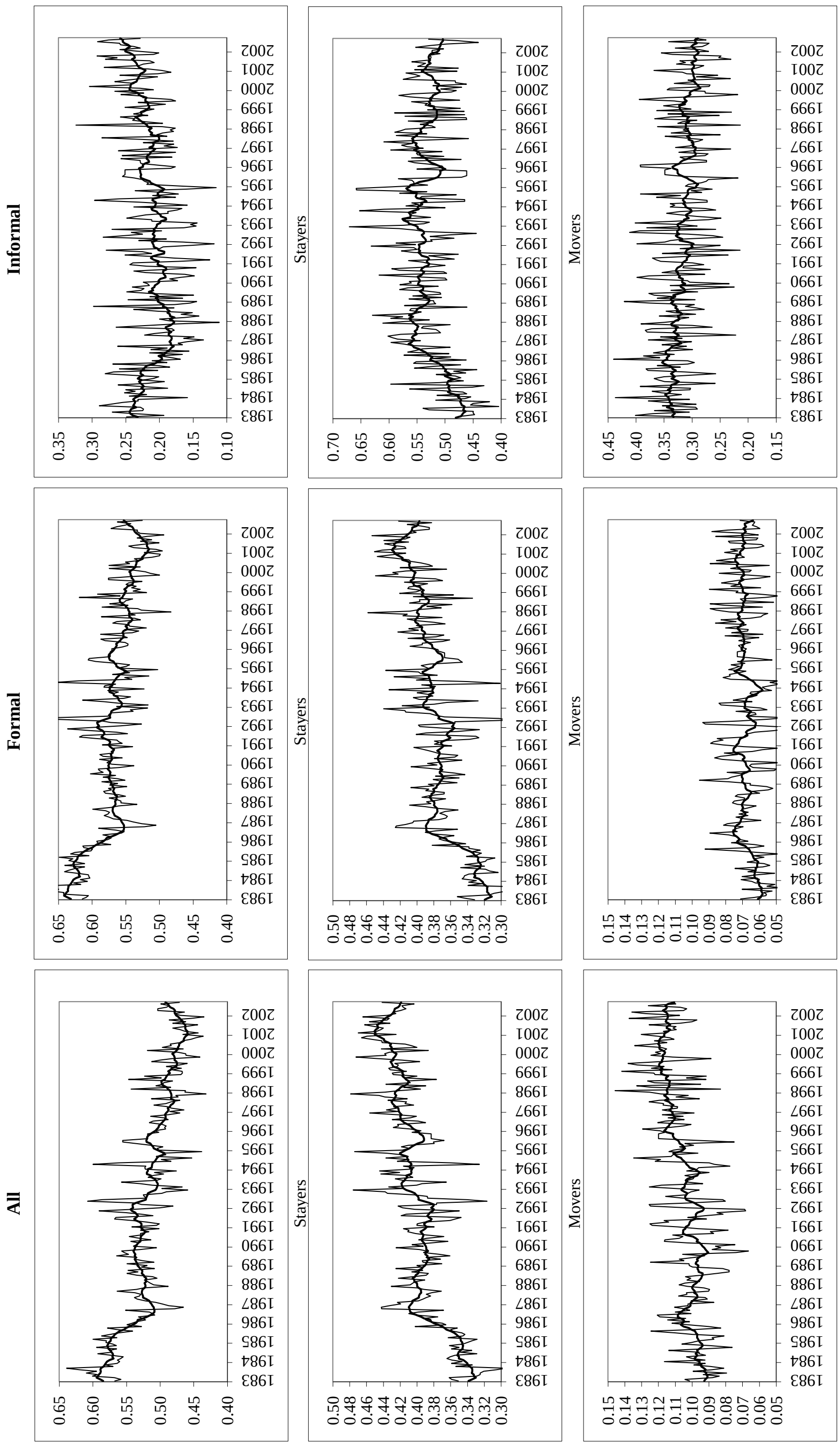
[31] On a related note, Menezes-Filho *et al.* (2008) find that Brazils high wage inequality cannot be explained by factors operating at the firm or industry level. Explanations must be sought in the characteristics of workers, both observable and unobservable.

[32] The subgroups belong to two dimensions: employment sector (All Salaried, Formal, Informal) and mobility group (Sector Stayers, Sector New Hires, Job to Job Sector Movers)

[33] A similar pattern of moderation during the nineties can be observed in the flows reported in Figure 2.1.

[34] In spite of the existence of a matched worker-firms dataset (Relação Anual de Informações Sociais or RAIS), regulations to disclose this information require coauthorship with a public officer of IPEA and a long process to get official authorization (as this dataset also identifies individuals and firms for taxation purposes). Thus the best subset of information we have access to is that of firms solely and of wages solely.

Figure 2.2: Labor Force Composition according to transition groups by Employment Sector



Note: The thick line represents the center value of a MA(12) for the actual shares (depicted with thin lines).

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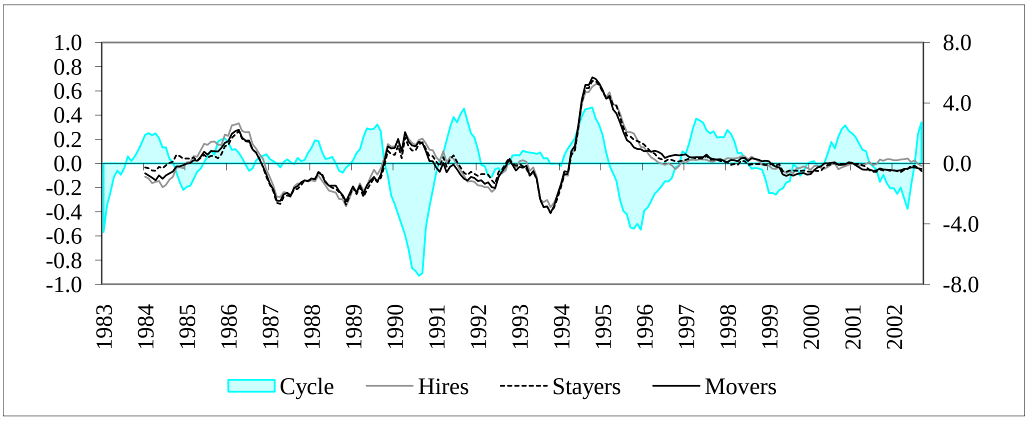
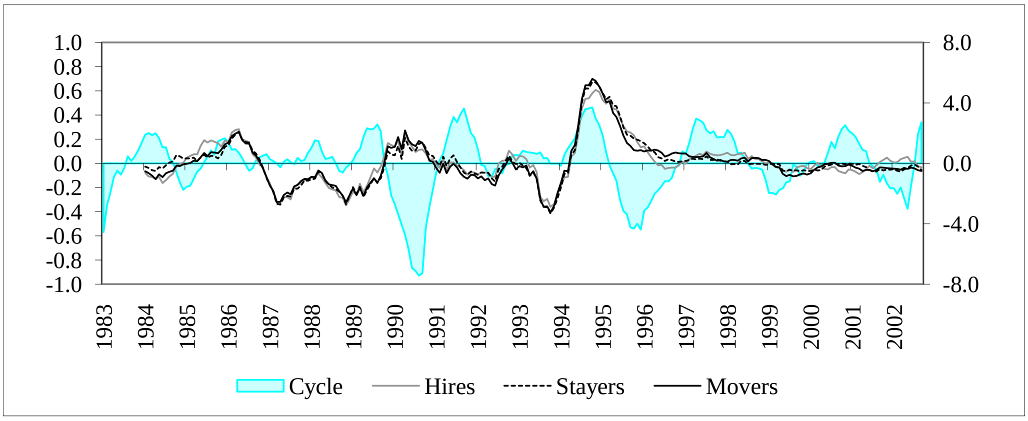
Figure 2.3: Cyclical relation between unconditional wages and aggregate manufacturing productivity

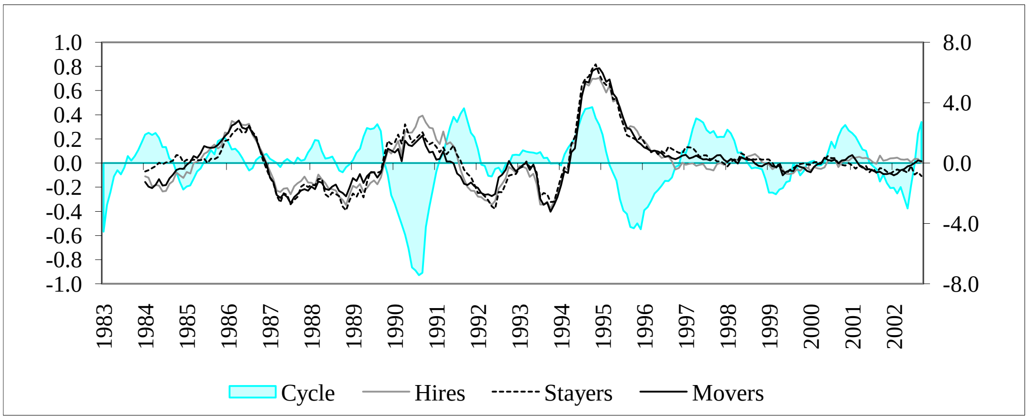
Unconditional



All



Formal



Informal

Note: The shaded area (right axis) depicts the aggregate manufacturing productivity series de-trended using a HP filter with smoothing parameter 14400. The lines (left axis) depict the yearly growth of MA(12) smoothed unconditional wages in the aggregate manufacturing sector.

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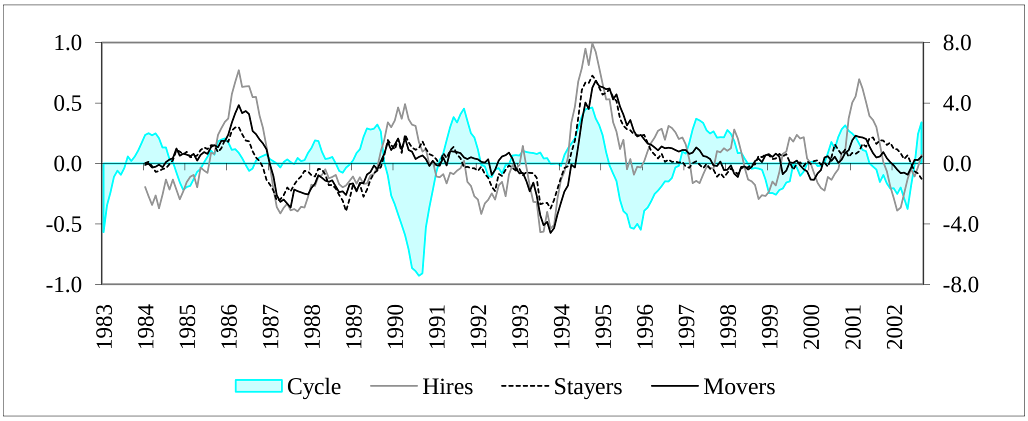
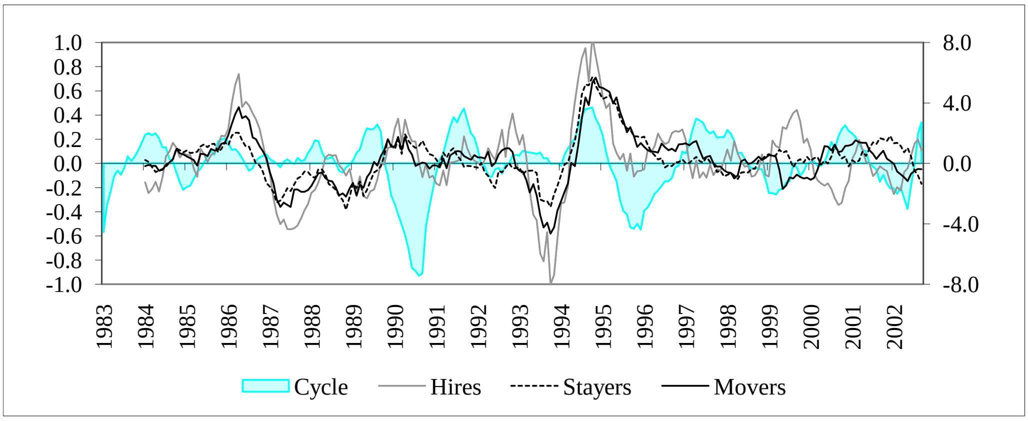
Figure 2.4: Cyclical relation between conditional wages and aggregate manufacturing productivity

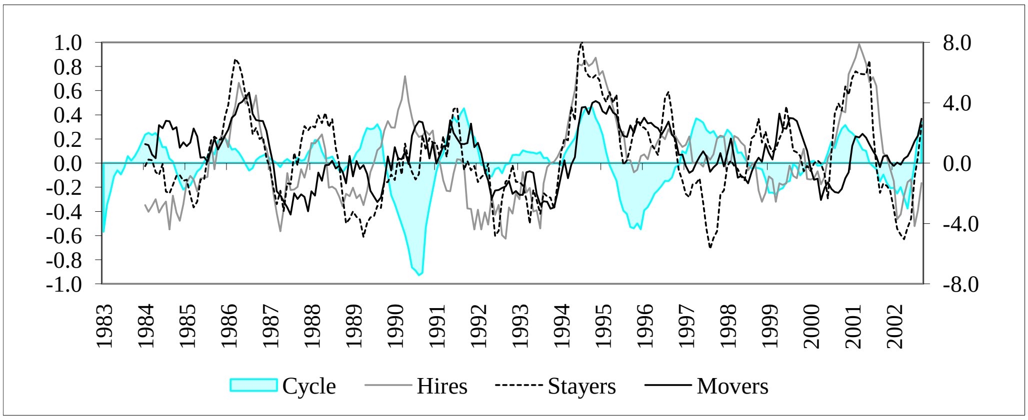
Conditional



All



Formal



Informal

Note: The shaded area (right axis) depicts the aggregate manufacturing productivity series de-trended using a HP filter with smoothing parameter 14400. The lines (left axis) depict the yearly growth of MA(12) smoothed conditional wages in the aggregate manufacturing sector.

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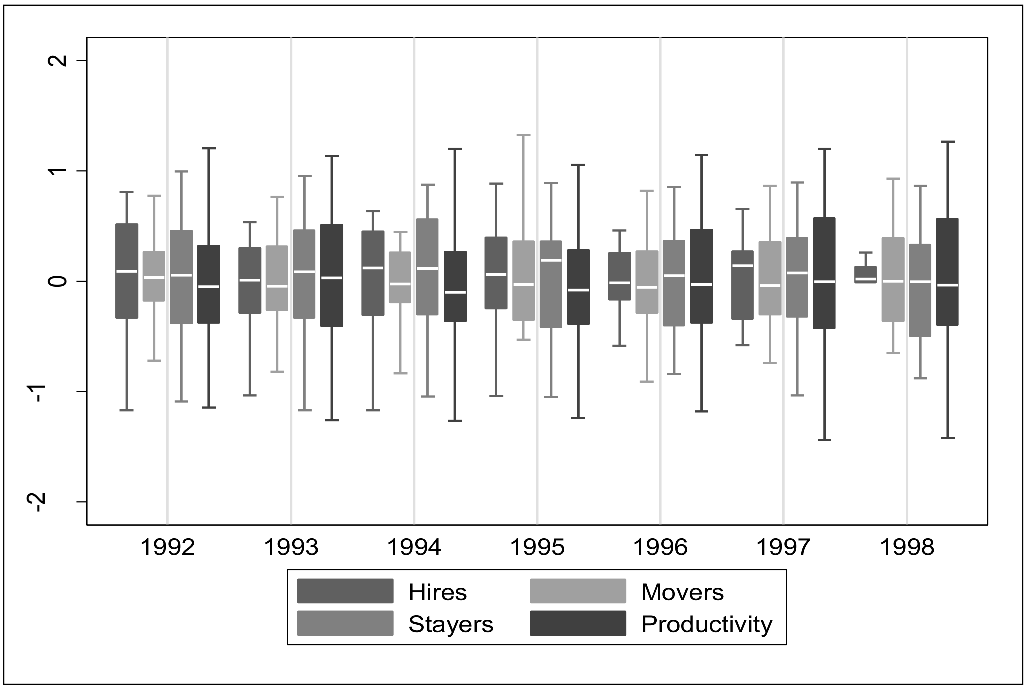
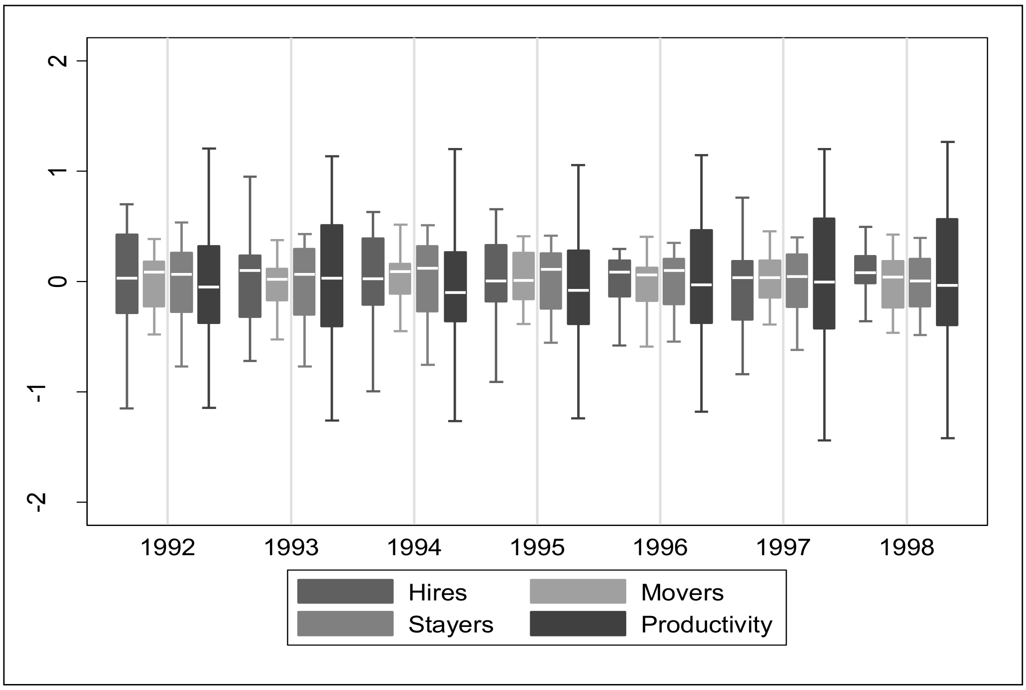
Figure 2.5: Wages Distribution's Support



Unconditional



Conditional

Note: The distributions represent dispersions across months within each year for aggregate measures of wages.

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To proceed formally, once the wage index and our indicator of firms productivity are measured in an homogenous frequency and for common observed units (manufacturing industries), we regress the change of log real wage index on the change of log real labor productivity. Among new hired workers, we expect to see the more flexible wages, this is, wages that would respond with the higher elasticities to changes in productivity. The opposite is expected among incumbents (higher stickiness or low elasticities)[35].

We run the following three specifications:

$$\ln \hat{w}^j_{st} = \alpha^j_s + \zeta^j \ln T\tilde{F}P_{st} + \epsilon^j_{st} \tag{2.1}$$

$$\ln \hat{w}^j_{st} = \alpha^j_s + \rho^j \ln \tilde{w}^j_{s(t-1)} + \zeta^j \ln T\tilde{F}P_{st} + \epsilon^j_{st} \tag{2.2}$$

$$\Delta \ln \hat{w}^j_{st} = \alpha^j_s + \zeta^j \tilde{\Delta} \ln TFP_{st} + \epsilon^j_{st} \tag{2.3}$$

where α^j_s and ϵ^j_{st} represent industry fixed effects and normal i.i.d. residuals respectively and $\tilde{X}$ denotes instrumented variable X[36]. Equation 2.1 represents a $log-log$ static model of wages against productivity which coefficient captures standard elasticities. Equation 2.2, our preferred specification, allows for dynamics and disentangles the effects of past earnings. Productivity coefficients in this case capture short run elasticities after inertia or stickiness in wages has been taken into account. Equation 2.3, is worked in differences in order to prevent spurious estimates of the elasticity if wages and productivity were integrated.

2.5 Results

Outcomes obtained from the estimation of Equations 2.1 to 2.3 are reported in Tables 2.2 and 2.3. Table 2.1 reports the results of a similar exercise for the baseline case in which microdata coming from firms surveys is disregarded[37]. Such results relies on aggregate industrial productivity measures[38] and correspond to those one should expect after accessing to common public sources of information.

Disregarding both, firms and workers heterogeneity, and abstracting from any kind of disaggregation, our preferred specification yields elasticities way below 1 and hovering around 0.54. After distinguishing between mobility groups, we already observe that the aggregate figure of 0.54 is closely resembled by that of the incumbent workers (0.55) whereas the new-hires group yields a higher value of 0.67. Nevertheless, these naive estimates are subject to potential biases stemming

[35] Elasticities for Movers are more difficult to predict given the absence of a theoretical model. One could expect an intermediate outcome as some movements may result from voluntary decisions and hence wage bargaining could be unrelated to the productivity observed in the absorbing sector (this is, one could expect wages of voluntary movers to be at least rigid to reductions) while some other movements may result from involuntary events that would put to the worker into a less favorable position to bargain a non-productivity-driven salary.

[36] In order to cope with potential endogeneity issues, internal instruments are used in all the cases. First lags are used as instruments for Equations 2.1 and 2.3. In the case of Equation 2.2 the instrumentation follows the Arellano and Bond system GMM estimation.

[37] In this case we abstract from Fixed Effects as aggregate data does not have cross sectional variation across industries. This also prevents the estimation of dynamic panels *a la* Arellano and Bond and hence the dynamic specification in the aggregate version includes in the Right Hand Side to the uninstrumented lag of the dependant variable and the instrumented productivity variable.

[38] See Section 2.2.2

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from the assumption of homogeneity, for both firms and workers.

Table 2.1 lifts the workers homogeneity constraint first (i.e. it keeps exploiting aggregate proxies of productivity instead of the actual estimates of firms labor productivity) and reveals a number of findings. First, Column labeled as *Unconditioned* suggests downward biased estimates when compared to those reported under the column of *Conditioned* wages. Like in Haefke *et al.* (2008) and Solon *et al.* (1994) we find strong evidence for composition bias because of worker heterogeneity. This negative heterogeneity bias prevails in every case throughout the Table in which the elasticity coefficient happens to be significant.

Second, a distinctive finding when using aggregate productivity measures is the high similarity between estimates for different mobility groups. This will change dramatically after refining the productivity measures and it will be likely reflective of the influence of firms idiosyncratic factors affecting differently to the productivity of specific firm at different industrial sectors. In fact, firms heterogeneity is not only a result after accounting for management or scale or technology or quality of the capital factor but also for the quality of its labor factor. Thus, workers heterogeneity should also play a role in this instance. Broader measures considering plain workers headcounts (such as product per worker) neglect differentiated complementarities between factors of different quality that should be translated into distinct productivity measures. Third, Panels A, B and C show the corresponding outcomes for All, Formal and Informal workers. It is not surprising to observe that Panels A and B keep much similarity as a significant portion of the manufacturing sector hires formal workers and aggregate measures of productivity are likely picking formal firms productivity[39]. Observed higher elasticities for Informal workers would suggest more flexibility of wages in this sector in case productivity measures were representative of informal firms but they are not. Hence we cannot make such a conclusion at this stage. Fourth, estimates obtained under the Dynamic specification are in between those high values coming from the Static estimation and those low and non-significant values coming from first differences[40].

Tables 2.2 and 2.3 discuss the values of the wage-productivity elasticities ζ observed for the different j groups when *conditional* wages and finer measures of productivity are taken into account.

Table 2.2 reports results corresponding to the estimation of Equations 2.1 to 2.3 for the whole sample of workers and for two alternative definitions of productivity. Columns (1) to (4) show results after considering Product per Worker (a broadly used proxy of productivity), but measured at a firm level rather than at the whole economy level. Given the limitations that this broad measure carries[41], another contribution of this study is the incorporation of a refined measure of productivity into the analysis: Columns (5) to (8) exploit our preferred definition of productivity, the Olley-Pakes productivity estimate computed by Muendler (2004)[42].

Similar to the case of aggregate productivity measures, we observe dynamic short run elasticities values in between those obtained under the static and first difference specifications. Yet we do not report the results of the unconditional estimates they also lie below those obtained after controlling for workers heterogeneity in most of the cases. In contrast to the previous case, recognition of firms

[39] Notice here that Panels A to C are distinguishing formality for the Left Hand Side only. This is refined later and is done here as a first approximation using aggregate measures of productivity.

[40] The dynamic specification remains as preferred as the Static one most likely contains unit roots and the first-differences are possible affected by over differentiation

[41] A measure of output per product does not fully recognize the effects of factor complementarity nor the effects of factor concentration. This is, it varies as a function of both other input factors and the efficiency with which the factors of production are used. Such other factors include -but are not limited to- physical capital and managerial activities, which in turn our actual labor productivity measure takes into account.

[42] See Section 2.2.2.

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Table 2.1: Wage-Productivity Elasticity. Conventional Aggregate Approach.

	Unconditioned				Conditioned			
	All (1)	New Hires (2)	Movers (3)	Stayers (4)	All (5)	New Hires (6)	Movers (7)	Stayers (8)
					A. All Salaried			
Static								
Elasticity	0.729***	0.852***	0.750***	0.750***	1.249***	1.118***	1.198***	1.270***
	[0.077]	[0.085]	[0.079]	[0.079]	[0.095]	[0.181]	[0.102]	[0.105]
R-squared	0.15	0.18	0.15	0.15	0.33	0.14	0.28	0.3
Dynamic								
Elasticity	0.543***	0.672***	0.564***	0.555***	0.935***	1.008***	0.937***	1.015***
	[0.089]	[0.101]	[0.088]	[0.093]	[0.130]	[0.184]	[0.128]	[0.135]
R-squared	0.24	0.24	0.24	0.24	0.39	0.14	0.33	0.35
In First Differences								
Elasticity	0.205	0.075	0.197	0.236	0.46	0.129	0.659	0.425
	[0.519]	[0.527]	[0.549]	[0.509]	[0.533]	[0.938]	[0.574]	[0.544]
					B. Formal			
Static								
Elasticity	0.774***	0.894***	0.752***	0.791***	1.213***	0.982***	1.119***	1.286***
	[0.078]	[0.087]	[0.078]	[0.079]	[0.093]	[0.204]	[0.103]	[0.105]
R-squared	0.17	0.2	0.15	0.17	0.33	0.08	0.26	0.31
Dynamic								
Elasticity	0.573***	0.676***	0.564***	0.587***	0.954***	0.985***	0.939***	1.045***
	[0.091]	[0.104]	[0.088]	[0.094]	[0.123]	[0.209]	[0.125]	[0.136]
R-squared	0.25	0.28	0.24	0.25	0.38	0.08	0.29	0.35
In First Differences								
Elasticity	0.23	0.156	0.247	0.225	0.541	0.9	0.867	0.321
	[0.516]	[0.494]	[0.550]	[0.508]	[0.532]	[1.155]	[0.578]	[0.543]
					C. Informal			
Static								
Elasticity	1.109***	1.059***	1.056***	1.297***	1.672***	1.642***	1.569***	1.415***
	[0.085]	[0.094]	[0.088]	[0.094]	[0.144]	[0.281]	[0.176]	[0.289]
R-squared	0.27	0.24	0.22	0.32	0.31	0.16	0.24	0.09
Dynamic								
Elasticity	0.826***	0.884***	0.795***	0.944***	1.293***	1.395***	1.365***	1.259***
	[0.116]	[0.117]	[0.112]	[0.132]	[0.183]	[0.293]	[0.197]	[0.306]
R-squared	0.35	0.28	0.31	0.39	0.35	0.17	0.26	0.1
In First Differences								
Elasticity	0.227	-0.136	0.094	0.772	-0.107	-0.153	-0.073	2.405**
	[0.556]	[0.617]	[0.564]	[0.597]	[0.698]	[1.194]	[1.047]	[1.129]

Notes:
225 monthly observations are included in the Static and Dynamic specifications and 224 in the specification in First (Annual) Differences.
Instrumental Variable Estimation uses as internal instrument the twelfth lag of the independent variable for the Static and Dynamic specifications. In the case of the model in First (Annual) Differences, the instrument is the first-month lag of the annual difference of the log-productivity.
Robust standard errors in brackets.
* significant at 10%; ** significant at 5%; *** significant at 1%.

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heterogeneity brings in differentiated conditioned elasticities according to workers mobility group. In particular, and paying attention to our preferred specification (Panel B) and to our preferred definition of productivity (Columns 5 to 8)[43], expected patterns emerge: a close resemblance of elasticities computed for all workers to those computed for stayers at either low or non significant values, highest (close to 0.7) and significant elasticities for new hires and also considerable (0.4) and significant elasticities for movers[44].

Results related to wages' inertia also merit attention. For instance, and consistent to the nature of the mobility groups, the memory of the process determining wage variations is nil and non-significant for new hires, significant and positive for movers and above 0.7 and significant for stayers.

Most of these findings give empirical grounds to the contention that in the Brazilian manufacturing labor market the less mobile the workers are, the more sticky their wages are and -also consistent with this- the lower the contribution of productivity to explain their wage variation.

These results add to the rent-sharing strand[45] with consistent findings to those obtained by Haefke *et al.* (2008) for the United States economy. Carlsson *et al.* (2009) also get results in favor to this strand for the Swedish case but without differentiated effects between mobility groups. As in Carneiro (1999) they stress the importance of insider effects yet they argue these are equally important to any kind of worker (entrant or incumbent). One reason that may drive the different outcome of Haefke *et al.* (2008) and Carlsson *et al.* (2009) exercises is the level of disaggregation of the data they exploit for the productivity variable. While the former exploits aggregate measures of productivity, the later got access to workers-firms matched data. In such a way Carlsson *et al.* (2009) can control for firms fixed effects besides workers unobserved heterogeneity. As they note, recent contributions have found mixed results in a related vein: Gertler & Trigari (2009), find that once one looks at equivalent workers within the same firm, there are no observable differences between incumbents and new hires in the response of wages to the aggregate unemployment rate in US data for the period 1990 to 1996. Carneiro *et al.* (2009) also controlling for firm and individual fixed effects find a higher elasticity of wages to the aggregate unemployment rate for new hires than for incumbents in Portugal for the period 1986 to 2005. However, Carneiro *et al.* (2009) also find no significant differences in the wage productivity elasticities across the two groups.

In our case and as explained before, sources of information for workers and firms are not matched even when both register information at an individual level. Hence, yet we cannot compare workers within firms, we can account for firms characteristics as it is implied in our productivity definition.

In these lines, another dimension in which firms characteristics allow a refinement of our analysis is firm's labor informality[46]. In particular even if it is not possible to measure the share of workers

[43] The limitations of the broad measure of productivity may be contaminating the results presented in Columns (1) to (4) and, for instance, probably driving the negative elasticity found in Column (4) of Panel B. Thus we present these results for expositional purposes to make the point that while results based on broad measures may seem at odds, refined measures or productivity yield better outcomes and reveal a coherent and consistent story.

[44] Elasticity is close to 1 for new hires and close to 0.6 for movers using the production per worker definition. In spite of this, negative and significant elasticities for incumbents and overall are found. This is most likely due to the fact that even after mitigating some of the imprecision that this ratio brings in for the aggregate case -deduced from the improvement of the elasticities for the different mobility groups- it still ignores the variation induced through factor complementarities that the Solow residual adjusted *a la* Olley-Pakes does take into account.

[45] As already suggested in section 2.4.1, two contending traditions can be referred as theoretical frames to study wage-productivity elasticites. In the competitive model, wages depend on labor market conditions and not on the particular situation of the firm or its sector of operation. On the other hand, in rent-sharing models wages depend on firms profitability and movements in demand and productivity.

[46] Provided the evidence in favor of higher and differentiated volatility of conditional wages across workers of different

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Table 2.2: Wage-Productivity Elasticity. All Salaried Workers and All Firms.

	Output per Worker				Labor Productivity			
	All (1)	New Hires (2)	Movers (3)	Stayers (4)	All (5)	New Hires (6)	Movers (7)	Stayers (8)
A. Static								
Elasticity	0.986***	1.221***	1.229***	0.696***	0.489***	0.550***	0.580***	0.348***
	[0.106]	[0.166]	[0.122]	[0.120]	[0.055]	[0.090]	[0.062]	[0.062]
R-squared	0.57	0.35	0.52	0.49	0.52	0.26	0.43	0.48
Observations	108	106	108	108	108	106	108	108
B. Dynamic								
Memory	0.923***	0.137	0.402***	0.905***	0.674***	0.105*	0.512***	0.745***
	[0.061]	[0.087]	[0.125]	[0.053]	[0.037]	[0.055]	[0.077]	[0.042]
Elasticity	-0.254**	1.047***	0.678**	-0.445***	0.194***	0.634***	0.382***	0.038
	[0.108]	[0.217]	[0.269]	[0.090]	[0.033]	[0.103]	[0.071]	[0.043]
Hansen	0.59	0.72	0.62	0.64	0.72	0.82	0.73	0.73
AR(2)	0	0.04	0.05	0.37	0	0.12	0.18	0.08
Observations	108	105	108	108	144	141	144	144
C. In First Differences								
Elasticity	2.356**	1.599	2.908	2.256*	0.874*	0.739	1.161**	0.915
	[1.182]	[1.429]	[1.918]	[1.220]	[0.486]	[0.708]	[0.584]	[0.559]
Observations	90	87	90	90	90	87	90	90

Notes:
Wages are conditional on workers characteristics.
Robust standard errors in brackets. Hansen and AR(2) p-values reported for the dynamic panel.
* significant at 10%; ** significant at 5%; *** significant at 1%.

within firms under informal arrangements, it is possible to proxy this -at a firm level- using two indicators: firms' size and the share of white-collar workers. ILO's definition of informality as well as regional evidence[47] relates informality to smaller firms. Likewise, same empirical evidence finds a close relation between informal labor and workers qualifications. Given the high similarity of the outcomes obtained when using the share of (blue)white-collar workers as the variable to approximate firm's (in)formality we just report results stemming from the use of firm's size.

Table 2.1 already advanced some discussion about the implication of informality in this context. Table 2.3 report results of a similar exercise after constraining the sample of firms. This is done under the suspicion that smaller firms (size being determined by the number of workers) are more keen to hire informally. We trim the firms sample dropping first to firms belonging to percentiles 1

labor sectors -as it is shown in Figure 2.2- we decompose the labor force according to formality sectors.

[47] Perry *et al.* (2007) show that 40% of the Brazilian informal labor force works in firms with 5 or less employees and 60% in firms with 10 or less workers.

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to 5[48] and second to firms belonging to percentiles 6 to 10. In such a way we are aimed at checking the sensitivity of the results after indirectly controlling for degrees of *firms'* formality.

Table 2.3 focuses on our preferred specification and reports a number of interesting findings. First, results considering Formal workers exclusively (Columns 1 to 4) for all firms (Panel A) are quite similar to those reported in Table 2.2. This is consistent with the fact that eventhough the high degree of informality in the Brazilian labor market, the manufacturing sector is still significantly populated by formal workers. Second, overall elasticity and elasticities for incumbent formal workers get closer while elasticity for New-hires goes up by about 50 percent getting close to 0.9. Third, earnings persistence keeps previous patterns: very high for stayers, moderate for Movers and only significant at 10 percent for New-Hires. Fourth, Formal New Hires start to show elasticities with increasing values (up to 1) after small firms (informal absorbers) have been put aside. Fifth, long run elasticities keep the same patterns after combining the effects of short run elasticities and the degree of income persistence. Columns (1) to (4) of Panel C relates formal workers earnings to productivity of firms of size above the first decile and thus these estimates should be preferred when analyzing the formal sector. Fifth, yet we are aware of the limitations of the data to identify the informal firms subspace, Panel A reports estimates for the sample containing all firms, including the smallest ones[49]. Going from Panel C to Panel A, one observes slight changes.

Overall, two messages emerge from the results obtained for Informal workers and All Firms: there is less variation across elasticities according to workers' mobility groups (elasticities goes from 0.6 for stayers to 0.8 for new hires) and earnings persistence is much lower or non significant as compared to those observed in the Formal case. Both features are consistent with the nature of Informal contracts, where long tenures and wage rigidities are unlikely, regardless the worker mobility group.

2.6 Income Risk

By using standard conditional and unconditional elasticities, previous section has documented the direct contribution of productivity volatility to income volatility for different groups of workers. Intrinsic to the study of income volatility is the notion of income risk and thus we expand our study to see what are the permanent and transitory effects of such productivity shocks into income risk for each type of worker.

To this end, we first define the income process as in Meghir & Pistaferri (2004) and then we compute the values of the variance of the permanent and transitory shocks using the Carroll & Samwick (1997) projection methodology.

As in Section 2.4.2 we posit the following Mincerian model for the conditional mean of log earnings:

$$lnw_{it}^{j} = x_{it}^{\prime j}\beta_t^j + \psi_{it}^j \tag{2.4}$$

where we use the same notation as before and $x_{it}^{\prime j}$ contains the same set of characteristics as before. In this case j does not include to the group on new hired workers as our observed unit is now an individual per month rather than an industry by year and hence we cannot follow the wages of newly hired before hiring. We also disregard time or industry fixed effects as we run the model for each available section and because we incorporate yearly industrial productivity later.

[48] The percentiles allude to the distribution of firms sorted according to their size.

[49] One cannot constraint the sample to only those firms belonging to the first percentiles as the loss of degrees of freedoms results dramatic.

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Table 2.3: Wage-Productivity Elasticity by Labor Sector and Firm Size.

	Formal				Informal			
	All	New Hires	Movers	Stayers	All	New Hires	Movers	Stayers
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
				A. All Firms				
Memory	0.587***	-0.276*	0.504***	0.771***	0.289***	-0.03	0.152	-0.194***
	[0.039]	[0.146]	[0.041]	[0.037]	[0.043]	[0.133]	[0.094]	[0.070]
Elasticity	0.103***	0.904***	0.229***	-0.004	0.610***	0.824***	0.704***	0.562***
	[0.036]	[0.146]	[0.048]	[0.036]	[0.064]	[0.139]	[0.084]	[0.202]
Long run Elasticity	0.25	0.71	0.46	0.00	0.86	0.80	0.83	0.47
Hansen	0.57	0.88	0.54	0.73	0.82	0.88	0.88	0.92
AR(2)	0.59	0.72	0.48	0.21	0.21	0.22	0.4	0.23
				B. Firms > 0.5th Decile				
Memory	0.530***	-0.371***	0.446***	0.778***	0.240***	-0.046	0.097	-0.203***
	[0.059]	[0.135]	[0.047]	[0.038]	[0.047]	[0.142]	[0.084]	[0.073]
Elasticity	0.151***	1.007***	0.279***	-0.012	0.656***	0.846***	0.759***	0.564***
	[0.056]	[0.173]	[0.057]	[0.037]	[0.082]	[0.164]	[0.072]	[0.205]
Long run Elasticity	0.32	0.74	0.50	-0.05	0.86	0.81	0.84	0.47
Hansen	0.60	0.84	0.55	0.73	0.81	0.86	0.86	0.84
AR(2)	0.38	0.97	0.67	0.21	0.24	0.33	0.52	0.22
				C. Firms > 1st Decile				
Memory	0.528***	-0.387***	0.403***	0.822***	0.235***	-0.034	0.081	-0.205***
	[0.074]	[0.134]	[0.048]	[0.036]	[0.051]	[0.145]	[0.086]	[0.075]
Elasticity	0.142**	0.981***	0.297***	-0.057*	0.621***	0.761***	0.727***	0.620***
	[0.062]	[0.144]	[0.053]	[0.033]	[0.074]	[0.158]	[0.072]	[0.200]
Long run Elasticity	0.30	0.71	0.50	-0.28	0.81	0.74	0.79	0.52
Hansen	0.60	0.89	0.54	0.73	0.81	0.86	0.87	0.93
AR(2)	0.50	0.41	0.66	0.23	0.39	0.39	0.72	0.2
Observations	108	141	108	144	136	131	136	124

Notes:
Wages are conditioned on workers' characteristics and Productivity is the Olley-Pakes adjusted productivity described in Section 2.2.2.
Results correspond to the estimation of Equation 2.2. Columns (1) and (3) consider the first lag of productivity as additional internal instrument.
Robust standard errors in brackets. Hansen and AR(2) p-values reported.
* significant at 10%; ** significant at 5%; *** significant at 1%.

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We assume that ψ_{it}^j, the residual current income after controlling by characteristics, can be decomposed into a martingale permanent component p_{it}^j and a transitory innovation with low persistence ξ_{it}^j.

$$\psi_{it}^j = p_{it}^j + \xi_{it}^j$$

where the permanent conditioned income p_{it}^j is assumed to follow a random walk:

$$p_{it}^j = p_{i,t-1}^j + \eta_{it}^j$$

Defining a d-period income difference as

$$\begin{aligned} \Delta^d \psi_{it}^j &= \psi_{i,t+d}^j - \psi_{i,t}^j \\ &= p_{i,t+d}^j + \xi_{i,t+d}^j - p_{it}^j - \xi_{it}^j \end{aligned}$$

recursive substitution yields

$$\Delta^d \psi_{it}^j = \{\eta_{i,t+1}^j + \eta_{i,t+2}^j + ... + \eta_{i,t+d}^j\} + \xi_{i,t+d}^j - \xi_{it}^j$$

Under the assumption that the errors ξ^j and η^j are white noise and uncorrelated with each other at all leads and lags, d-period variance can be derived as

$$Var(\Delta^d \psi_{it}^j) = d_{it}\sigma_{\eta^j}^2 + 2_{it}\sigma_{\xi^j}^2$$

where $\sigma_{\eta^j}^2$ and $\sigma_{\xi^j}^2$ are the variances of the permanent and transitory shocks to income respectively and $Var(\Delta^d \psi_{it}^j)$ can be estimated for each individual by

$$\hat{Var}(\Delta^d \psi_{it}^j) = (\Delta^d \psi_{it}^j)^2 = Var(\Delta^d \psi_{it}^j) + \mu_{it}^{dj}$$

with μ_{it}^{dj} is a mean-zero disturbance.

Thus the specification to be estimated is:

$$\hat{Var}(\Delta^d \psi_{it}^j) = d_{it}\sigma_{\eta^j}^2 + 2_{it}\sigma_{\xi^j}^2 + \mu_{it}^{dj} \tag{2.5}$$

where observations are distinguished by the length of the difference *d*. The projection method simply does OLS individual by individual of $\hat{Var}(\Delta^d \psi_{it}^j) = \{\hat{Var}(\Delta^1 \psi_{it}^j), ..., \hat{Var}(\Delta^n \psi_{it}^j)\}'$ on $[d_{it} \quad 2_{it}]$, where $d_{it} = \{1_{it}, ..., n_{it}\}'$, $2_{it} = \{2_{it}, ..., 2_{it}\}'$.

Given the frequency of our panel data, $\Delta^d \psi_{it}^j$ can represent income differences of $d = 1, d = 2, d = 3$ up to $d = n = 4$ months for each individual. The coefficients obtained for this regression give estimates s_η^2 and s_ξ^2 for $\sigma_{\eta^j}^2$ and $\sigma_{\xi^j}^2$ respectively.

Table 2.4 presents estimates corresponding to Equation 2.5 for the earnings of Formal and Informal workers for a number of exercises. First, we run the experiment for the entire period 1982-2002 (Columns 1 and 5). This is to have a measure of the variance of the permanent and transitory components exploiting as much information as available on the income dimension. Presumably, this measure yields the most consistent of our estimates. Second, we constrain the sample to those periods for which we dispose firm's productivity information but without including the volatility in productivity (Columns 2 and 6). This is to have benchmark measures and use them to quantify the impact after including the productivity shocks. Finally we incorporate the information on productivity[50] into the specification of Equation 2.4[51] and compute the corresponding estimates of Equation

[50] Starting from the information originally recorded at firm level we aggregate our different measures of productivity by industrial sector and year disregarding firm size. Controlling for firm size does not yield noticeable variations in the estimates discriminating by formality sectors and hence we decide to include all firms when computing these productivity aggregates.

[51] We include productivity and squared productivity measures as additional covariates.

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Table 2.4: Permanent and Transitory shocks by Labor Sector and Mobility Group

	Formal				Informal			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
				A. All workers				
s_η^2	0.016***	0.017***	0.016***	0.016***	0.029***	0.033***	0.033***	0.033***
	[0.001]	[0.001]	[0.001]	[0.001]	[0.004]	[0.005]	[0.005]	[0.005]
s_ε^2	0.086***	0.092***	0.091***	0.091***	0.136***	0.127***	0.124***	0.123***
	[0.001]	[0.001]	[0.001]	[0.001]	[0.004]	[0.006]	[0.006]	[0.006]
Observations	632713	202259	202259	202259	78086	27022	27022	27022
R-squared	0.124	0.164	0.165	0.165	0.138	0.158	0.161	0.161
				B. Stayers				
s_η^2	0.014***	0.016***	0.016***	0.016***	0.008	0.014**	0.013**	0.014***
	[0.001]	[0.002]	[0.002]	[0.002]	[0.005]	[0.006]	[0.006]	[0.005]
s_ε^2	0.077***	0.081***	0.081***	0.081***	0.111***	0.096***	0.090***	0.086***
	[0.001]	[0.002]	[0.002]	[0.002]	[0.005]	[0.007]	[0.006]	[0.006]
Observations	392266	124549	124549	124549	24689	8353	8353	8353
R-squared	0.115	0.159	0.159	0.159	0.12	0.182	0.192	0.201
				C. Movers				
s_η^2	0.017***	0.015***	0.015***	0.015***	0.027***	0.030***	0.029***	0.030***
	[0.001]	[0.002]	[0.002]	[0.002]	[0.004]	[0.007]	[0.006]	[0.006]
s_ε^2	0.101***	0.109***	0.108***	0.108***	0.147***	0.141***	0.138***	0.137***
	[0.002]	[0.003]	[0.003]	[0.003]	[0.005]	[0.007]	[0.007]	[0.007]
Observations	234495	75746	75746	75746	49441	17312	17312	17312
R-squared	0.141	0.177	0.18	0.179	0.156	0.19	0.193	0.193

Note: Estimates correspond to Equation 2.5. Columns (1) and (5) include all available months from 1982 to 2002 but they do not control for productivity shocks. Columns (2) and (6) include periods for which data on productivity shocks is available but they do not control for productivity shocks. Columns (3) to (4) and (7) to (8) control for productivity shocks according to different proxies of productivity: output per worker (3 and 7) and labor productivity (4 and 8).
Standard errors in brackets.
* significant at 10%; ** significant at 5%; *** significant at 1%.

2.5 (Columns 3, 4, 7 and 8).

Results show a number of salient features. First, a comparison between the outcomes obtained including data for the eighties (Columns 1 and 5, for which productivity data at firms level is not available) and those obtained after constraining the period to that for which there is information about firm's productivity but do not control for it (Columns 2 and 6) reveals a slight upward bias in the estimates of the later sample for both, permanent and transitory components in the case of Formal workers. In the case of Informal workers we observe a slight bias in the same direction for just the estimates of the permanent component whereas a slight negative bias for the estimates of the transitory one. Given that major reforms with effects on wages and productivity took place in early and mid nineties, it is not surprising to observe that permanent and transitory effects on income volatility seems to be slightly higher when constraining the analysis to just that period. Second, and

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consistent with the estimates found for other Latin American countries[52] we observe a dominant transitory component (four times -or more- the size of the variance of the permanent income shocks) in the constitution of income volatility[53]. This dominance of the transitory component persists regardless the mobility group or the formality sector. Third, the variance decomposition of formal workers earnings display slightly lower values in the transitory component for stayers compared to that observed for movers while the permanent component is virtually the same. In the case of informal workers, the permanent component for movers is as much as two times the value observed among stayers. Variance of earnings of informal movers show the highest participation of temporary volatility. Overall, Informal workers display more income risk and among them, movers depict earnings volatility with the highest permanent and transitory components. Fourth, the effects after including productivity are negligible for Formal workers. Columns 3 and 4 show unchanged figures compared to Column 2. In the Informal group, estimates of the permanent component remain unchanged as well whereas a reduction of up to 10% (Column 8, Panel B) becomes noticeable in the transitory counterpart.

Related to the last point two observations merit further attention. First, unfortunately and due to data limitations explained before, we cannot appropriately match productivity shocks that are specific to a firm with those shocks observed in specific workers' incomes. Hence we cannot conclude precisely, as in Guiso *et al.* (2005), about how likely are the firms to insure workers against permanent and transitory shocks. However and in spite of the data limitations we can still capture part of the idiosyncratic components of the firms' productivity values[54] allowing for industry cross sectional variation rather than using a systemic measure of productivity as in other studies. Thus, results shown in Table 2.4 broadly support Guiso *et al.* (2005) findings in the sense that for Formal workers (and hence firms), inclusion of (non-systemic) productivity shocks is not reflected into a recomposition of workers income risk. In other words, this could be suggestive of formal firms providing insurance to workers against such shocks. On the other hand, Informal workers earnings volatility reflects an increase of the transitory component when productivity is left out as part of the residual (Column 6) of Equation 2.4. In other words, roughly 10% of the transitory volatility observed in informal earnings can be attributed to firms' productivity shocks. A second observation is that the income risk analysis that we portray at this section is deeply rooted into the second moments space. This is in contrast to the analysis done in the previous section where wage-productivity elasticities are drawn from the space of first moments. This is not to confuse orthogonality observed between productivity shocks and workers income volatility with unitary elasticities observed between wages and productivity.

Overall a deeper analysis of this section merits a study on its own and here we are aimed at just getting broad estimates of measures relating workers wages to firms productivity, taking as much advantage as possible of the sources of variation that we got from both, firms and workers data.

2.7 Conclusion

An empirical assessment of the wages-productivity relation is done for the Brazilian manufacturing sector. Two main messages emerge. First, an analysis of wage-productivity elasticities suggests an heterogeneous response of workers wages to firms labor productivity. Such a response varies accord-

[52] Krebs *et al.* (2010) find comparable estimates for Argentina and Mexico.

[53] This comes at stark contrast with findings for the US by Heathcote *et al.* (2008) where the permanent component is dominant. Also for the US, Meghir & Pistaferri (2004) offers a set of flexible results, according to individuals' educational attainment (less educated individuals display income volatilities with higher permanent components relative to the transitory counterpart) whereas in Carroll & Samwick (1997) the transitory component doubles the permanent one.

[54] In part because we aggregate based upon individual firms, in part because we aggregate at an intermediate -industry-level.

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ing to the degree of workers' mobility and formality but it overall supports a rent sharing approach in the setting of wages: almost unitary elasticities observed for entrants whereas low or non significant elasticities for incumbents along low (for new hires) and high (for stayers) persistence of incomes in the formal sector. Correspondingly, yet keeping new hires earnings as the most flexible, we observe less dissimilar elasticities across mobility groups for informal workers along very low persistence of earnings regardless the degree of mobility. Second, an analysis of income risk suggests that the more formal and less mobile the workers are, the less their income risk (with higher participation of transitory volatility - compared to permanent volatility - in the income variance decomposition). It also suggests that productivity shocks enhance the transitory component on informal workers income volatility whereas it is allegedly insured by firms in the formal sector, whose workers' earnings volatility remains unaffected after controlling for such shocks.

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Chapter 3

R&D and Development. An Instrumented Semiparametrical Approach [1]

Edwin A. Goñi-Pacchioni[2]

Abstract

The contribution of this study is twofold. First, using a world-wide cross country dataset covering the last forty years, it maps the contribution of R&D to the process of economic growth of countries across the whole spectrum of economic development by exploiting methods of varying estimation. Second it addresses the potential existence of double causation between economic growth and investment in R&D by exploiting internal (past investments in R&D) and external instruments (Intellectual Property Rights). Our findings suggest the existence of an inverted U relationship between the level of development of a country and the contribution of investments in R&D to its economic growth. Our findings also support the existence of complementarity effects between production factors when making adjustments by quality to the labor factor.

3.1 Introduction

This paper provides empirical evidence of nonlinear contributions of investment in R&D to the process of economic growth: low but increasing at low levels of economic development and low and non-increasing in countries close to the economic frontier. At early phases of development, low-income countries endowed with poor institutions, unskilled labor and many other barriers to economic progress fail to avail themselves of technological improvements. At mid advanced stages of development, the complementarity effect between production factors (i.e. skilled labor ripping the benefits of innovative technologies) and the lack of stealing and crowding out effects in the process of technological absorption (rather than creation) make the marginal return of R&D to reach its peak. Countries at the highest *plateau* also perceive low returns to R&D investments, most probably due to the stealing and crowding out effects in the process of technological creation in spite of

[1] We thank to Luigi Guiso, Omar Licandro, Massimiliano Marcellino and Morten Ravn for their helpful academical advice in the theoretical and econometric arenas; to Bill Maloney and Daniel Lederman for sharing part of the dataset used in this paper and for insightful and provocative discussions that conduced to the exploration of the ideas studied in this paper; to Zongwu Cai, Mitali Das, Huaiyu Xiong and Xizhi Wu for sharing their SPlus codes useful for the estimation of the first step of the IV procedure; to the participants of the EUI Macro Seminars for their valuable suggestions. The usual disclaimers apply.

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their most favorable conditions for economic growth. Because this heterogeneity in the responses to technological improvements might be induced by the heterogeneity - measured by the distance to the economic frontier - among recipients and originators we revisit the simplest specification used to model economic growth. In our model we allow for the contribution of each argument (investment in physical capital, investment in R&D and growth in the labor force) to vary according to the distance to the economic frontier.

In this sense, the paper attempts to control for a variety of possible estimation issues in estimating R&D returns. First, it updates and extends the parametrical approach followed by Lederman & Maloney (2005), where the use of the GMM estimators allows maximum exploitation of the number of observations and to control for unobserved fixed effects in a dynamic panel context. Further, in the semiparametrical approach, the use of Varying Coefficients allows the assessment of more accurate pointwise estimates that vary according to the level of development of a specific country in a period of time. Second, we also attempt to control for the possible endogeneity of R&D that cast doubt on many estimates of R&D. Barro & SalaiMartin (1995, pp.352), for instance, find the reported rates of return to be implausibly high and speculate that they are due to reverse causality going from productivity growth to R&D expenditures. Existing industry or firm level studies within country rely on internal instruments which weak exogeneity has been questioned. Yet, we also start with internal instruments, and since we work at the country level, we are also able to exploit policy changes that could have a better claim to be exogenous. Thus, as in Lederman & Maloney (2005), we also use the evolution of intellectual property rights protection as external instrument [3]. Our findings suggest that using either internal or external instruments the average non-varying returns to R&D are on the order of magnitude found in other studies. We also find that this non-varying returns constitute too aggregate measures as when we allow the returns to vary according to the level of economic development, a non constant contribution becomes apparent.

In this extent, the paper also follows Griffith *et al.* (2004) in testing to see if rates of return are substantially higher in non frontier countries due to the importance of the second face of R&D. Using level of development as a measure of distance we confirm their finding of increasing returns with distance from the frontier but just up to certain level of development. Beyond this point, more distant countries experience decreasing returns.

In the last section of the paper we speculate on what complementarities may be driving the effect. We adjust labor with quality (attainment) and find that the heterogeneity in the responses of growth to R&D investments persists yet the contributions of the other factors oscillate close around the parametrical non varying estimates.

The remainder of this paper is organized as follows. Section 3.2 debriefs existing theoretical and empirical contributions related to our topic. Section 3.3 provides the sources of our information and some stylized facts related to this data. Section 3.4 presents the general specification to be estimated. Section 3.5 briefly explains the econometric methods necessary to perform varying estimation. Results of these estimations are described in Section 3.6. Section 3.7 explores the complementarity effects of productive factors after adjusting by labor quality. Section 3.8 concludes.

[3] They show that this variable is specifically correlated with R&D, and at the same time it has a strong exogenous component due to pressures exerted by advanced country governments and their strengthening under the TRIPS provisions negotiated within the context of the Uruguay trade round.

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3.2 R&D and Heterogeneity

3.2.1 Previous theoretical contributions

From a microfounded perspective, theoretical contributions are rooted in the notion of creative destruction coined by Schumpeter (1934) and thoroughly examined by the work of Aghion and Howitt during the last decade. As they explain, their work portrays a free enterprise economy that is constantly being disturbed by technological innovations from which some people gain and others lose, an economy in which competition is a Darwinian struggle whose survivors are those that succeed in creating, adopting and improving new technologies. A vast cascade of work stemming from the Schumpeterian concept of creative destruction exists in the theoretical arena.

Our work does not add in this particular theoretical strand but complements it from empirical grounds. Our study addresses the empirical counterpart of existing extensions of the creative destruction models aimed at explaining potential convexities in the contribution of knowledge and innovation into the process of economic growth. One of these extensions is the offsetting effect of distance from the frontier. A number of potential arguments can be contended to explain this.

For instance, Young (1993a) argues that while the traditional approach to creative destruction models implies substitutability of older technologies by newer ones, it ignores the equally important complementarity effect between technologies[4]. This implies a life cycle of new technologies in which first they are complements and after substitutes of old technologies. The tension between these two effects produces a stylized life cycle in which new technologies initially find further invention complementary, but then, as they develop mature and large markets, find that the dominant effect of further invention is substitution.

In a companion study, Young (1993b) is more generic and integrates the two major types of models of endogenous technical change: invention (Romer, 1990) and learning by doing Lucas (1988)[5] A recognition that most products of research at the moment of their invention may be broadly inferior to more mature technologies alerts one to the important role of both production experience and complementary inventive activity in actualizing the productive potential for new technologies. Experience in production increases productivity of new technologies. Thus both, the rate of invention and the rate at which production experience accumulates determines the life cycle - and hence discounted profitability - of new technologies.

Acemoglu *et al.* (2007) provide an alternative model. They analyze the impact of contractual incompleteness and technological complementarities on the equilibrium technological choice: firms choose technology corresponding to the range of intermediate inputs used in production (a greater range increases productivity by allowing greater specialization) and offer contracts to suppliers specifying the required investments in contractible activities. Suppliers choose non-contractible activities in anticipation of the ex-post bargaining payoffs. Greater contractual incompleteness reduces investments in non-contractible and contractible activities and leads to the adoption of less advanced technologies. Their model shows that among countries with identical technological opportunities, those with better contracting institutions specialize in sectors with greater complementarities among inputs.

[4]This is because oftentimes new inventions creates rather than destroy rents for older technologies: when created, new technologies are too expensive or with many few applications and they impulse for older technologies to become more efficient before they are definitely displaced.

[5]As Young explains, models of the first group, focus on factors that influence the incentive to consciously innovate such as institutional framework or market size. Models of the second group focus on factors that incentive to produce different types of goods.

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3.2.2 Previous empirical contributions

In the context of the theoretical background referred in the previous section, and keeping close attention to the effects of factor complementarities, our work adds to two families of empirical contributions.

Heterogeneity

Empirical studies on growth have become popular and important references to understand and test in reality models produced by macro theorist. However, many critiques have been posted to conventional applied methods. Pack (1994) conveys some of them and concludes that the production function interpretation is further muddled by the assumption that all countries are on the same international production frontier, that regression equations that attempt to sort out the sources of growth also generally ignore interaction effects and that the recent spate of crosscountry growth regressions also obscures some of the lessons that have been learnt from the analysis of policy in individual countries. Durlauf (2001), also argue that a problem with conventional growth analysis is the assumption of parameters homogeneity. All in all, if true, problems of the implementation of these empirical efforts suffer from imposing strong homogeneity assumptions, they stem from conventional theoretical setups that are epitome in the mainstream of the literature but that hardly find echo in the less developed part of world. For instance, it is hard to argue that complementarity between production factors occurs at the same intensity in OECD countries as in LDCs for many reasons (institutional barriers, lack of education, multiple externalities and non-convexities reflected in evident market imperfections, etc. as it is argued in the studies referred in Section 3.2.1), and in this context, the case of R&D investments turns to be of our particular interest: if these expenditures assure high economic growth rates what prevents LDCs to spend every available resource into this kind of investment?[6]

Kourtellos (2002) remarks that at a statistical level, evidence of misspecification (generally in the form of nonlinearities) is stressed in a number of studies that suggest that the assumption of a single linear model when applied to all countries turns to be invalid. As he digests, Durlauf & Johnson (1995) employs a tree-regression approach to uncover multiple regimes in the data while Hansen (2000) proposes a threshold regression model that leads to a formal test for the presence of a regime change. Liu & Stengos (1999) employ a semiparametric specification test and an additive semiparametric partially linear model to identify nonlinear growth patterns. Canova (1999) uses a predictive density approach, Desdoigts (1999) employs an exploratory projection pursuit (density estimation) while Kourtellos (2001) uses a projection pursuit regression. All of these find evidence to argue in favor of the existence of multiple steady-state equilibria that should conduce to classify countries into different convergence clubs. In the same vein, theories such as the one of Azariadis & Drazen (1990) suggest that countries that are identical in their structural characteristics but differ in initial conditions may cluster around different steady state equilibria in the presence of increasing returns to scale from some factor of production, market imperfections, non-convexities in the production function, etc. In other words, the introduction of initial conditions such as level of initial human capital, initial income distribution, non-convexities, externalities and capital market imperfections may lead to the emergence of club convergence (Galor, 1996). Our study lifts the restriction of constant contributions of productive factors to production's growth and implicitly it admits multiple equilibria by allowing the coefficients of return to vary across countries and across time.

Complementarity

In line with the theoretical contributions referred earlier, our argument has a second component as heterogeneity translates into non linear returns of R&D allegedly because of complementarity

[6]See Maloney & Rodriguez-Clare (2007) for a discussion in this vein.

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effects.

Arguable poor complementarity between productive factors in countries where the innovation process fails to avail economic growth was already sustained by Acemoglu & Zilibotti (2001), who argue that technologies used in LDCs are developed in OECD and thus are designed to make optimal use of skilled labor force. This induces a mismatch between factors (skill complementarity or skill bias) that makes LDCs economies less productive. Technology-skill mismatch can lead to significant productivity differences even when LDCs have access to all the technologies used in the North. Investors know this and therefore they don't invest in newest technologies prompting asymmetric returns.

In general, cross country analysis of the effects of R&D on growth is recent in the literature. As just explained, significant part of this effort has been placed in a North-South framework in which technological improvements take place into two different forms: (i) Northern developed economies tend to innovate and create new technologies while (ii) the Southern developing economies tend to imitate and absorb (adopt) such innovations (Chin & Grossman, 2004; Grossman & Helpman, 1991; Deardorff, 1992; Chen & Puttitanun, 2005). Yet among the existing empirical contributions, most of them have focused in the U-shape relations between protection to intellectual property and development (Maskus, 2000; Primo Braga & Fink, 2000; Chen & Puttitanun, 2005), some recent contributions follow the approach of differentiated technological development. For instance, Griffith *et al.* (2004) claim that R&D has two faces: in addition to the conventional role of stimulating innovation, it enhances technology transfer (absorptive capacity); Aghion *et al.* (2006) also examine the contribution of human capital to economy-wide technological improvements through these two different channels (imitation and innovation) under the assumption that more developed countries are endowed with more skilled labor and hence is suited to seize the investments in tangible and intangible capital.

3.3 Data and some stylized facts

3.3.1 Data

We use an unbalanced panel of 75 countries covering the period 1960 - 2000[7]. The dataset comprises a number of variables coming from different sources.

R&D

The R&D series from 1960-2000 were compiled by Lederman & Saenz (2005) from national surveys that use a common definition of expenditures that includes fundamental and applied research as well as experimental development. In the lines of the discussion of the purpose of R&D investment (aimed at creating/innovating or aimed at absorbing/imitating technology), the data of R&D considers not only the traditional investments for development of new technologies expected in advanced countries, but also investments in the adoption and adaptation of existing technologies more likely to be labeled as R&D into developing countries[8]. Lederman & Saenz (2005) constructed these series

[7] The spell covering 2001-2005 is disregarded because the few data points available for R&D investment for that period.

[8] Referring to the same dataset, Lederman & Maloney (2005) argue that though it would be desirable to study the evolution, rate of return, and determinants of private R&D, it is still justifiable to work with aggregate R&D for two reasons. First, because the data sources divide R&D not into private and public R&D, but rather into productive and non-productive sectors, the latter accounting for roughly 20% of the total. They point out that the definition of "productive sector" includes both public and private for profit and not-for profit firms while "non-productive sector" includes R&D financed or undertaken by the executive branch of government. Since the productive sector may well include public utilities or other state owned enterprises, the exercise of analyzing how its R&D evolves and its rate of return relative to that of non-productive sector is less interesting than the public/private sector split would be. Second, Lederman & Maloney (2005) also argue that such a

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combining data published by UNESCO, the OECD, the Ibero American Science and Technology Indicators Network (RICYT) and the Taiwan Statistical Data Book, following the definitions convened in the OECD Frascati Manual [9].

Educational attainment and other

We draw this information from Barro & Lee (2010). Attainment accounts for the average years of education of the adult population (25-64 years) as proxy of total human capital. Data on other productive factors come from 2009 Penn World Tables V6.3[10].

IPR

This data comes from an updated version of Ginarte & Park (1997)[11]. This quinquennially index of property rights protection encompasses five components measuring each country's IPR laws coverage and enforcement. The components are the coverage of patent laws across seven industries, membership in three key international agreements, loss of protection due to three potential reasons (namely working requirements, compulsory licensing, and revocation of patents), three types of enforcement mechanisms, and the duration of patents relative to international standards. The composite index ranges between zero and five with higher values indicating stronger IPR protections and enforcement.

3.3.2 Stylized facts

Figures 3.1 and 3.2 are borrowed from Lederman & Maloney (2005). It is clear from Figure 3.1 that R&D expenditures/GDP rise with development as does the rate of increase. Figure 3.2 shows how a few select countries from several regions compare to the predicted value.

Lederman & Maloney (2005) emphasize that countries such as Korea, Finland, and Israel show substantial "take offs" relative to the median trajectory. Two Latin American countries, Argentina and Mexico, which had similar levels of income as Korea and Israel prior to their take off hover on or below the predicted value for their level of development. Both China and India appear to be following more in the footsteps of the "take off" countries than the Latin Americans. As we suggested previously referring to the real role of R&D to yield economic growth, Lederman & Maloney (2005) also stresses the concern of whether the unusually high levels of R&D in some countries, and particularly the dramatic takeoffs of Finland, Israel, Korea, and Taiwan were justifiable investments and to what degree do they owe this to their R&D investments.

Such an evidence is indicative of strong heterogeneity in the contributions and effectiveness of R&D investments regardless the clustering criteria one may use (e.g. natural resource intensive, export oriented, manufacturing intensive, etc.) to look for common patterns. In an extent, by sorting according to economic development we indirectly encompass this myriad of clustering criteria.

In line with the previous discussion, Figure 3.3 summarizes our initial prior. The Figure reveals two regions of high mass concentration when comparing the distribution of countries according to

division seems to occasionally lead to some critical issues in categorization. For instance, if a public company finances its R&D from retained earnings, this will count as productive sector R&D. If instead that R&D is financed by a transfer from the treasury to the firm, it counts as "non-productive" R&D. For several countries in their (and our) sample, there were striking shifts in composition from one year to the next suggesting sensitivity to accounting practices. In contrast, the total R&D series were reasonably stable. Besides these two reasons, many developing countries tabulate only the aggregate values and as they are a primary focus of this paper, we want to include as many as possible.

[9] Proposed Standard practice for Surveys on Research and Experimental Development 2002.

[10] `http://pwt.econ.upenn.edu/php_site/pwt_index.php`

[11] We are grateful to Walter Park who kindly shared the updated version of his original indicator (Park, 2001, 2008)

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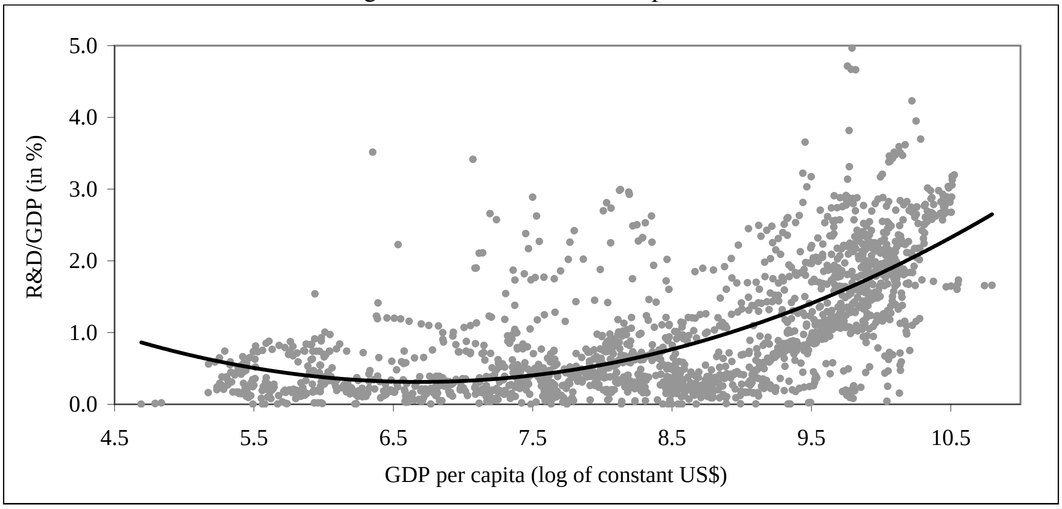
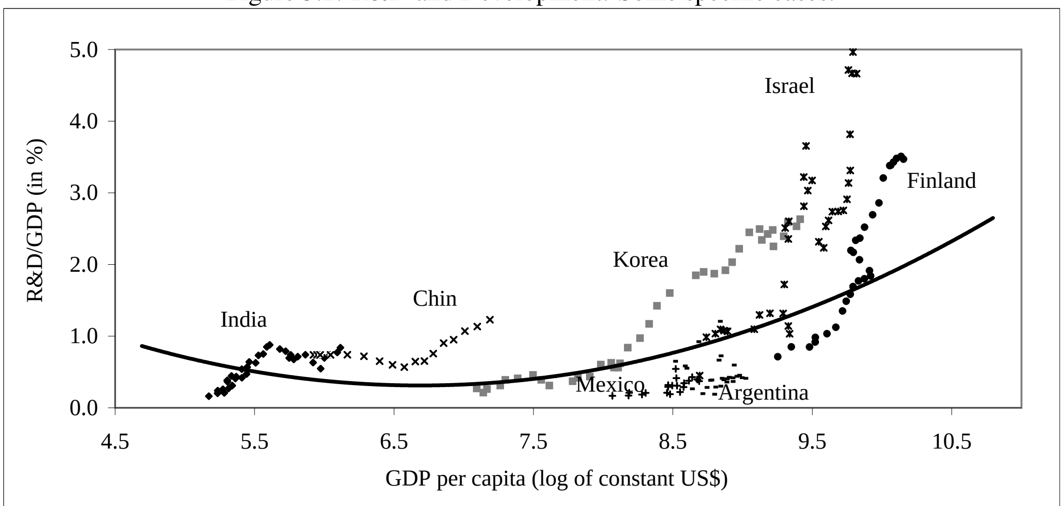
Figure 3.1: R&D and Development



Source: Lederman & Maloney (2005)

Note:Figure 3.1 plots the predicted and observed levels of R&D (expressed as a share of GDP) as function of the log GDP per capita. The predicted value is generated from a regression of the log of the ratio of total R&D expenditures to GDP on log GDP per capita and its squared term.

Figure 3.2: R&D and Development. Some specific cases.



Source: Lederman & Maloney (2005)

their levels of development (distance to the economic frontier) with their distribution according to a set of covariates that potentially influence economic growth (proportion of R&D over GDP, Educational Attainment, IPR). This initial evidence suggests that there exists substantial heterogeneity across countries. As we show later, this heterogeneity is reflected in the substantially varying contribution of some productive factors on growth. Such variation appears to be related to the distance to the economic frontier and also to the quality of the labor factor approximated with its average educational attainment. Preliminary evidence shown in Figure 3.3 portrays the presence of a two steady state equilibria in the process determining aggregate economic output which appears to be induced by certain initial conditions. This, in turn, resembles the the twin peaks found by Quah (1997) in the

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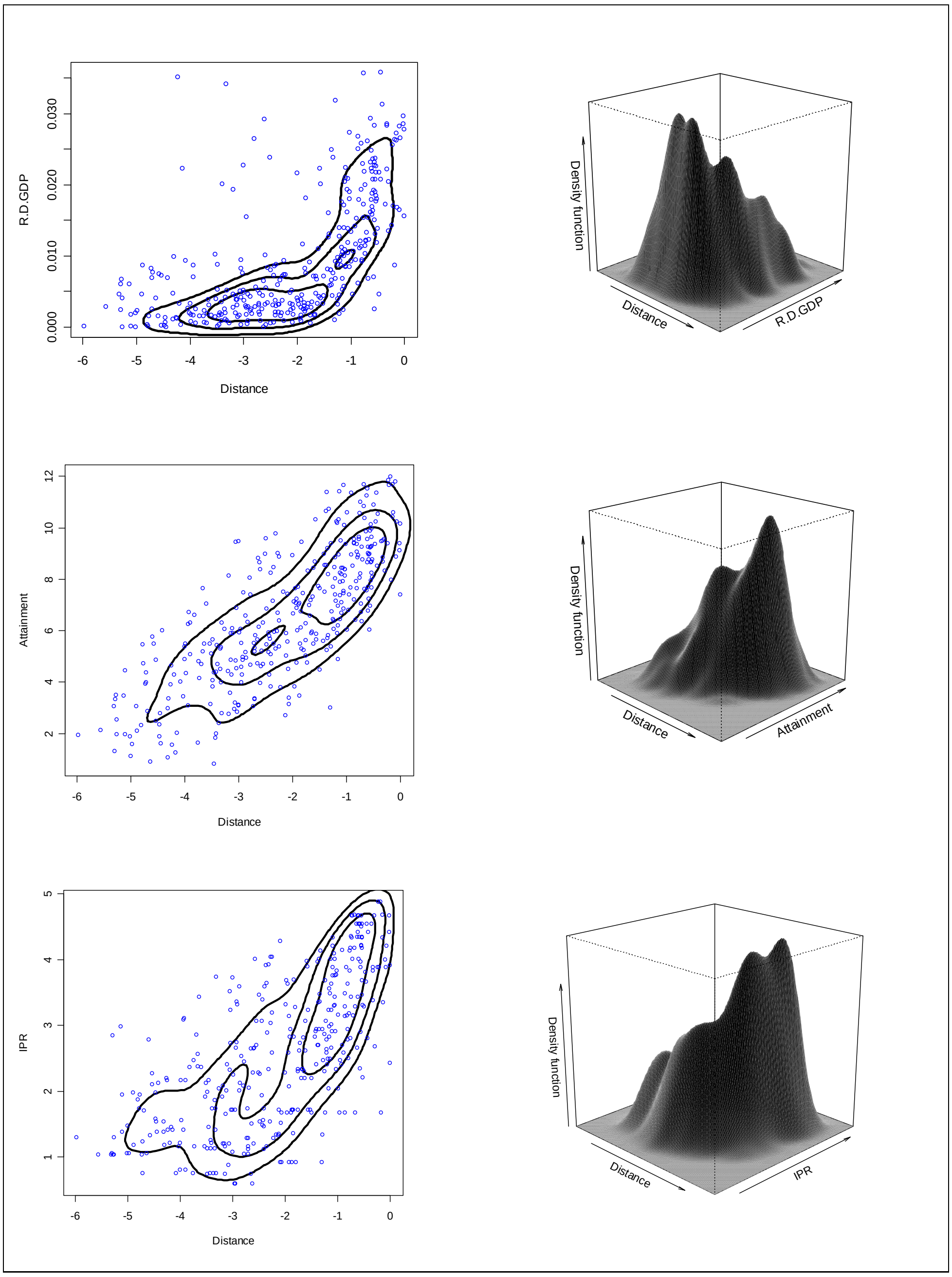
Figure 3.3: Twin Peaks, Kernel Smoothed Bivariate Density



Note: Computations are based (Duong, 2007) ks procedure implemented for the software R. The method implements diagonal and unconstrained data-driven bandwidth matrices for kernel density estimation. Data used in each graph spans the whole sample (casewise and timewise) as described in the Data Section. Graphs on the left column represent the contour levels of the bivariate densities plotted in the right column. Ordered pairs lying outside the outer contour line represent combinations of a Variable (R&D/GDP, IPR, Attainment) and the Distance to the Economic frontier which are not frequently observed in different periods/countries.

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cross-country limiting distribution of income.

Before starting our analysis, we report quinquennial averages of the most important variables involved in the study, for two groups of countries: OECD and non-OECD. Table 3.1 shows the evolution of annual economic growth rates, educational attainment and IPR. These points estimates provides a rough summary of all the initial findings about heterogeneity and complementarity referred in the previous discussion.

Table 3.1: Summary of Indicators
(Average per quinquennium)

	Growth	Investment	Labor growth	Attainment	R&D	IPR
OECD countries						
1961 - 1965	0.061	0.275	0.014	7.144	0.013	2.603
1966 - 1970	0.054	0.272	0.011	7.223	0.012	2.719
1971 - 1975	0.039	0.278	0.012	8.020	0.012	2.703
1976 - 1980	0.032	0.255	0.010	8.466	0.013	3.069
1981 - 1985	0.021	0.228	0.010	8.739	0.016	3.278
1986 - 1990	0.030	0.228	0.008	9.197	0.017	3.523
1991 - 1995	0.020	0.206	0.008	9.670	0.018	4.173
1996 - 2000	0.035	0.213	0.007	10.090	0.020	4.410
Non OECD countries						
1961 - 1965	0.059	0.200	0.025	2.873	0.003	1.505
1966 - 1970	0.051	0.198	0.024	3.011	0.002	1.446
1971 - 1975	0.047	0.219	0.026	3.809	0.004	1.371
1976 - 1980	0.058	0.261	0.027	4.269	0.006	1.357
1981 - 1985	0.024	0.235	0.025	4.968	0.005	1.622
1986 - 1990	0.038	0.229	0.023	5.477	0.006	1.643
1991 - 1995	0.044	0.231	0.022	5.926	0.005	2.520
1996 - 2000	0.037	0.229	0.020	6.689	0.004	3.179

Notes: Growth denotes annual GDP growth rates; Investment is expressed as a share of GDP; Labor growth denotes labor force annual growth rates; Attainment denotes educational attainment in years of schooling; R&D denotes gross investments in R&D as a fraction of GDP; IPR denotes the Park's Intellectual Property Rights Index.

3.4 Model

We follow the now standard approach to estimating rates of return based on a simple production function[12].

$$Y = AK^{\gamma_k} L^{\gamma_l} S^{\gamma_s} \tag{3.1}$$

where Y is the level of output, A is the efficiency with which factors are used (TFP), K the level of physical capital, L the labor stock and S the stock of accumulated R&D, can be rewritten as

[12] See (Jones & Williams, 1998).

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$$\Delta \ln Y = r_k(\frac{I}{Y}) + r_s(\frac{\dot{S}}{Y}) + \gamma_l \Delta \ln L + \Delta \ln A$$

by using the fact that

$$\gamma_x \Delta \ln(X) = r_x(\frac{\dot{X}}{Y}) = r_x(x)$$

Here r_x is the rate of return on factor X, x is the share of investment in X over Y, and γ_x is the output elasticity of factor X.

Consistent with the approach followed by other authors (Lederman & Maloney, 2005; Bravo-Ortega & Lederman, 2008), we follow Griliches (1995) and compute the rate of return of production factors by estimating the following specification:

$$y_{i,t} = \alpha_t(z_{i,t}) + \beta(z_{i,t})X_{i,t} + \delta(z_{i,t})Y_{i,t-1} + \varepsilon_{i,t}$$

where $y_{i,t}$ represents real growth rates[13], α_t is a sample-wide fixed time effect, $Y_{i,t-1}$ is the lagged income[14], $X_{i,t}$ the matrix of covariates or conditioning variables[15][16] and $\varepsilon_{i,t}$ is a time-country specific white noise. Finally, $z_{i,t}$ represents the varying inducing covariate which in our baseline case is the distance to the economic frontier. For expositional purposes, we simplify the former expression into:

$$y_{i,t} = \gamma(z_{i,t})X_{i,t} + \varepsilon_{i,t} \tag{3.2}$$

The suggested approach to specify the determinants of economic growth dates back to Mansfield (1965) and Griliches (1986). As Jones & Williams (1998) note, most of the literature aimed at estimating returns in this tradition is based on neoclassical growth models in which R&D is simply an alternative form of capital investment. Such an approach ignores many of the distortions associated with research that are formalized by the new growth theory[17]. We take advantage of this in the sense that yet the differentiated effect among creators and absorbers, in part induced by the creative destruction process, is not captured by any argument in the last equation, it is reflected in the varying estimates of the contribution of R&D (and indirectly of IPR after instrumenting) to growth. This variation is shown to depend on the degree of economic development which in turn is related to the clustering of originators and imitators of technological progress.

3.5 Estimation

Our estimation strategy follows two related avenues. First we estimate a simple Varying Coefficient Model in which the inducing source of variation is given by the initial level of GDP per capita of each cross section unit relative to the GDP per capita of the country at the frontier at each quinquennium. Then we acknowledge the possibility of endogeneity attributable to simultaneous causation of the R&D investment argument in our specification. We do so by instrumenting this variable using two kind of instruments: internal and external. We use the lagged value of R&D Investments as the internal instrument and the Intellectual Property Rights Index provided by W. Park (Park, 2001, 2008) as the external one.

[13]Measured as the log difference between real GDP in a quinquennium t and real GDP in the precedent quinquennium $t-1$ for each country i.

[14]Log of real GDP for country i. Average value for the quinquennium $t-1$.

[15]In this case, growth of labor and investment in both physical and innovative capital expressed as a share of income, all expressed as quinquennium averages.

[16]All, left hand side and right hand side variables are demeaned by country to remove country fixed effects.

[17]Including monopoly pricing, intertemporal knowledge spillovers, congestion externalities and creative destruction.

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Table 3.2: Impact of R&D into growth. A Non Varying Parametrical Approach

	POOL			POOL WITH FE			PANEL		
	No IV	Internal	External	No IV	Internal	External	No IV	Internal	External
R&D (Gross)	0.262	0.176	1.074**	0.743***	0.503*	-0.103	0.743***	0.846**	1.009**
	[0.180]	[0.202]	[0.492]	[0.220]	[0.302]	[0.805]	[0.265]	[0.402]	[0.510]
Investment	0.208***	0.208***	0.208***	0.268***	0.269***	0.272***	0.268***	0.271***	0.266***
	[0.023]	[0.023]	[0.024]	[0.025]	[0.025]	[0.026]	[0.030]	[0.026]	[0.027]
Labor Growth	0.840***	0.812***	1.095***	0.239	0.223	0.185	0.239	0.4	0.420*
	[0.131]	[0.134]	[0.198]	[0.226]	[0.227]	[0.238]	[0.272]	[0.244]	[0.246]
Initial GDP	0	0	-0.002	-0.050***	-0.049***	-0.045***	-0.050***	-0.042***	-0.041***
	[0.001]	[0.001]	[0.001]	[0.006]	[0.006]	[0.008]	[0.007]	[0.008]	[0.008]
Observations	227	227	227	227	227	227	227	227	227
F(Cragg-Donald)			37.72			13.6			
Sarpan p-value								0.11	0.10
AR(2) p-value								0.43	0.42

	POOL			POOL WITH FE			PANEL		
	No IV	Internal	External	No IV	Internal	External	No IV	Internal	External
R&D (Net)	0.757**	0.619*	2.109**	1.309***	1.679***	-0.182	1.309***	2.333***	2.049**
	[0.313]	[0.371]	[0.955]	[0.339]	[0.543]	[1.426]	[0.408]	[0.745]	[0.936]
Investment	0.204***	0.205***	0.197***	0.262***	0.259***	0.273***	0.262***	0.255***	0.254***
	[0.023]	[0.023]	[0.024]	[0.025]	[0.025]	[0.028]	[0.029]	[0.028]	[0.028]
Labor Growth	0.858***	0.840***	1.038***	0.218	0.226	0.188	0.218	0.395	0.395
	[0.124]	[0.127]	[0.176]	[0.224]	[0.225]	[0.235]	[0.269]	[0.252]	[0.251]
Initial GDP	-0.001	0	-0.002	-0.050***	-0.052***	-0.045***	-0.050***	-0.045***	-0.043***
	[0.001]	[0.001]	[0.001]	[0.006]	[0.006]	[0.008]	[0.007]	[0.009]	[0.009]
Observations	227	227	227	227	227	227	227	227	227
F(Cragg-Donald)			29.16			10.28			
Sarpan p-value								0.25	0.14
AR(2) p-value								0.38	0.37

Notes:
Dependent variable: Average intra-quinquennial yearly GDP growth.
The first column (No IV) in each block of results shows standard estimates without instrumenting R&D. The second column shows the results of the IV estimation using internal instruments (intra-quinquennial average of the first lag of R&D Investments) and the third column shows the results of the IV estimation using external instruments (Park's IPR).
The first block (POOL) shows the OLS estimates on a pooled sample including only an intercept in addition to the explanatory variables. The second block (POOL WITH FE) includes in addition country and period Fixed Effects. The third block includes country Fixed Effects and period dummy variables. The last two columns of the third block reports GMM system estimates.
Standard errors in brackets.
* significant at 10%; ** significant at 5%; *** significant at 1%

3.5.1 Non Varying Coefficient (Parametrical Benchmark)

Table 3.2 reports the conventional parametrical estimates of the non varying coefficient version of Equation 3.2. These constitute our benchmark estimates. A number of issues merit attention. First, we report results for gross and net (of depreciation) R&D investments (see Appendix 3.9.3). Hall & Mairesse (1995) in their work on French and the US firms construct stocks of physical and R&D capital through a perpetual inventory method rather than looking at investments[18]. They, as Griliches & Lichtenberg (1984), find that the point estimate on R&D rises somewhat with an assumed rate of depreciation. We found similar results as the net R&D returns double the gross ones. Second, returns of physical investment prove to be quite robust to all of our specifications hovering between 0.2 and 0.3. The estimated coefficients for the rest of the factors vary considerably across specifications. For instance, significant returns to gross R&D varies between 0.5 and 1.1. These figures are in line with those found by Coe & Helpman (1995) and also by Lederman & Maloney (2005). However,

[18] Griffith *et al.* (2004) assume a negligible rate of depreciation, and Jones & Williams (1998) assume a zero depreciation rate, partly on the grounds that existing studies suggest the best fit occurs without adjustment.

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the sensitivity of the results to respecifications, as well as the diverse estimates available from other existing contributions (see Section 3.1) constitute a first reason to explore varying returns. Third, introduction of Fixed Effects to the pooled specifications yields higher estimates of the returns to R&D and turns estimates of coefficient associated to labor's growth non-significant. Finally, instrumenting yields significant variation in R&D coefficients yet the directions are not clear as the results of the pool regressions with Fixed Effects and the results of the panel GMM system estimators are opposite. These limitations constitute a second reason to pursue the Varying Coefficient estimation as it may be the case that constancy is a too strong assumption that prompts unstable coefficients in the parametrical setup. Yet the panel GMM systems constitute our preferred specification in the non varying parametrical setup, pool Fixed Effects constitute our benchmark specifications to compare the results of the semiparametric Varying Coefficient approach. This is because current state of art in the technology of IVVC estimation allow to mimic pooled estimation.

Table 3.3: Impact of R&D into growth. A Semi Varying Parametrical Approach

	POOL			POOL WITH FE			PANEL		
	No IV	Internal	External	No IV	Internal	External	No IV	Internal	External
R&D III	1.335	0.655	6.42	1.399	3.684**	6.492	1.444	2.404	-0.852
	[0.906]	[1.424]	[6.494]	[1.185]	[1.716]	[11.556]	[1.358]	[1.737]	[0.772]
R&D IV	0.089	0.074	5.589	0.459	0.531	-8.313	0.937*	0.293	0.356
	[0.383]	[0.451]	[4.267]	[0.484]	[0.638]	[12.163]	[0.555]	[0.961]	[0.393]
R&D V	-0.541	-1.066	6.95	0.028	-0.798	-17.466	0.439	-0.342	-0.801
	[0.529]	[0.650]	[4.837]	[0.678]	[0.936]	[21.588]	[0.777]	[1.176]	[0.553]
R&D VI	1.277***	0.938*	6.426*	0.934**	0.545	-1.609	1.233**	1.098	1.839***
	[0.448]	[0.555]	[3.412]	[0.435]	[0.569]	[4.554]	[0.498]	[0.687]	[0.395]
R&D VII	0.306	0.302	4.291	0.3	0.328	-1.608	0.637	0.855*	0.463*
	[0.277]	[0.333]	[2.762]	[0.337]	[0.439]	[3.429]	[0.386]	[0.519]	[0.250]
R&D VIII	0.282	0.263	3.065	0.273	0.279	-0.878	0.638*	0.969**	0.487**
	[0.225]	[0.269]	[1.893]	[0.323]	[0.420]	[2.323]	[0.370]	[0.488]	[0.193]
R&D IX	0.383*	0.339	2.931*	0.136	0.15	-0.894	0.455	0.825*	0.384**
	[0.200]	[0.237]	[1.652]	[0.288]	[0.367]	[1.899]	[0.330]	[0.433]	[0.166]
R&D X	0.191	0.136	2.333*	0.007	-0.004	-0.859	0.277	0.721*	0.136
	[0.174]	[0.206]	[1.401]	[0.286]	[0.353]	[1.551]	[0.328]	[0.416]	[0.144]
Investment	0.215***	0.200***	0.119*	0.247***	0.252***	0.342***	0.234***	0.282***	0.193***
	[0.020]	[0.023]	[0.068]	[0.028]	[0.028]	[0.129]	[0.032]	[0.031]	[0.016]
Labor Growth	0.738***	0.692***	1.278***	0.438*	0.428*	0.308	0.423	0.393	0.797***
	[0.114]	[0.128]	[0.408]	[0.250]	[0.246]	[0.783]	[0.287]	[0.266]	[0.087]
Initial GDP	-0.001	0	-0.003	-0.035***	-0.043***	-0.009	-0.048***	-0.046***	0
	[0.001]	[0.001]	[0.002]	[0.007]	[0.007]	[0.051]	[0.008]	[0.009]	[0.001]
Observations	227	227	227	222	227	227	227	227	227
R-squared	0.557			0.879			0.606		

Note: Standard errors in brackets. * significant at 10%; ** significant at 5%; *** significant at 1%.

Assuming that the contribution of all other covariates but R&D investments into economic growth remains constant at any stage of economic development (semiparametrical results will show later that this assumption holds empirically), we allow R&D returns to vary within the parametrical setup but in a discrete arbitrary fashion. In particular, we interact R&D with dummies that take values of 1 when the observation belongs to a specific decile of the distribution of our variation inducing variable (distance to the economic frontier[19]). Thus, lower deciles correspond to observations very far from the frontier. For reasons that will become apparent later we deliberately omit the interactions of the first two deciles. Table 3.3 shows the results of such an exercise for gross R&D[20]. Two

[19]The distance is computed as the difference of logs of GDP per capita of a country and GDP per capita of the frontier country for each quinquennium.

[20]Yet we do not report them, results for net R&D are qualitatively equivalent.

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facts merit attention here. First, coefficients for all the explanatory variables except R&D remain in the neighborhood of the initial estimates. Second, R&D contributions noticeably vary across deciles showing an inverted U pattern with a significant peak in the sixth decile. This finding confers us the third reason to pursue the varying estimation.

This initial informal assessment shows that conventional parametrical methods suffer of at least two problems in this context. First, non-varying parametrical estimation imposes constancy (and hence homogeneity) of the returns across observed units and second, varying parametrical estimation consumes too many degrees of freedom and hence affects both the point estimates and the standard errors. According to Ahmad *et al.* (2005), the semiparametric varying coefficient model has the advantage that it allows more flexibility in functional forms than a parametric linear model or a semiparametric partially linear model, and, at the same time, it avoids much of the "curse of dimensionality" problem, as the nonparametric functions are restricted only to part of the variable z.

The rest of this section briefly describes the essentials of the two methods of estimation, Varying Coefficients (VC) and Instrumental Variables Varying Coefficients (IVVC) that are relevant for our analysis. It also describes the procedure to estimate the optimal variable bandwidth within a varying coefficient environment.

3.5.2 Varying Coefficient (VC)

We start the estimation procedure with the simplest approach. This approach is based on the work on varying coefficients of Hastie & Tibshirani (1993) and Fan & Zhang (1999, 2000), which follows the conditional linear assumptions given by Eq.(2) with

$$E(y_i|\mathscr{X}_i = X_i, z_i = z_i) = \gamma(z_i)'X_i$$

$$Var(y_i|\mathscr{X}_i = X_i, z_i = z_i) = \sigma^2_{g_i}(z_i)$$

such that the random sample $\{z_i, X_i\}_{i=1}^n$ is drawn from a distribution $F_{z,\mathscr{X}}$. Notice that from now own i refers to a country-period observation rather than denoting to just a country as it was before. This is just to avoid notational complications in the next subsections.

For each given point z_0[21], we approximate the functions $\gamma_j(z), j = 1, ..., p$, locally as

$$\gamma_j(z) \approx \textstyle\sum_{l=0}^q c_{jl}(z - z_0)^l$$

for sampling points z in a neighborhood of Z_0. This results in the following locally weighted least squares problem:

$$\min_{a_1,...,a_p,b_1,...,b_p} \sum_{i=1}^n [y_i - \sum_{j=1}^p \sum_{l=0}^q c_{jl}(z_i - z_0)^l X_{ij}]^2 K_h(z_i - z_0) \tag{3.3}$$

where $K_h(.) = (1/h)K(./h)$ and $K(.)$ is the Epanechnikov kernel $K(z) = (3/4)(1-z^2)I(|z| \leq 1)$. Let $\mathbf{g} = (g_1, ..., g_n)'$, $\mathbf{W} = diag(\frac{1}{h}K(\frac{z_1-z_0}{h}), ..., \frac{1}{h}K(\frac{z_n-z_0}{h}))$ and

$$\mathbf{X}_q = \begin{pmatrix} X_{11} & \dots & (z_1 - z_0)^q X_{11} & \dots & X_{1p} & \dots & (z_1 - z_0)^q X_{1p} \\ \vdots & \ddots & \vdots & \ddots & \vdots & \ddots & \vdots \\ X_{n1} & \dots & (z_n - z_0)^q X_{n1} & \dots & X_{np} & \dots & (z_n - z_0)^q X_{np} \end{pmatrix}$$

[21] This moving *anchor* is the argument that will induce the variation in the weighting functions of the optimization problem 3.3 and hence generating varying estimates.

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The solution to the optimization problem stated in the Expression 3.3 is

$$\hat{\gamma}_j(z) = e'_{2j-1,2p}(\mathbf{X}'_1\mathbf{W}\mathbf{X}_1)^{-1}\mathbf{X}'_1\mathbf{W}\mathbf{y} \tag{3.4}$$

where $e_{k,m}$ denotes the unit vector of length m with 1 at the k_{th} position. The conditional variance is given by:

$$\hat{\sigma}^2(z) = \frac{\sum_{i=1}^n (y_i - \hat{y}_i)^2 K_h(z_i - z_0)}{tr\{\mathbf{W} - \mathbf{W}\mathbf{X_1}(\mathbf{X'_1}\mathbf{W}\mathbf{X_1})^{-\mathbf{1}}\mathbf{X'_1}\mathbf{W}\}} \tag{3.5}$$

where

$$\hat{\mathbf{y}}_i = (\hat{y}_1, ..., \hat{y}_n)' = \mathbf{X_1}(\mathbf{X'_1}\mathbf{W}\mathbf{X_1})^{-\mathbf{1}}\mathbf{X'_1}\mathbf{W}\mathbf{y}$$

and, under local homoskedasticity,

$$V\hat{a}r(\hat{\gamma}_j(z)|\Im) = e'_{2j-1,2p}(\mathbf{X'_1}\mathbf{W}\mathbf{X_1})^{-\mathbf{1}}(\mathbf{X'_1}\mathbf{W^2}\mathbf{X_1})(\mathbf{X'_1}\mathbf{W}\mathbf{X_1})^{-\mathbf{1}}e'_{2j-1,2p}\hat{\sigma}^2(z)$$

where $\Im = (x_{11}, ..., x_{1n}, ..., x_{p1}, ..., x_{pn})$ denotes the vector of observed covariates.

3.5.3 Instrumental Variables Varying Coefficient (IVVC)

The Instrumental Variables Varying Coefficient approach we follow in this section is based on the work of Cai *et al.* (2006) on functional coefficient IV models that assume the following form:

$$\begin{cases} Y = g(\mathbf{X}, \mathbf{Z_1}) + \varepsilon \\ g(\mathbf{X}, \mathbf{Z_1}) = \sum_{j=0}^d g_j(\mathbf{Z_1})X_j = \mathbf{X}'\mathbf{g}(\mathbf{Z_1}) \\ E(\varepsilon|\mathscr{Z}) = 0 \end{cases} \tag{3.6}$$

where Y is an observable scalar random variable, $g_j(.)$ are the unknown structural functions of interest, $X_0 \equiv 1, \mathbf{X} = (X_0, X_1, ..., X_d)'$ is a $(d+1)$ dimension vector consisting of d regressors, $\mathbf{g}(\mathbf{Z_1}) = (g_0(\mathbf{Z_1}), ..., g_d(\mathbf{Z_1}))'$, and $\mathscr{Z}$ is a $(m+l)$ dimension vector consisting of a m-dimension vector $\mathbf{Z_1}$ of exogenous variables and a l-dimension vector $\mathbf{Z_2}$ of instrumental variables.

The reduced form associated to the previous problem can be stated as:

$$E(Y|\mathscr{Z}) = \sum_{j=0}^{d} g_j(\mathbf{Z_1})E(X_j|\mathscr{Z}) = \sum_{j=0}^{d} g_j(\mathbf{Z_1})\pi_j(\mathscr{Z}) = \pi'(\mathscr{Z})\mathbf{g}(\mathbf{Z_1}) \tag{3.7}$$

where $\pi_j(\mathscr{Z}) = E(X_j|\mathscr{Z})$ and $\pi(\mathscr{Z}) = (1, \pi_1(\mathscr{Z}), ..., \pi_d(\mathscr{Z}))' = E(\mathbf{X}|\mathscr{Z})$

Comparing equation 3.2 to equation 3.7, one observes that $g_j(.)$ is just a vector of functional coefficients of $\pi(\mathscr{Z})$. Hence, $g_j(.)$ could be recovered from a local linear regression of Y on $(\mathscr{Z})$. Given that $\pi_j(\mathscr{Z})$ are unknown, this suggests a two-stage estimation method. The first step estimates $\pi_j(\mathscr{Z})$ by a regression of $\mathbf{X}$ on (Z), while the second step estimates $g_j(.)$ by a regression of Y on (Z) and the first estimate $\hat{\pi}_\mathbf{j}(\mathscr{Z})$. In other words, while the first stage involves estimation of the conditional expectations $E(X_j|\mathscr{Z})$, the second stage substitutes the estimated conditional expectations in place of $\pi(\mathscr{Z})$ in 3.7.

Details of the two stages estimation procedure are provided in Appendix 3.9.1.

The estimate of the asymptotic covariance matrix takes the form:

$$\hat{\Sigma} = (1/n)\hat{S}_{n,0}^{-1}(\hat{\mho}_\eta + \hat{\mho}_\epsilon + \hat{\mho}_{\eta\epsilon})\hat{S}_{n,0}^{-1} \tag{3.8}$$

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where:

$\hat{S}_{n,0}$ is defined as before
$\hat{\mho}_{\eta} = (1/n)\sum_{i=1}^{n} \hat{\pi}_{-i}\hat{\pi}'_{-i}\hat{\eta}_i^2 L_{h_2}(\mathbf{Z}_{i1} - \mathbf{z}_1)$
$\hat{\mho}_{\epsilon} = (1/n)\sum_{i=1}^{n} \hat{\pi}_{-i}\hat{\pi}'_{-i}\mathbf{g}(\hat{\mathscr{X}})_{\mathbf{i}}\hat{\epsilon}_i\hat{\epsilon}'_i\mathbf{g}(\hat{\mathscr{X}})_{\mathbf{i}} L_{h_2}(\mathbf{Z}_{i1} - \mathbf{z}_1)$
$\hat{\mho}_{\epsilon} = (1/n)\sum_{i=1}^{n} \hat{\pi}_{-i}\hat{\pi}'_{-i}\hat{\epsilon}'_i\mathbf{g}(\hat{\mathscr{X}})_{\mathbf{i}} L_{h_2}(\mathbf{Z}_{i1} - \mathbf{z}_1)$
$\hat{\eta}_i = Y_i - \sum_{j=0}^{d} \hat{g}_j(\mathbf{Z}_{i1})\hat{\pi}_{j,-i}(\mathscr{X}_i)$
$\hat{\epsilon}_i = \mathbf{X}_i - \hat{\pi}_{-i}$

3.5.4 Bandwidth Selection

Kernel smoothing methods - like the one just described - are non trivially affected by the bandwidth selection. Larger bandwidth may gain on variance side, but loses on bias side. For the case of IVVC, Cai *et al.* (2006) note that the second stage estimation is not sensitive to the choice of the first stage bandwidth so long as h_1 is chosen small enough such that the bias in the first stage is not too large. To select the bandwidth for the first stage fitting and for the initial pilot bandwidth required in the second stage we look for a minimizer of the integrated residual squares criterion (IRSC) as in Fan & Gijbels (1995). In implementing the second step, the choice of bandwidth can be carried out as in a standard nonparametric regression. Yet there appears to be no results in the literature for a data-driven bandwidth selection with optimal properties (Newey *et al.*, 1999) we follow Zhang & Lee (2000) optimal variable bandwidth selection procedure as they show that this procedure is superior to the theoretical optimal constant bandwidth and to the bandwidth obtained by cross validation methods. Section 3.9 describes the technical details of the procedure of varying optimal bandwidth selection.

We implement all the estimation procedures using the software R. The next section describes the results.

3.6 Results

Our findings support two main ideas: first, that the heterogeneity observed in the economic development of different countries matters in their capacity to seize opportunities to grow thanks to technological improvements and second, that this might happen because productive factors fail to complement one with the other in the productive process.

3.6.1 Varying Coefficients

Figure 3.4 shows the evolution of the VC related to Investment in Physical Capital, Labor Growth and Initial Output when mapped against the distance to the economic frontier. The three horizontal lines crossing each graph correspond to the OLS pool estimates with Fixed Effects along the the 95% Confidence Interval. Top Panel of Figure 3.5 has the corresponding results for gross R&D returns[22]. A number of interesting observations arise.

First, all coefficients plotted in Figure 3.4 show relative constancy for almost the whole spectrum of the varying inducing variable. However, while investment in physical capital yields a roughly positive homogenous contribution to the growth of underdeveloped and developing economies, this marginal contribution decays for very large economies (Oulton & Young 1996 find evidence that

[22]For the case of R&D, the three horizontal lines depicted for the case of the external instrument correspond to the coefficients and confidence interval coming from the parametrical estimation using pooled fixed effects with internal instruments. This is because the parametrical estimates with external instruments yield non significant and odd estimates.

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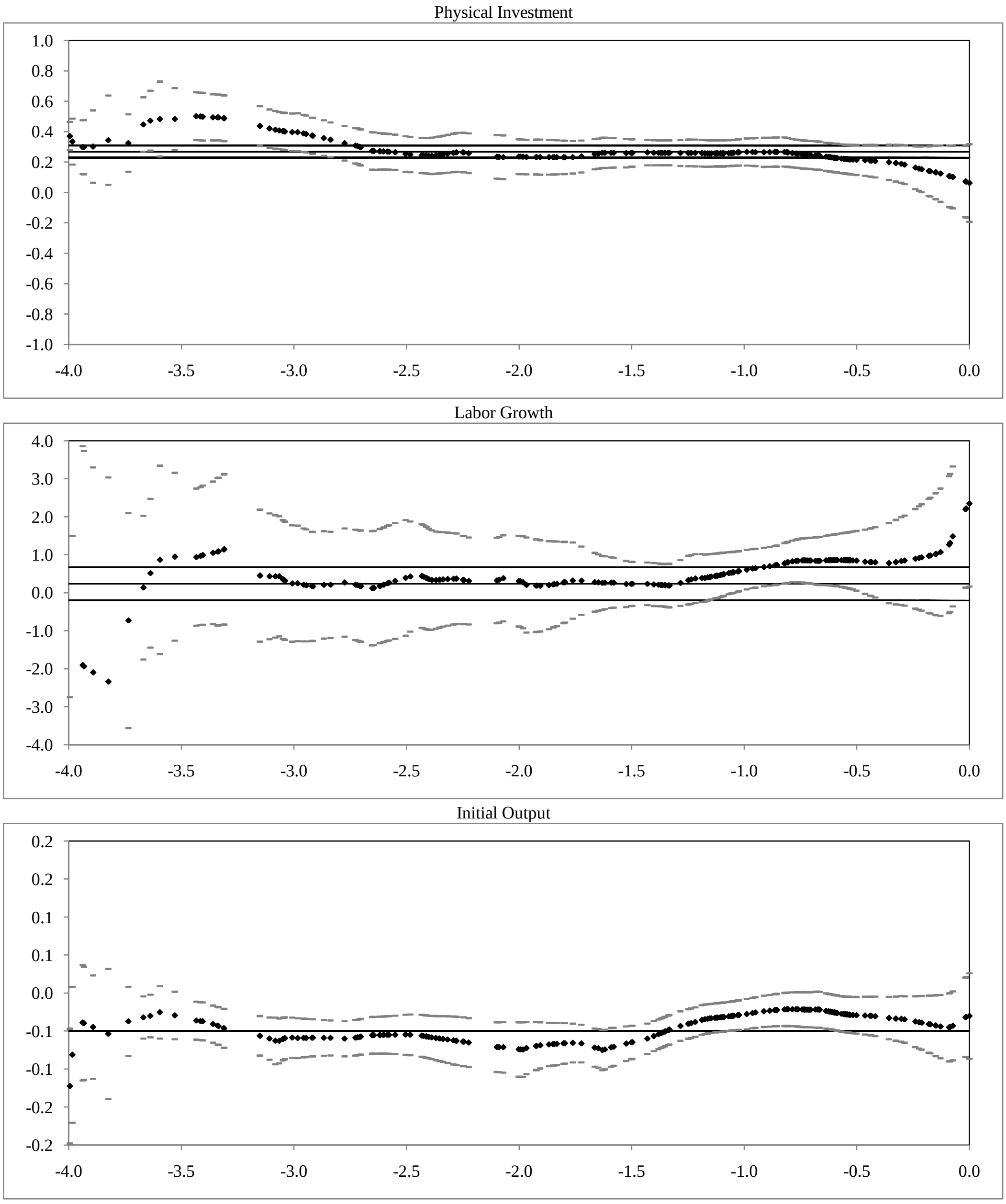
the importance of investment in physical capital may fall as countries grow richer). Something analogous is observed in the contribution of labor growth: roughly constant contributions for underdeveloped and developing countries but increasing at the very high end of the spectrum (Becker *et al.* 1990 and Galor & Weil 2000 models stylize economies where higher *plateaus* of development are characterized by higher educational levels, lower fertility rates and higher rates of return to human capital investment).

Figure 3.4: Varying Coefficients induced by the Distance to the Economic Frontier.



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Second, we observe that in all cases, the non varying estimates are consistent with the average values of the varying ones. Moreover, the confidence interval of the non varying results lie within the corridor depicted by the confidence interval of the varying coefficients. Again, this happens almost everywhere except for the very high end of the spectrum. We exploit the argument of complementarity and cope with this in a later section.

Third, evident non linear returns to investment in R&D are observed across the spectrum of the distance to the frontier (non varying estimates and confidence interval cross more than once to the varying estimates and varying confidence intervals): allegedly, at early stages of development, lack of complementarity effects between factors prevents economies to seize the opportunities provided by newer technologies. In turn, mid developed countries might face reduced costs of catching up via absorption and imitation while most advanced economies show similar returns than those observed in LDCs probably due to higher costs of innovation both, in the process of creation (expanding the frontier) and in the process of destruction (stealing and crowding out effects). Arguments found in the theoretical literature for these results are in the lines of the work of Aghion *et al.* (2001) and all the related literature of Schumpeterian growth, while consistent evidence is found in empirical work such as Griffith *et al.* (2004), Acemoglu & Zilibotti (2001), etc.

Among the Varying Coefficients just described, one is of our utmost interest[23]: the contribution of investments in R&D to growth. For this reason and given the potential reverse causation between economic growth and the R&D investments variable that would generate inconsistent estimates, we refine our computations by implementing the IVVC estimation. To this end, we use the lag of R&D investments and an Intellectual Property Rights index as internal and external instruments respectively.

3.6.2 Instrumental Variables Varying Coefficients

Valid instrumental variables must be relevant and exogenous. As described by Conley *et al.* (2007) and Kray (2008), it is very difficult to perfectly satisfy the orthogonality and exclusion conditions in empirical work. Yet IPR is not immune to this critique, we contend it might fulfill both required conditions. Empirical studies on the relationship between IPRs and innovations/growth, including Deolalikar & Lars-Hendrik (1989); Gould & Gruben (1997); Lach (1995); Thompson & Rushing (1996, 1999); Maskus & McDaniel (1999), and Crosby (2000) have taken IPR as exogenous. We continue in the same tradition in spite of compelling evidence collected in contributions that argue in favor of the endogenous determination of this variable (Ginarte & Park, 1997; Maskus, 2000; Maskus & Penubarti, 1995; Chen & Puttitanun, 2005) contending that IPR and institutional development - well known determinant of growth- are close related. We still use IPR as exogenous instruments for a number of reasons (we endorse here the arguments provided by Lederman & Maloney (2005)).

First, there is a strong correlation between IPR and R&D investments (instrument relevance), in other words, if there is a channel through which IPRs would exert some effect on economic growth that is the innovation channel, i.e. through the incentives it confers to make innovation-oriented technological investments after protection of the appropriability rights of rents are established.

Second, regarding the instrument exogeneity, it seems unlikely that the correlation is driven by those potential R&D investors demanding protections since much of the variation of the index appears driven by external forces: most advances in IPR protection in recent years appear driven by pressure from the United States, Europe and Japan in the context of trade rounds where protection of IPRs

[23]For which in turn, a non varying evolution seems improbable after the evidence provided by both, the semi-varying parametric and varying semiparametric approaches

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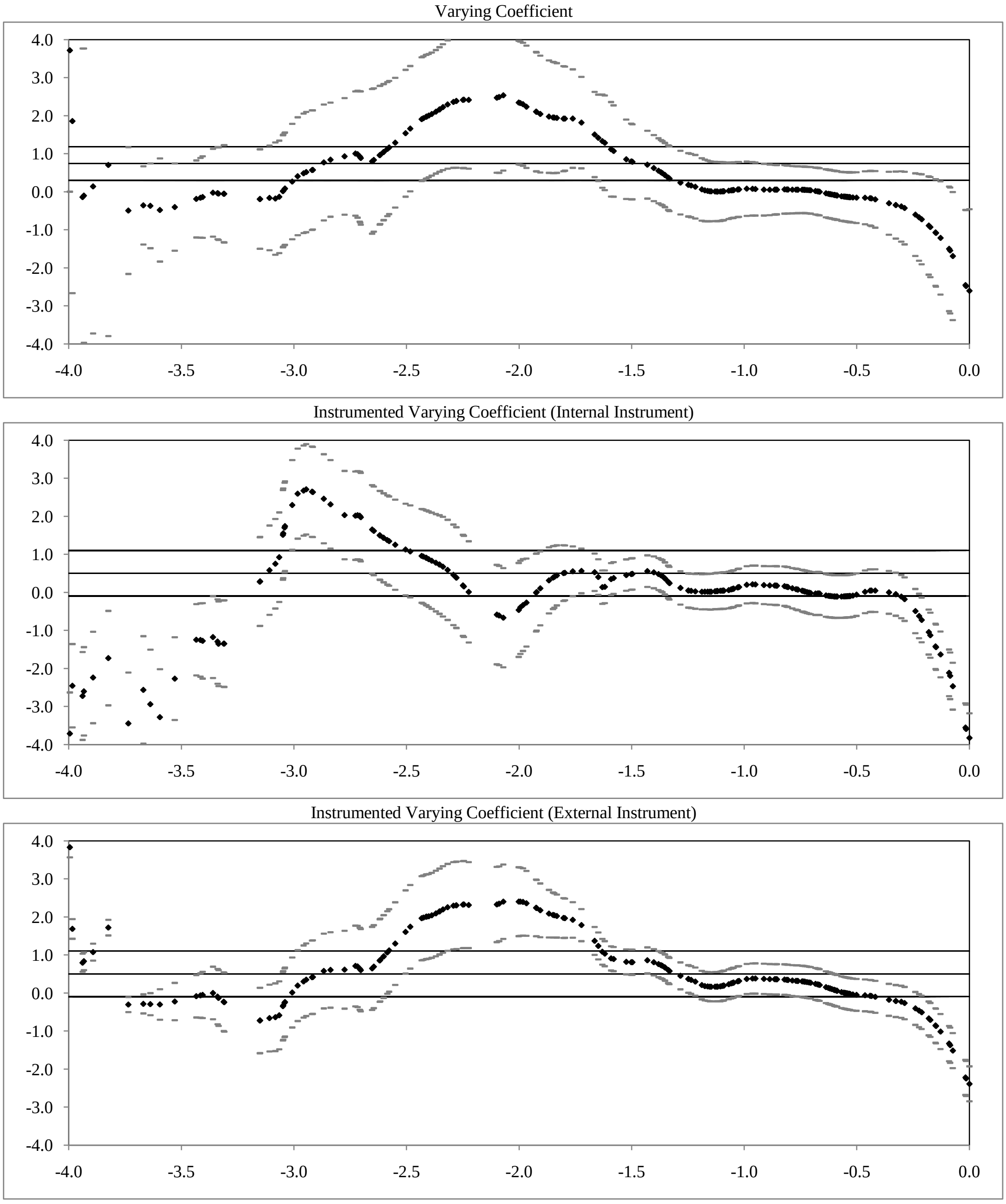
Figure 3.5: Varying Return of R&D Investment



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has, in fact, not been driven by a perception of their benefit *per se* (in fact, perhaps it is the reverse) for developing countries, but rather as a price to pay for greater access to advanced country markets.

Third, regarding the exclusion restriction, Lederman & Maloney (2005) run a number of standard parametrical experiments to verify the validity of the exclusion restriction of IPRs in the equation that determines economic growth and conclude that even if "...it might be argued that governments would have implemented stricter IPRs protections in anticipation of further growth, thus contaminating the instrument, ... *(evidence suggests that)* the concern that lagged R&D or growth would lead to expectations of higher innovative effort in the future, and hence a need to establish intellectual property rights, seems ill founded."

Bottom Panels in Figure 3.5 shows the evolution of the IVVC related to Investment in R&D when mapped against the distance to the economic frontier and when lagged R&D Investments and IPR are used as internal and external instrument for R&D investment respectively. Varying Coefficient estimates of R&D, still follow an inverted U shape, yet with some different patterns for the different set of instruments. This is consequence of the quality of the instruments in the correction of the endogeneity bias.

Figure 3.5 combined with information reported in Appendix 3.9.4 allow us to read the evolution of the point estimates at specific anchors used to compute each estimate[24]. A number of observations are worth it to remark.

First, we can delimit three regions according to the distance to the economic frontier: LDCs (observations to the left of -3 -approximately- in the horizonal axis), developing (somewhere between -3 and -1.5 approximately) and developed economies (observations lying to the right of -1.5). Each of these clusters show a very distinct contribution of R&D to growth as we discussed previously.

Second, after instrumenting we observe three distinct behaviors according to the region of development. For regions closer to the frontier, both internal and external instruments indicate biased VC that underestimate the actual parameters (IV coefficients show higher values) with more incidence when the internal instrument is used. For mid-developed regions, the external instrument still points towards a negative bias -yet more modest than that observed for developed countries- whereas the internal instrument favors lower instrumented estimates. For regions far from the frontier, internal instruments point towards negative bias again while external instruments cause meager effects on the new estimates. All of these patterns will persist or get strengthened after adjusting labor by its quality as will be shown in the next section. Limitations of internal instruments[25] as well as a stable pattern observed in the direction of the bias across the whole spectrum of economic development make us to consider the estimates using external instruments as our preferred outcomes[26].

Third, assuming that imitation/absorption of technology is more distinctive of developing countries while innovation/creation is more distinctive of developed ones, we can track the performance of

[24]Each point estimate does not strictly refer to the coefficient of "a country-period" unit but to the estimated coefficient based on all observations after they have been weighted according to a function of the difference between the value of the variation-inducing variable of the specific country-period used as anchor (z_0 in the terminology used in the Estimation section) and the value of the variation-inducing variable - distance to the economic frontier - of the other observations not used as anchor (z_i in the terminology used in the Estimation section). Therefore it is reasonable to refer as each point estimate as representative of the "country-period" unit of the anchor as all observations are weighted according to their distance to it.

[25]Namely, weak exogeneity, loss of degrees of freedom, etc.

[26]Focusing on the externally instrumented estimates, the asymmetric response to the instrumenting process observed according to levels of development may be reflecting the degree of effectiveness of the instrument to act according to the level of development. This property is desirable as it does not violate the exclusion restriction (growth rates are not necessarily correlated to development) and at the same time allows for a flexible -and arguably more accurate- determination of the size of the bias.

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specific countries and see if their classification was unchanged, upgraded or downgraded across periods. For example, we observe that very developed countries like the US, UK or Germany have experienced the same rate of return to R&D during the last forty years and they have been always in the region of innovating countries. In the other end, Sub Saharan countries like Zambia, Togo or Kenya have remained in the less developing area with negative returns to R&D. There are also upgrading cases like Brazil, Chile or Colombia which jumped from the less developed cluster to the developing pool and the recent Western European emerging economies Spain, Ireland and Finland which made their way from developing to developed countries hand in hand with innovation oriented policies. Finally, Argentina is the only country in our sample that appears to have jumped back and gotten more distant to the frontier after forty years and hence the country exhibits better returns to R&D now than during the sixties. Alongside with these re-classifications, returns to R&D for same countries at different stages of development reflect our contentions: upgrading LDCs absorbers jump from meager returns (below 0.5 and even negative for the lowest end) to overwhelming rates (up to 2 approximately) when they become developing imitators. Similarly, upgrading imitators that become developed innovators experience a reduction in their returns (that then hover between 0.5 and 1).

3.7 Complementarity effects and Robustness Check. Adjustment by labor quality.[27]

Evidence shown in the previous section supports our argument about the heterogeneous contribution of R&D investments to growth according to the stage of development of the recipient/originator of the innovation efforts. Besides the clear inverted U shape we observe in the coefficient accompanying R&D investments, Figure 3.4 shows some degree of heterogeneity in the returns of physical investment and in the share of the labor factor. One possible explanation is the high complementary effect that is specially notorious in high income countries. This complementary effect should be at least partially acknowledged (i.e. the heterogeneity picked up in the returns of physical investment and labor growth at the high end should be mitigated) after adjusting by factor quality. In the case of labor, quality can be approximated by average educational attainment of the population.

Typical production functions in growth models (up to 1966) consider output determined by tangible capital and effective labor. However, even then, some models had already recognized the marginal productivity of education. In old dated models this marginal productivity is a function of the inputs and of the current technology and can remain positive forever even if the technology is stationary. Nelson & Phelps (1966) model considers that education has a positive payoff only if the technology is always improving (i.e. education and technology are in great extent complementary). Their hypothesis is that educated people make good innovators so that education speeds the process of technological diffusion. To address this, they propose a model of technological diffusion in which the rate at which latest theoretical technology is realized in improved technological practice depends upon educational attainment and upon the gap between the theoretical level of technology and the level of technology in practice. With this idea, we follow Caselli (2004), who in the tradition of Hall & Jones (1999), augments the labor factor according to its quality (understood as a the average human capital)[28]. In other words, the production function that originates Equation 3.1 is transformed into:

$$Y = AK^{\gamma_k}(hL)^{\gamma_l}S^{\gamma_s}$$

where $h = exp(\phi(s))$, s is the average years of schooling, and $\phi(s)$ a piecewise linear function

[27] This Section is significantly indebted to the advise and guidance of Bill Maloney who suggested the quality adjustment of Labor factor.

[28] We thank Aart Kray for his suggestions on this regard.

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Table 3.4: Impact of R&D into growth. A Non Varying Parametrical Approach controlling by Labor Quality

	POOL			POOL WITH FE			PANEL		
	No IV	Internal	External	No IV	Internal	External	No IV	Internal	External
R&D (Gross)	0.262	0.176	1.074**	0.743***	0.503*	-0.103	0.743***	0.846**	1.009**
	[0.180]	[0.202]	[0.492]	[0.220]	[0.302]	[0.805]	[0.265]	[0.402]	[0.510]
Investment	0.208***	0.208***	0.208***	0.268***	0.269***	0.272***	0.268***	0.271***	0.266***
	[0.023]	[0.023]	[0.024]	[0.025]	[0.025]	[0.026]	[0.030]	[0.026]	[0.027]
Labor Growth	0.840***	0.812***	1.095***	0.239	0.223	0.185	0.239	0.4	0.420*
	[0.131]	[0.134]	[0.198]	[0.226]	[0.227]	[0.238]	[0.272]	[0.244]	[0.246]
Initial GDP	0	0	-0.002	-0.050***	-0.049***	-0.045***	-0.050***	-0.042***	-0.041***
	[0.001]	[0.001]	[0.001]	[0.006]	[0.006]	[0.008]	[0.007]	[0.008]	[0.008]
Observations	227	227	227	227	227	227	227	227	227
F(Cragg-Donald)			37.72			13.6			
Sarpan p-value								0.11	0.10
AR(2) p-value								0.43	0.42

	POOL			POOL WITH FE			PANEL		
	No IV	Internal	External	No IV	Internal	External	No IV	Internal	External
R&D (Net)	0.757**	0.619*	2.109**	1.309***	1.679***	-0.182	1.309***	2.333***	2.049**
	[0.313]	[0.371]	[0.955]	[0.339]	[0.543]	[1.426]	[0.408]	[0.745]	[0.936]
Investment	0.204***	0.205***	0.197***	0.262***	0.259***	0.273***	0.262***	0.255***	0.254***
	[0.023]	[0.023]	[0.024]	[0.025]	[0.025]	[0.028]	[0.029]	[0.028]	[0.028]
Labor Growth	0.858***	0.840***	1.038***	0.218	0.226	0.188	0.218	0.395	0.395
	[0.124]	[0.127]	[0.176]	[0.224]	[0.225]	[0.235]	[0.269]	[0.252]	[0.251]
Initial GDP	-0.001	0	-0.002	-0.050***	-0.052***	-0.045***	-0.050***	-0.045***	-0.043***
	[0.001]	[0.001]	[0.001]	[0.006]	[0.006]	[0.008]	[0.007]	[0.009]	[0.009]
Observations	227	227	227	227	227	227	227	227	227
F(Cragg-Donald)			29.16			10.28			
Sarpan p-value								0.25	0.14
AR(2) p-value								0.38	0.37

Note: Standard errors in brackets. * significant at 10%; ** significant at 5%; *** significant at 1%.

with varying slopes[29].

Table 3.4 is a modified version of Table 3.2 after adjusting the labor factor. Besides the fact that the returns to physical investment remain almost unaffected, two major effects are observed when comparing these two tables. First, the returns of R&D are slightly lower in the new sets of results.

Second, the labor term is reduced dramatically, reaching even negative (see Pritchett 2001 for a possible explanation) and significant (or barely significant) values. Again we do not argue much in favor of these parametrical non-varying outcomes as our argument endorses a varying estimation. Still, we report these results as they constitute new reference values drawn from the parametrical setup that we use to compare our IVVC results after adjusting by quality of labor.

Figures 3.6 and 3.7 are modified versions of Figures 3.4 and 3.5 respectively after adjusting the labor factor. In Figure 3.6 we observe that the divergence of the coefficients related to the adjusted labor growth rates in countries close to the frontier has been mitigated. In Figure 3.7 we observe that the inverted U is still present in the returns of R&D investments. The confidence interval for these estimates slightly gets better compared to those observed in the case of non adjusted labor and the direction of the bias observed after doing external instrumentation becomes evidently non-positive. Thus, yet the parametrical estimates do not necessarily improve after adjusting the labor factor, the non varying estimates do. overall, the adjusted IVVC results strengthen the idea of complementarity of factors (labor adjusted by quality complements at a non varying rate to the physical capital) and of heterogeneous contribution of R&D (the inverted U is still present with slightly better accuracy).

[29] ($\phi(s) = 0.134s$ if $s \leq 4$; $\phi(s) = 0.134{\times}4{+}0.101(s{-}4)$ if $4 < s \leq 8$; $\phi(s) = 0.134{\times}4{+}0.101{\times}4{+}0.068(s{-}8)$ if $8 < s$)

Figure 3.6: Varying Coefficients induced by the Distance to the Economic Frontier (controlling by Labor Quality)

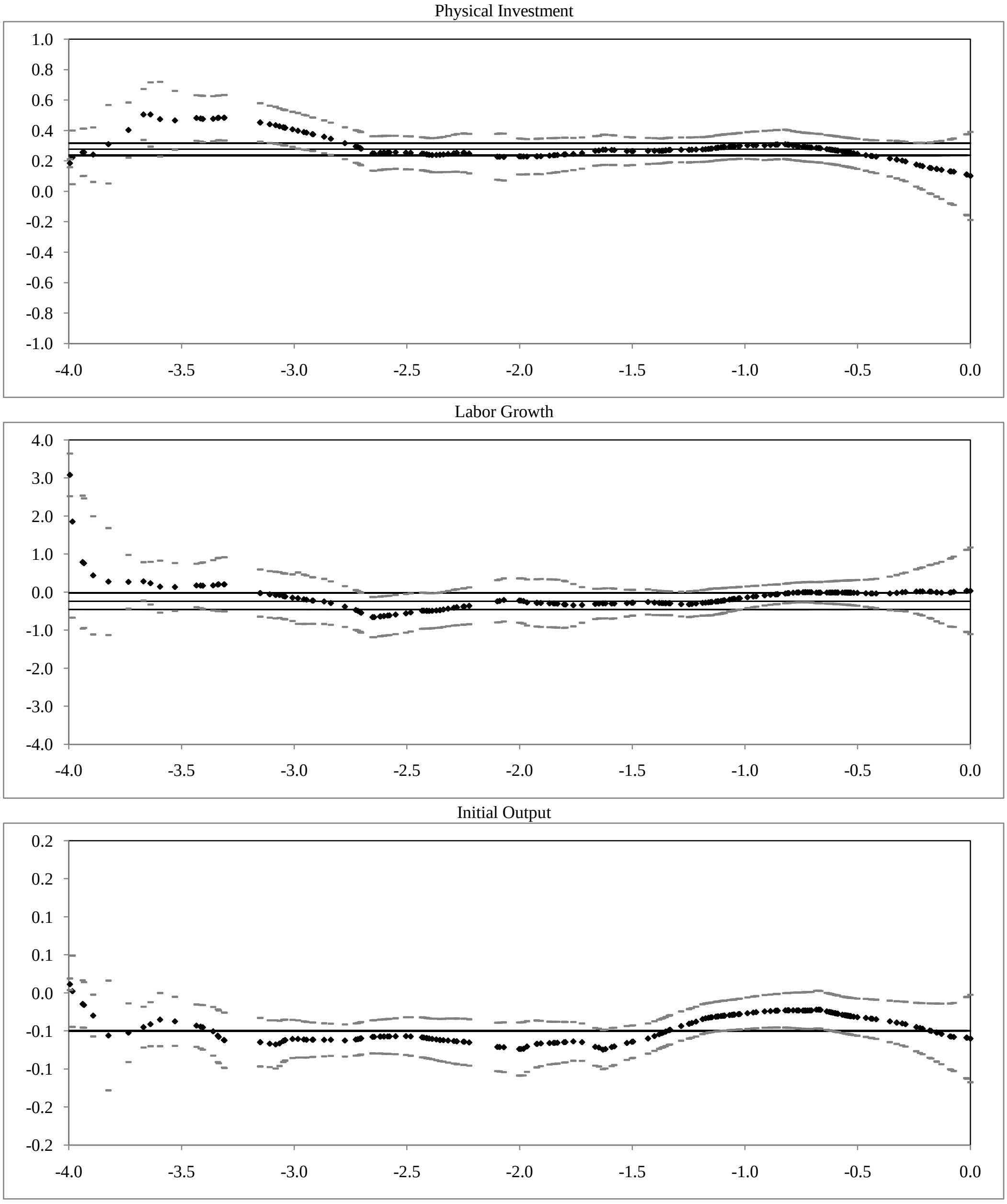

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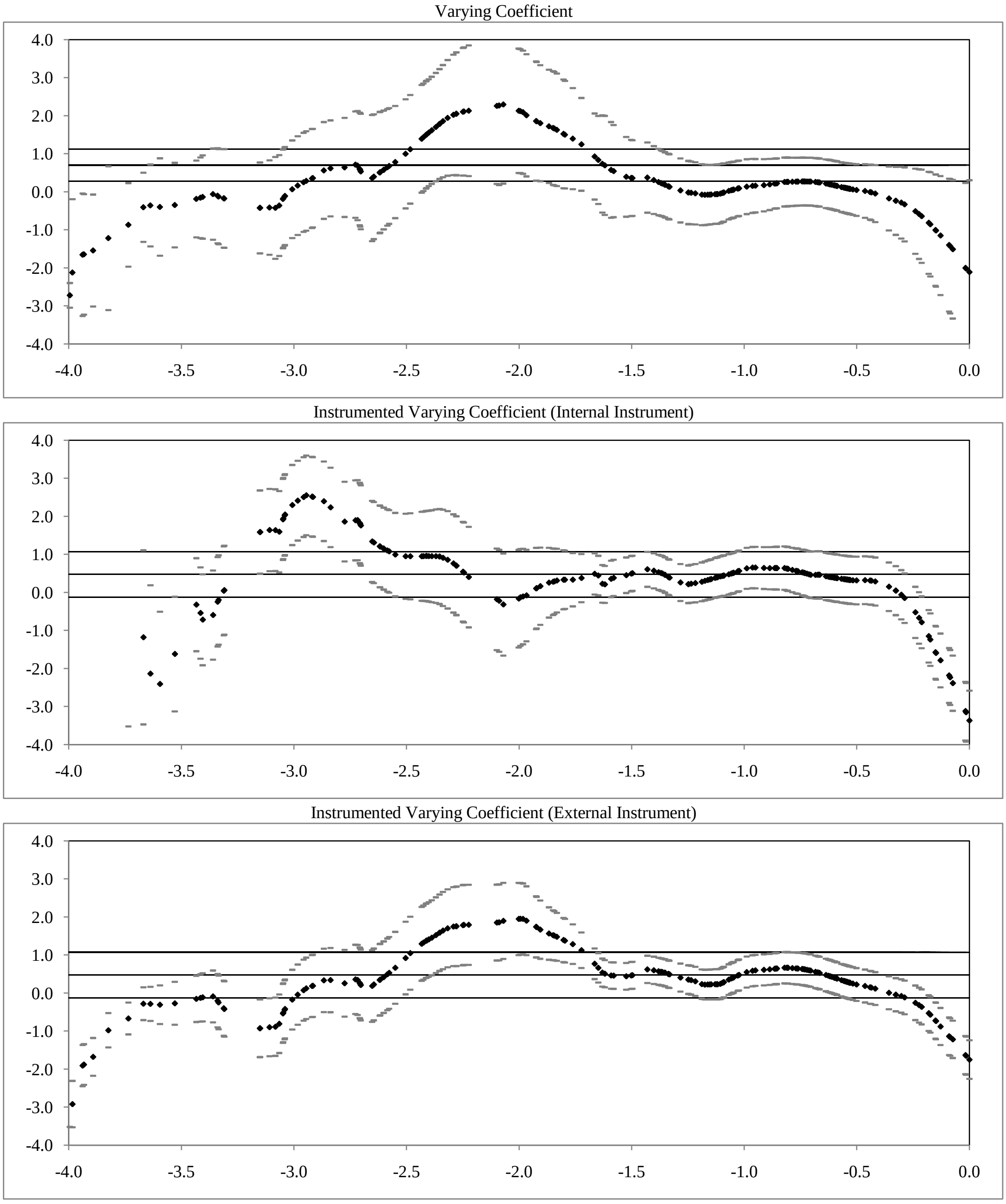
Figure 3.7: Varying Return of R&D Investment (controlling by Labor Quality)



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3.8 Conclusions

When mapping the contribution of R&D to economic growth in a cross country environment: (i) Acknowledgment of *heterogeneity* matters: clusters, clubs, regions, types or any other qualification denoting heterogenous effects for heterogeneous observations play an important role to understand the contributions of innovation/imitation to growth. Our findings suggest the existence of an inverted U relationship between the level of economic development of a country and the contribution of investments in R&D to its economic growth. (ii) Acknowledgment of *endogeneity* matters: endogeneity bias originated by reverse causation of the innovation variable pushes the VC estimates downwards for developed economies. Our IVVC estimates cope with this issue by exploiting internal and external instruments. (iii) Evidence supports the existence of *complementarity effects* between productive factors: even when all countries have equal access to new technologies or comparable amounts of R&D investments, the technology-skill mismatch in productive factors can lead to sizeable differences in the returns to innovation efforts, we experiment by adjusting the labor factor by educational attainment and partially cope with this issue.

Yet it is not our aim to prescribe optimal policies, our findings may have some policy implications. These implications are related to the acknowledgement of the differentiated potential gains from R&D investments according to the particular initial conditions that each country has. Our results would suggest that efforts coming from governments aimed at promoting this kind of investments are expected to yield better outcomes in contexts with highly complementary factors (i.e. educated labor force) and a sector of technology developers/innovators (absorbers/imitators) of low (high) density. It becomes apparent that for countries that lie too far from the economic or technological frontier, government priorities should focus more intensively into improving the quality of human capital rather than into devoting resources to prompt technological innovations that would not be fully exploited due to the lack of complementarity effects.

3.9 Appendix

3.9.1 Two Steps in IVVC estimation

First Stage

Assuming that $\pi_j(.)$ has a continuous second derivative, for $\mathscr{Z}_k$ in a neighborhood of $\mathscr{Z}_i$, a Taylor expansion can approximate $\pi_j(\mathscr{Z})_k \approx \alpha_{ij} + (\mathscr{Z}_k - \mathscr{Z}_i)'\beta_{ij}$ where $\alpha_{ij} = \pi_j(\mathscr{Z}_i)$ and $\beta_{ij} = \pi'_{ij}(\mathscr{Z}_i)$. By jackknife sampling, we use all observations except the ith observation in estimating $\pi_j(\mathscr{Z}_i)$. We get a typical local weighted least squares estimator as follows:

$$\sum_{k\neq i}^{n}\{X_{kj} - \alpha_{ij} - (\mathscr{Z}_k - \mathscr{Z}_i)'\beta_{ij}\}^2 K_{h_1}(\mathscr{Z}_k - \mathscr{Z}_i)$$

where $K_{h_1}(.) = K(./h_1)/h_1^m + l$. $K(.)$ is a kernel function in $\Re^m + l$, and $h_1 = h_{1n} > 0$ is the bandwidth in the first step that controls the degree of smoothing in estimation. Minimizing the last expression with respect to α_{ij} and β_{ij} gives the local linear estimate of $\pi_j(\mathscr{Z}_i)$:

$$\hat{\pi}_{j,-i}(\mathscr{Z}_i) = \hat{\alpha_{ij}} = \mathbf{e}_1'(\tilde{\mathscr{Z}_i}'\mathbf{W}_i\tilde{\mathscr{Z}_i})^{-1}\tilde{\mathscr{Z}_i}\mathbf{W}_i\tilde{\mathbf{X}}_j \tag{3.9}$$

where
$\mathbf{e}_1' = (1, 0, ..., 0)$, for $1 \leq i \leq n$ and $1 \leq j \leq d$,

$$\tilde{\mathscr{Z}_i} = \begin{pmatrix} 1 & (\mathscr{Z}_1 - \mathscr{Z}_i)' \\ \vdots & \vdots \\ 1 & (\mathscr{Z}_n - \mathscr{Z}_i)' \end{pmatrix}, \qquad \tilde{\mathbf{X}}_j = \begin{pmatrix} X_{1j} \\ \vdots \\ X_{nj} \end{pmatrix}$$

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and
$\mathbf{W}_i = diag\{K_{h_1}(\mathscr{Z}_1 - \mathscr{Z}_i), ..., K_{h_1}(\mathscr{Z}_n - \mathscr{Z}_i)\}$

Second Stage

Assuming that the functions $\{g_j(\mathbf{Z}_1)\}$ have a continuous second derivative at any given point $\mathbf{z}_1$, a Taylor expansion can approximate: $g_j(\mathbf{Z}_1) \approx b_j + (\mathbf{Z}_1 - \mathbf{z}_1)'\mathbf{c}_j$ where $b_j = g_j(\mathbf{z}_1)$ and $\mathbf{c}_j = \partial g_j(\mathbf{z}_1)/\partial \mathbf{z}_1$. Let

$$\mathbf{\Pi} = \begin{pmatrix} \pi_1' & \pi_1' \otimes (\mathbf{Z}_{11} - \mathbf{z}_1)' \\ \vdots & \vdots \\ \pi_n' & \pi_n' \otimes (\mathbf{Z}_{n1} - \mathbf{z}_1)' \end{pmatrix},$$

where $\otimes$ is the Kronecker product. Then the conditional mean in model (2) can be approximated by $E[\mathbf{Y}|\mathscr{Z}_1, ..., \mathscr{Z}_n] \approx \mathbf{\Pi\Theta}$, where

$\mathbf{Y} = \begin{pmatrix} y_1 \\ \vdots \\ y_n \end{pmatrix}$ and $\mathbf{\Theta} = \mathbf{\Theta}(\mathbf{z}_1) = \begin{pmatrix} \mathbf{b} \\ \mathbf{c} \end{pmatrix}$ with $\mathbf{b} = \begin{pmatrix} b_0 \\ \vdots \\ b_d \end{pmatrix}$ and $\mathbf{c} = \begin{pmatrix} \mathbf{c}_0 \\ \vdots \\ \mathbf{c}_d \end{pmatrix}$.

The local linear estimator $\hat{\mathbf{\Theta}}$ is defined as the minimizer of:

$$\sum_{i=1}^{n}[Y_i - \sum_{j=0}^{d}\{b_j + (\mathbf{Z}_{i1} - \mathbf{z}_1)'\mathbf{c}_j\}\hat{\pi}_{j,-i}(\mathscr{Z}_i)]^2 L_{h_2}(\mathbf{Z}_{i1} - \mathbf{z}_1) \tag{3.10}$$

,

where $L_{h_2}(.) = L(./h_2)/h_2^m$. $L(.)$ is a kernel function in $\Re^m$, and $h_2 = h_{2n} > 0$ is the bandwidth in the second step that controls the degree of smoothing in estimation. Minimizing the last expression with respect to $\mathbf{b}$ and $\{\mathbf{c}_j\}$ gives the local linear estimate of $b_j(\mathbf{z}_1)$ and $\mathbf{c}_j(\mathbf{z}_1)$.Therefore the minimizers of (4) are given by

$$\hat{\mathbf{\Theta}} = \mathbf{H}_2^{-1}\hat{\mathbf{S}}_n^{-1}\hat{\mathbf{T}}_n$$

,

where $\mathbf{H}_2 = diag\{1, ..., 1, h_2, ..., h_2\}$ is a $(d+1)(m+1) \times (d+1)(m+1)$ matrix with the first $(d+1)$ diagonal elements being 1's and the rest diagonal elements h_2's,

$$\hat{\mathbf{S}}_n = \hat{\mathbf{S}}_n(\mathbf{z}_1) = \begin{pmatrix} \hat{\mathbf{S}}_{n,0} & \hat{\mathbf{S}}_{n,1}' \\ \hat{\mathbf{S}}_{n,1} & \hat{\mathbf{S}}_{n,2} \end{pmatrix} \quad \text{and} \quad \hat{\mathbf{T}}_n = \hat{\mathbf{T}}_n(\mathbf{z}_1) = \begin{pmatrix} \hat{\mathbf{T}}_{n,0}(\mathbf{z}_1) \\ \hat{\mathbf{T}}_{n,1}(\mathbf{z}_1) \end{pmatrix}.$$

Further,

$\hat{\mathbf{S}}_{n,0} = (1/n)\sum_{i=1}^{n} L_{h_2}(\mathbf{Z}_{i1} - \mathbf{z}_1)\hat{\pi}_{-i}\hat{\pi}_{-i}'$,
$\hat{\mathbf{S}}_{n,1} = (1/nh_2)\sum_{i=1}^{n} L_{h_2}(\mathbf{Z}_{i1} - \mathbf{z}_1)\{\hat{\pi}_{-i} \otimes (\mathbf{Z}_{i1} - \mathbf{z}_1)\}\hat{\pi}_{-i}'$,
$\hat{\mathbf{S}}_{n,2} = (1/nh_2^2)\sum_{i=1}^{n} L_{h_2}(\mathbf{Z}_{i1} - \mathbf{z}_1)\{\hat{\pi}_{-i} \otimes (\mathbf{Z}_{i1} - \mathbf{z}_1)\}\{\hat{\pi}_{-i} \otimes (\mathbf{Z}_{i1} - \mathbf{z}_1)\}'$,
$\hat{\mathbf{T}}_{n,0} = (1/n)\sum_{i=1}^{n} L_{h_2}(\mathbf{Z}_{i1} - \mathbf{z}_1)\hat{\pi}_{-i}Y_i$ and
$\hat{\mathbf{T}}_{n,1} = (1/nh_2)\sum_{i=1}^{n} L_{h_2}(\mathbf{Z}_{i1} - \mathbf{z}_1)\{\hat{\pi}_{-i} \otimes (\mathbf{Z}_{i1} - \mathbf{z}_1)\}Y_i$.

That gives the two-stage local linear estimated of the coefficient functions and in particular,

$\hat{\mathbf{g}}(\mathbf{z}_1) = \hat{\mathbf{b}} = \mathbf{e}'\mathbf{H}_2^{-1}\hat{\mathbf{S}}_n^{-1}\hat{\mathbf{T}}_n = \mathbf{e}'\hat{\mathbf{S}}_n^{-1}\hat{\mathbf{T}}_n$,

where is $\mathbf{e} = (1, ..., 1, 0, ..., 0)'$ is a $(d+1)(m+1)$ vector whose first $(d+1)$ elements are one and the remaining elements are zero.

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3.9.2 Optimal Bandwidth Estimation

The optimal variable bandwidth for the varying coefficient model is given by:

$$\hat{\mathbf{h}}_{opt} = Argmin_h \widehat{MSE}(\hat{\gamma}(\mathbf{z})|\mathscr{X})$$

where $\widehat{MSE}(\hat{\gamma}(\mathbf{z})|\mathscr{X})$ is a good estimator of the mean squared error $MSE(\hat{\gamma}(z)|\mathscr{X})$ defined by

$$MSE(\hat{\gamma}(z)|\mathscr{X}) = \mathbf{b}'(z)\Omega(z)\mathbf{b}(z) + tr(\Omega(z)Cov(z)|\mathscr{X}))$$

where $\mathbf{b}(z) = bias((\hat{\gamma}_1(\mathbf{z})|\mathscr{X}), ..., bias((\hat{\gamma}_p(\mathbf{z})|\mathscr{X}))$ and Ω is a matrix with $(i,j)th$ elements equal to $r_{ij}(z)$ with $r_{ij}(z) = E(X_iX_j|z=z)$, for $i,j = 1,2,...,p$.

Bias

Based on the Taylor expansion of order m, the conditional bias can be approximated by $(I_p \otimes e'_{1,q})(X'_qWX_q)^{-1}X'_qW\eta$, where η is a n vector with ith element equal to

$$\sum_{j=1}^{p}\sum_{k=1}^{m}(c_{j,q+k}(z_i - z_0)^{q+k})X_{ij}$$

For convenience, we take $m = 2$; then $(\mathbf{I_p} \otimes \mathbf{e'_{1,q}})(\mathbf{X'_q W X_q})^{-1}\mathbf{X'_q W}\eta$ is simplified as

$$(\mathbf{I_p} \otimes \mathbf{e'_{1,q}})(\mathbf{X'_q W X_q})^{-1}\mathbf{X'_q W X^*_q s}$$

where

$$s = (c'_1 \otimes (1,0) + c'_2 \otimes (0,1))'$$

$$c_i = (c_{1,q+i}, ..., c_{p,q+i})', \text{ for } i = 1,2$$

and

$$\mathbf{X}^*_q = \begin{pmatrix} (z_1 - z_0)^{q+1}X_{11} & (z_1 - z_0)^{q+2}X_{11} & \dots & (z_1 - z_0)^{q+1}X_{1p} & (z_1 - z_0)^{q+2}X_{1p} \\ \vdots & \vdots & \ddots & \vdots & \vdots \\ (z_1 - z_0)^{q+1}X_{n1} & (z_n - z_0)^{q+2}X_{n1} & \dots & (z_1 - z_0)^{q+1}X_{np} & (z_n - z_0)^{q+2}X_{np} \end{pmatrix}$$

The quantity s can be estimated by using a local polynomial regression of order $g(g > q)$ with a bandwidth h_*:

$$\hat{\mathbf{s}} = (\mathbf{I_p} \otimes (\mathbf{e_{q+2,g}}, \mathbf{e_{q+3,g}})')(\mathbf{X'_g W_* X_g})^{-1}\mathbf{X'_g W_* y}$$

where $W_* = diag(K_{h_*}(z_1 - z_0), ..., K_{h_*}(z_n - z_0))$. The initial bandwidth h_* is obtained by the minimizer of the integrated residual squares criterion (IRSC) as in Fan & Gijbels (1995)[30].

[30] $h_* = Argmin(IRSC(h)) = \int RSC(z_0;h)dz_0$,
where $RSC(z_0;h) = (1 + (q+1)V)\frac{\sum_{i=1}^{n}(y_i - \hat{y_h}_i)^2 K_h(z_i - z_0)}{tr\{W_h - W_hX(X'W_hX)^{-1}X'W_h\}}$ and V is the first diagonal element of the matrix $(X'W_hX)^{-1}(X'W_h^2X)(X'W_hX)^{-1}$

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Covariance

The conditional covariance is given by

$$Cov(\hat{\gamma}(z)|\mathscr{X}) = (\mathbf{I_p} \otimes \mathbf{e}'_{\mathbf{1,q}})(\mathbf{X}'_{\mathbf{q}}\mathbf{W}\mathbf{X}_{\mathbf{q}})^{-\mathbf{1}}(\mathbf{X}'_{\mathbf{q}}\mathbf{W}\mathbf{\Psi}\mathbf{W}\mathbf{X}_{\mathbf{q}})(\mathbf{X}'_{\mathbf{q}}\mathbf{W}\mathbf{X}_{\mathbf{q}})^{-\mathbf{1}}(\mathbf{I_p} \otimes \mathbf{e}'_{\mathbf{1,q}})$$

where $\Psi = diag(\sigma^2(z_1), ..., \sigma^2(z_n))$. Under local homoscedasticity the covariance can be approximated by:

$$Cov(\hat{\gamma}(z)|\mathscr{X}) = (\mathbf{I_p} \otimes \mathbf{e}'_{\mathbf{1,q}})(\mathbf{X}'_{\mathbf{q}}\mathbf{W}\mathbf{X}_{\mathbf{q}})^{-\mathbf{1}}(\mathbf{X}'_{\mathbf{q}}\mathbf{W}^{\mathbf{2}}\mathbf{X}_{\mathbf{q}})(\mathbf{X}'_{\mathbf{q}}\mathbf{W}\mathbf{X}_{\mathbf{q}})^{-\mathbf{1}}(\mathbf{I_p} \otimes \mathbf{e}'_{\mathbf{1,q}})$$

where σ^2 can be estimated by the normalized weighted residual sum of squares from a gth order polynomial fit as:

$$\hat{\sigma}^2 = \frac{\sum_{i=1}^{n}(y_i - \hat{y}_{*i})^2 K_h(z_i - z_0)}{tr\{\mathbf{W}_* - \mathbf{W}_*\mathbf{X_g}(\mathbf{X}'_{\mathbf{g}}\mathbf{W}_*\mathbf{X_g})^{-\mathbf{1}}\mathbf{X}'_{\mathbf{g}}\mathbf{W}_*\}} \tag{3.11}$$

where $\hat{y}_{*i} = (\hat{y}_{*1}, ..., \hat{y}_{*n})' = \mathbf{X_g}(\mathbf{X}'_{\mathbf{g}}\mathbf{W}_*\mathbf{X_g})^{-\mathbf{1}}\mathbf{X}'_{\mathbf{g}}\mathbf{W}_*\mathbf{y}$

Omega

The estimate of the element r_{ij} in Ω will be obtained based on $(Z_l, X_{li}X_{lj}), l = 1, ..., n$, using the local polynomial fit of order g with bandwidth $h_{0*} = Op(h_*)$:

$$\hat{r}_{ij} = \mathbf{e}'_{\mathbf{1,g}}(\mathbf{V}'\mathbf{W}_{\mathbf{0}*}\mathbf{V})^{-\mathbf{1}}\mathbf{V}'\mathbf{W}_{\mathbf{0}*}\mathbf{U_{ij}}$$

where $W_{0*} = diag(K_{h_{0*}}(z_1 - z_0), ..., K_{h_{0*}}(z_n - z_0))$ and

$$V = \begin{pmatrix} 1 & \dots & (z_1 - z_0)^g \\ \vdots & \ddots & \vdots \\ 1 & \dots & (z_n - z_0)^g \end{pmatrix}, \qquad U_{ij} = \begin{pmatrix} X_{1i}X_{1j} \\ \vdots \\ X_{ni}X_{nj} \end{pmatrix}$$

The estimates of the three elements combined (bias, covariance and omega), considering $q = 1$ and $g = 3$ allow us to estimate the Mean Squared Errors as:

$$\widehat{MSE}(\hat{\gamma}(z)|\mathscr{X}) =$$
$$(\hat{\mathbf{s}}'\mathbf{X}_{\mathbf{1}}^{*\prime}\mathbf{W}\mathbf{X_1}(\mathbf{X}'_{\mathbf{1}}\mathbf{W}\mathbf{X_1})^{-\mathbf{1}})\hat{\mathbf{\Omega}}((\mathbf{X}'_{\mathbf{1}}\mathbf{W}\mathbf{X_1})^{-\mathbf{1}}\mathbf{X}'_{\mathbf{1}}\mathbf{W}\mathbf{X}^*_{\mathbf{1}}\hat{\mathbf{s}}) + \mathbf{tr}((\mathbf{X}'_{\mathbf{1}}\mathbf{W}\mathbf{X_1})^{-\mathbf{1}}(\mathbf{X}'_{\mathbf{1}}\mathbf{W}^{\mathbf{2}}\mathbf{X_1})(\mathbf{X}'_{\mathbf{1}}\mathbf{W}\mathbf{X_1})^{-\mathbf{1}}\hat{\mathbf{\Omega}}\hat{\sigma}^{\mathbf{2}}_*)$$

Given the weighting matrix $\mathbf{e}'_{\mathbf{1,g}}(\mathbf{V}'\mathbf{W}_{\mathbf{0}*}\mathbf{V})^{-\mathbf{1}}\mathbf{V}'\mathbf{W}_{\mathbf{0}*}$, $\hat{r}_{ij}$ could be negative semi-definite (leading to a negative semi-definite Ω matrix and therefore to potential non-positive MSE) we use the covariance matrix repairing algorithm proposed by Dong & Yaoi (2007). To repair ill-posed covariance matrices, the algorithm adds a positive value to the diagonal of Ω. The value we chose is the absolute value of the minimum eigenvalue (which is negative if Ω is ill-posed) of Ω, multiplies a self-adapted positive coefficient of 1.5 (as in Dong & Yaoi 2007). Alternatively we also computed $\hat{r}_{ij}$ as $\hat{r}_{ij} = |\mathbf{e}'_{\mathbf{1,g}}(\mathbf{V}'\mathbf{W}_{\mathbf{0}*}\mathbf{V})^{-\mathbf{1}}\mathbf{V}'\mathbf{W}_{\mathbf{0}*}|\mathbf{U_{ij}}$. Both approaches yield virtually the same optimal variable bandwidths (except for the cases of extreme left values of the ordering variable in which the repair method cannot yield reliable outputs).

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3.9.3 Depreciated (net) investment in physical capital and R&D.

Starting from gross values of R&D expenditures, we get "net" values, by applying the perpetual inventory method with a pre-sample growth rate of 0.05 as in Hall & Mairesse (1995). However, we depart from Hall and Mairesse in two regards. First, we work with average data by quinquennium rather than by year and therefore we assume constant expenditures within quinquennia. Second, we use a depreciation rate of 0.05 instead of 0.15[31]. Though not ideal, this correction gives us an idea of the order of magnitude of differences between the contributions of gross and net R&D investments.

Based on the following equation:

$$S_t = (1-\delta)S_{t-1} + R_{t-1}$$

where S_t is the capital stock of R&D, R_t is the gross flow, and δ is the depreciation rate, it is easy to show that:

$$Rnet_t = R_t - \delta R_{t-1} - \delta(1-\delta)R_{t-2} - \delta(1-\delta)^2 R_{t-3} - \delta(1-\delta)^3 R_{t-4} - ...$$

and hence we can incorporate depreciation issues without the need of generating stocks. With the assumption of a constant R_t within quinquennium we have:

$R_t = Rq$ for $t = t, ..., t-4$

$Rnet_t = Rq_j(1-\delta-\delta(1-\delta)-\delta(1-\delta)^2-\delta(1-\delta)^3) - Rq_{j-1}(\delta(1-\delta)^4+...) - ... - Rq_1(\delta(1-\delta)^{n-4}+\delta(1-\delta)^{n-3}+\delta(1-\delta)^{n-2}+\delta(1-\delta)^{n-1}+\delta(1-\delta)^n\frac{1}{g+\delta})$

where $Rnet_t$ is the variable in net terms, Rq_1 is the observation of the first quinquennium, and g is the pre-sample growth rate.

Finally, we calculate this equation for every period within a quinquennium and calculate the average of the coefficient multiplying each Rq_j

$$Rnet_{t-1} = Rq_j(1-\delta-\delta(1-\delta)-\delta(1-\delta)^2) - Rq_{j-1}(\delta(1-\delta)^3+...) - ...$$

$$Rnet_{t-2} = Rq_j(1-\delta-\delta(1-\delta)) - Rq_{j-1}(\delta(1-\delta)^2+...) - ...$$

$$\begin{aligned} Rnetq &= Rq_j(1-\tfrac{4}{5}\delta-\tfrac{3}{5}\delta(1-\delta)-\tfrac{2}{5}\delta(1-\delta)^2-\tfrac{1}{5}\delta(1-\delta)^3) \\ &-Rq_{j-1}\begin{pmatrix} \frac{1}{5}\delta+\frac{2}{5}\delta(1-\delta)+\frac{3}{5}\delta(1-\delta)^2+\frac{4}{5}\delta(1-\delta)^3+\delta(1-\delta)^4+ \\ +\frac{4}{5}\delta(1-\delta)^5+\frac{3}{5}\delta(1-\delta)^6+\frac{2}{5}\delta(1-\delta)^7+\frac{1}{5}\delta(1-\delta)^8 \end{pmatrix} - ... \end{aligned}$$

[31] With a depreciation rate of 0.15, average net R&D investments amounts one third of their gross values whereas they amount one half when $\delta = 0.05$

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3.9.4 Values of Inducing Variables

Distance to the Economic Frontier

Country	Period	Distance to Economic Frontier	Country	Period	Distance to Economic Frontier	Country	Period	Distance to Economic Frontier	Country	Period	Distance to Economic Frontier	Country	Period	Distance to Economic Frontier
MWI	1966-1970	-5.98	MUS	1981-1985	-3.33	GRC	1966-1970	-2.31	ITA	1981-1985	-1.24	CAN	1996-2000	-0.61
IND	1966-1970	-5.56	GTM	1981-1985	-3.31	URY	1971-1975	-2.29	BEL	1971-1975	-1.22	FRA	1996-2000	-0.60
IND	1976-1980	-5.32	ECU	1996-2000	-3.31	MYS	1996-2000	-2.28	CYP	1991-1995	-1.22	BEL	1991-1995	-0.60
PAK	1966-1970	-5.29	ECU	1991-1995	-3.26	HUN	1996-2000	-2.28	AUT	1971-1975	-1.21	BEL	1996-2000	-0.60
UGA	1991-1995	-5.28	IRN	1991-1995	-3.19	HUN	1991-1995	-2.25	SWE	1966-1970	-1.21	BEL	1986-1990	-0.59
IND	1981-1985	-5.28	TUR	1971-1975	-3.19	ZAF	1986-1990	-2.25	ISR	1961-1965	-1.20	GBR	1991-1995	-0.59
IND	1971-1975	-5.26	TUR	1981-1985	-3.15	BRA	1986-1990	-2.22	AUS	1976-1980	-1.19	FRA	1991-1995	-0.59
UGA	1996-2000	-5.12	JOR	1981-1985	-3.15	PAN	1986-1990	-2.17	FRA	1971-1975	-1.18	NOR	1961-1965	-0.57
KEN	1966-1970	-5.10	TUN	1991-1995	-3.13	PRT	1971-1975	-2.13	AUS	1971-1975	-1.17	NLD	1991-1995	-0.57
PAK	1976-1980	-5.02	CHL	1976-1980	-3.12	MEX	1981-1985	-2.12	FIN	1976-1980	-1.16	FRA	1986-1990	-0.56
BGD	1971-1975	-5.01	THA	1991-1995	-3.11	ARG	1966-1970	-2.11	ESP	1986-1990	-1.15	AUT	1991-1995	-0.55
PAK	1971-1975	-4.98	GTM	1986-1990	-3.08	CHL	1996-2000	-2.10	AUS	1981-1985	-1.15	DEU	1996-2000	-0.55
TGO	1991-1995	-4.89	TUR	1976-1980	-3.07	VEN	1981-1985	-2.09	NZL	1991-1995	-1.13	NLD	1986-1990	-0.55
PAK	1981-1985	-4.88	TUN	1996-2000	-3.05	PRT	1976-1980	-2.07	FRA	1976-1980	-1.12	NLD	1996-2000	-0.54
BGD	1991-1995	-4.80	ROM	1991-1995	-3.05	HUN	1986-1990	-2.00	BEL	1976-1980	-1.12	DEU	1991-1995	-0.54
ZMB	1996-2000	-4.78	PER	1991-1995	-3.04	VEN	1996-2000	-1.99	AUT	1976-1980	-1.12	AUT	1986-1990	-0.54
LKA	1971-1975	-4.77	SLV	1991-1995	-3.04	MEX	1996-2000	-1.98	GBR	1971-1975	-1.12	CHE	1976-1980	-0.53
IND	1986-1990	-4.72	SLV	1971-1975	-3.02	PRT	1981-1985	-1.97	CAN	1971-1975	-1.11	GBR	1986-1990	-0.53
THA	1966-1970	-4.71	ROM	1996-2000	-3.01	TTO	1996-2000	-1.95	ESP	1991-1995	-1.10	AUT	1996-2000	-0.53
IND	1991-1995	-4.71	BRA	1971-1975	-3.00	CYP	1981-1985	-1.93	ESP	1996-2000	-1.09	GBR	1996-2000	-0.53
TGO	1986-1990	-4.69	KOR	1971-1975	-2.98	MEX	1991-1995	-1.92	GBR	1976-1980	-1.09	FIN	1986-1990	-0.52
LKA	1981-1985	-4.59	PER	1981-1985	-2.96	URY	1991-1995	-1.92	GBR	1981-1985	-1.09	CAN	1986-1990	-0.50
CHN	1986-1990	-4.57	SLV	1986-1990	-2.95	VEN	1976-1980	-1.92	BEL	1981-1985	-1.07	HKG	1996-2000	-0.47
IND	1996-2000	-4.54	PER	1971-1975	-2.95	VEN	1991-1995	-1.91	NLD	1971-1975	-1.06	CHE	1981-1985	-0.46
SEN	1996-2000	-4.53	PER	1976-1980	-2.95	ISR	1966-1970	-1.90	FRA	1981-1985	-1.06	SWE	1996-2000	-0.44
SEN	1991-1995	-4.47	JAM	1981-1985	-2.93	VEN	1971-1975	-1.87	FIN	1981-1985	-1.05	CHE	1971-1975	-0.44
CMR	1966-1970	-4.47	SLV	1996-2000	-2.92	ARG	1981-1985	-1.87	NZL	1996-2000	-1.05	SWE	1991-1995	-0.42
ZMB	1966-1970	-4.45	PER	1996-2000	-2.92	ARG	1976-1980	-1.85	AUT	1981-1985	-1.05	DNK	1991-1995	-0.36
KEN	1961-1965	-4.44	COL	1996-2000	-2.87	MEX	1986-1990	-1.84	CAN	1976-1980	-1.05	USA	1961-1965	-0.33
CHN	1991-1995	-4.41	CHL	1981-1985	-2.87	ITA	1966-1970	-1.83	NLD	1976-1980	-1.03	SWE	1986-1990	-0.33
ZMB	1971-1975	-4.32	KOR	1961-1965	-2.86	KOR	1986-1990	-1.83	NLD	1981-1985	-1.03	ISL	1996-2000	-0.33
GHA	1961-1965	-4.31	THA	1996-2000	-2.84	URY	1996-2000	-1.80	IRL	1986-1990	-1.02	DNK	1996-2000	-0.30
PAK	1986-1990	-4.30	HUN	1971-1975	-2.80	VEN	1986-1990	-1.80	USA	1966-1970	-1.02	ISL	1991-1995	-0.29
PAK	1996-2000	-4.25	CRI	1971-1975	-2.79	ARG	1971-1975	-1.76	CAN	1981-1985	-0.99	DNK	1986-1990	-0.24
PHL	1966-1970	-4.25	KOR	1976-1980	-2.78	HKG	1961-1965	-1.75	ISL	1971-1975	-0.96	USA	1991-1995	-0.22
THA	1976-1980	-4.19	BRA	1976-1980	-2.73	GRC	1961-1965	-1.75	NOR	1971-1975	-0.95	NOR	1991-1995	-0.21
TUN	1966-1970	-4.16	MUS	1986-1990	-2.72	ESP	1971-1975	-1.73	JPN	1961-1965	-0.92	USA	1996-2000	-0.18
EGY	1976-1980	-4.14	CRI	1981-1985	-2.71	ARG	1991-1995	-1.72	NZL	1986-1990	-0.91	ISL	1986-1990	-0.17
THA	1961-1965	-4.08	CRI	1976-1980	-2.71	IRL	1971-1975	-1.69	IRL	1991-1995	-0.91	CHE	1996-2000	-0.15
GUY	1981-1985	-4.00	BRA	1981-1985	-2.70	GRC	1976-1980	-1.68	SWE	1971-1975	-0.89	USA	1986-1990	-0.15
THA	1981-1985	-3.99	MEX	1966-1970	-2.70	GRC	1981-1985	-1.66	ISL	1976-1980	-0.86	NOR	1986-1990	-0.13
PHL	1971-1975	-3.98	MYS	1986-1990	-2.68	ESP	1981-1985	-1.65	ISR	1991-1995	-0.86	NOR	1996-2000	-0.09
IDN	1991-1995	-3.97	JOR	1986-1990	-2.65	FIN	1966-1970	-1.64	SWE	1981-1985	-0.86	JPN	1986-1990	-0.09
CHN	1996-2000	-3.94	TUR	1991-1995	-2.65	IRL	1976-1980	-1.63	SWE	1976-1980	-0.85	CHE	1991-1995	-0.07
EGY	1981-1985	-3.93	ROM	1986-1990	-2.63	ESP	1976-1980	-1.62	NOR	1976-1980	-0.82	JPN	1996-2000	-0.02
LKA	1996-2000	-3.93	PER	1986-1990	-2.62	ARG	1996-2000	-1.59	ISR	1986-1990	-0.81	JPN	1991-1995	-0.01
PHL	1976-1980	-3.89	HUN	1976-1980	-2.62	KOR	1991-1995	-1.58	JPN	1971-1975	-0.81	CHE	1986-1990	0.00
PHL	1981-1985	-3.82	TUR	1986-1990	-2.60	ARG	1986-1990	-1.52	DNK	1981-1985	-0.81	NOR	1986-1990	-0.13
GTM	1966-1970	-3.75	KOR	1981-1985	-2.60	PRT	1986-1990	-1.50	DNK	1971-1975	-0.79	NOR	1996-2000	-0.09
IDN	1996-2000	-3.74	JAM	1986-1990	-2.59	IRL	1981-1985	-1.50	DNK	1976-1980	-0.79	JPN	1986-1990	-0.09
ECU	1971-1975	-3.73	TUR	1996-2000	-2.58	ITA	1971-1975	-1.43	USA	1971-1975	-0.77	CHE	1991-1995	-0.07
PHL	1991-1995	-3.67	MYS	1991-1995	-2.55	CAN	1966-1970	-1.41	ISR	1996-2000	-0.76	JPN	1996-2000	-0.02
BOL	1991-1995	-3.66	ZAF	1981-1985	-2.55	JPN	1966-1970	-1.41	CHE	1966-1970	-0.76	JPN	1991-1995	-0.01
BOL	1996-2000	-3.64	HUN	1981-1985	-2.50	ISR	1971-1975	-1.40	AUS	1991-1995	-0.76	CHE	1961-1965	0.00
PHL	1986-1990	-3.60	POL	1991-1995	-2.49	GRC	1996-2000	-1.38	USA	1976-1980	-0.74	CHE	1986-1990	0.00
ECU	1976-1980	-3.53	URY	1966-1970	-2.49	NZL	1971-1975	-1.38	JPN	1976-1980	-0.74			
IRN	1981-1985	-3.52	CHL	1986-1990	-2.48	NZL	1976-1980	-1.37	ITA	1996-2000	-0.73			
KOR	1966-1970	-3.51	MUS	1996-2000	-2.44	PRT	1991-1995	-1.36	ITA	1991-1995	-0.73			
COL	1971-1975	-3.48	ZAF	1991-1995	-2.43	PRT	1996-2000	-1.35	ITA	1986-1990	-0.72			
SYR	1996-2000	-3.48	BRA	1991-1995	-2.42	ISR	1976-1980	-1.35	IRL	1996-2000	-0.71			
JOR	1976-1980	-3.48	CRI	1991-1995	-2.41	GRC	1991-1995	-1.34	NOR	1981-1985	-0.71			
PHL	1961-1965	-3.46	MEX	1971-1975	-2.41	ITA	1976-1980	-1.34	USA	1981-1985	-0.68			
GTM	1971-1975	-3.43	PAN	1991-1995	-2.40	KOR	1996-2000	-1.33	AUS	1996-2000	-0.67			
COL	1976-1980	-3.42	TTO	1966-1970	-2.40	VEN	1961-1965	-1.30	ISL	1981-1985	-0.67			
HND	1961-1965	-3.41	BRA	1996-2000	-2.39	ARG	1961-1965	-1.29	AUS	1986-1990	-0.67			
THA	1986-1990	-3.41	CHL	1991-1995	-2.37	ISR	1981-1985	-1.28	FIN	1996-2000	-0.64			
EGY	1991-1995	-3.40	CRI	1986-1990	-2.35	NOR	1966-1970	-1.28	GBR	1961-1965	-0.63			
GTM	1976-1980	-3.36	PAN	1996-2000	-2.35	FIN	1971-1975	-1.25	CAN	1991-1995	-0.63			
EGY	1996-2000	-3.34	CRI	1996-2000	-2.34	GRC	1986-1990	-1.25	JPN	1981-1985	-0.62			
COL	1981-1985	-3.33	POL	1996-2000	-2.32	ISL	1966-1970	-1.24	FIN	1991-1995	-0.62			

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Chapter 4

Fiscal redistribution and income inequality in Latin America[1]

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Abstract

This paper documents and compares the redistributive performance of Latin American and Western European fiscal systems. It finds that (i) taxes and transfers widen the difference in income inequality between the two country groups, because (ii) the redistributive impact of the fiscal system is very large in Europe and very small in Latin America; and (iii) where fiscal redistribution is significant, it is achieved mostly through transfers rather than taxes. While the priorities of pro-equity fiscal reforms vary across Latin American countries, overall the prospects for major fiscal redistribution lie mainly in raising the volume of resources available for transfers, and improving their targeting, rather than increasing the progressivity of Latin America's tax systems.

4.1 Introduction

According to the World Bank's *World Development Report 2006*, Latin America ranks at the top among world regions in terms of inequality, second only to Sub Saharan Africa. Inequality in Latin America is pervasive - it extends to every aspect of life, from the distribution of income and assets, to access to education and health services, and political voice and influence.

High inequality is viewed by many as intrinsically bad on moral and ethical grounds[5]. But in ad-

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[5]See Bank (2006) for a review of the philosophical arguments for equity.

dition high inequality can be a powerful drag on development and prosperity, for several reasons. First, for given average income levels, higher income inequality means higher poverty. And the magnitude of this effect is far from negligible: according to Perry & Servén (2006), if Latin America had the levels of inequality found in developed countries, its current poverty rate would be halved. High poverty is not only a tragedy in itself but, as a rapidly expanding literature has argued, it can also be a source of underdevelopment traps in which financial market imperfections and institutional constraints prevent the poor from contributing to the growth process[6]. This in turn retards growth and makes poverty self-perpetuating. Indeed, the international evidence is consistent with this growth-deterring effect of poverty (e.g., López & Servén (2006b)).

Second, the other side of the coin is that high inequality also weakens the impact of aggregate income growth on poverty: the more unequal income distribution is, the faster the rate of growth required to achieve a given reduction in poverty[7]. Combined with the mechanism just described, the result is that inequality lies at the core of the vicious circles of stagnation and poverty in which many developing countries appear to be stuck.

Third, high inequality can also be a source of distributive conflict and social tension, which tend to undermine the legitimacy of policies and institutions as well as their stability, and in particular weaken property rights, thus discouraging investment and thereby growth[8].

These considerations have brought social equity to center stage in the Latin American policy debate. Indeed, recent major contributions to that debate portray social equity as one, or even the, key organizing principle for the region's development strategy, and some retrospective assessments of Latin America's performance under the so-called 'Washington Consensus' of the 1990s blame the reforms' neglect of equity considerations for the region's limited achievements over the 1990s and 2000s[9]. On both counts, equity has become a central concern for development policy across the region, echoing the views of Latin American citizens, who overwhelmingly perceive the distribution of income as 'unfair' or 'very unfair' and believe that the state should take responsibility for reducing the gap between rich and poor[10].

High income inequality usually reflects an unequal distribution of assets, such as land and human capital. Across countries, asset inequality and income inequality are closely associated - e.g., the cross-country correlation between the Gini coefficients of the distribution of income and the distribution of human capital, as proxied by years of schooling is above .75, while the correlation between the Gini indices of the distribution of income and the distribution of land is around .50 (de Ferranti *et al.*, 2004). And asset distribution is highly unequal in Latin America. For example, the Gini coefficient of the distribution of operational holdings of agricultural land is about .81 for Latin America (Deininger & Olinto, 2000), while in other world regions it hovers around .60; similarly, for the distribution of years of schooling, the Gini coefficient is .42 in Latin America, against .27 in industrial countries[11].

[6] See for example Azariadis & Stachurski (2005).

[7] See Bourguignon (2004) and López & Servén (2006a) for some quantitative estimates of the order of magnitude of this effect.

[8] Alesina & Rodrik (1994) and Alesina & Perotti (1996) present empirical evidence of the adverse effect of inequality on investment through these channels.

[9] See for example Birdsall & de la Torre (2001); Kuczynski & Williamson (2003); de Ferranti *et al.* (2004); Perry & Servén (2006).

[10] According to the regional opinion poll Latinobarometro, 79 percent of those surveyed in 2009 regarded the distribution of income in their countries as unfair or very unfair. In an earlier 1996 poll, over 90 percent of all respondents agreed or strongly agreed with the view that the state should take charge of reducing inequality.

[11] Asset inequality has deep roots in Latin America, which some authors (Engerman & Sokoloff, 2000; Acemoglu *et al.*, 2001) trace to the pattern of specialization during colonial times, heavily reliant on natural resources - especially mining and sugar production - and on the use of subordinate labor. In this view, the colonists developed institutions (related, for

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But high inequality can also reflect the failure of fiscal policy to perform its redistributive function - one of Musgrave (1959)'s three classic fiscal functions[12]- through appropriate use of taxes and transfers to correct socially-undesirable distributive outcomes arising from market forces, given the prevailing distribution of assets. The evidence shows that industrial-country governments are highly effective at this redistributive function, but developing-country governments are not, and in fact they are often part of the problem rather than the solution[13].

If inequality is above socially tolerable levels, as the survey evidence cited above suggests, in general policy action has to take a dual approach. On the one hand, it should target the deep roots of inequality, through interventions aiming to broaden asset ownership and equality of opportunity - e.g., by expanding access to education and health. But this is likely to be a long-term process with much of the payoff accruing in the distant future. On the other hand, policy should ensure, through the necessary fiscal reforms, that the fiscal system performs its redistributive function effectively.

In this paper we document and compare the performance of Latin American and Western European fiscal systems. Our use of internationally-comparable data on tax and transfer incidence, allows us to adopt a comparative perspective that extends the literature in two dimensions. First, and perhaps the most novel result of this paper, we show that the biggest contrast between the two regions in terms of income inequality concerns the distributional impact of the tax and transfer system. In contrast with industrial countries, in most Latin American countries the fiscal system is of little help in reducing inequality. This is the combined result of two adverse factors. One is that transfers do help redistribution, but in general not much, especially if only cash transfers are considered. The other is that the scope for active fiscal redistribution is severely constrained by the region's low levels of tax collection, which (with a couple of exceptions) are well below the international norm - a fact that also underlies the shortcomings of Latin America's public sectors in the other classic functions of efficiency and stability.

Second, wherever there is significant fiscal redistribution, it is achieved through transfers - especially in-kind transfers - rather than taxes. This result is consistent with the conclusions of Harberger (2003) and, more recently, OECD (2008), and extends to other Latin American countries the earlier findings of Engel *et al.* (1999) for Chile. In addition, however, the paper's quantitative results show that the redistributive impact of transfers can be quite considerable, in contrast with Harberger's conclusion that they were at best moderately more potent than taxes in this regard. In effect, we find that transfers largely responsible for the wide disparity in income inequality between the European and Latin American countries examined. From the policy perspective, the finding that the redistributive role of the tax system is in fact relatively minor also means that distributional concerns should not dictate the choice between income tax and VAT-based strategies to raise revenue collection, even for policy makers mindful of equity concerns.

On the whole, the evidence we present shows ample room for enhancing the distributive performance of Latin America's fiscal systems through appropriate fiscal reform - with specific reform priorities determined by country circumstances. Our analysis is subject to some caveats, however. First, we focus on annual income rather than lifetime income, thus neglecting intertemporal issues such as redistribution over the life-cycle). This may lead to overstating inequality as well as the regressivity of indirect taxation (Fullerton & Rogers, 1993). Second, we rely on standard incidence analysis, leaving aside incentive effects of fiscal interventions whose proper consideration would

example, to land use and political control) that helped perpetuate their political dominance and their wealth, resulting in a highly unequal distribution of assets that has persisted until today.

[12] The other two are efficiency and stabilization; see Musgrave (1959).

[13] For a discussion of the anti-poverty effect of taxes and transfers in industrial countries see Smeeding (2006). For a comparison of industrial and developing countries see Chu & Gupta (2000).

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require a fully-specified general equilibrium model. Similarly, our discussion of the incidence of transfers on income inequality focuses on who benefits from spending on *average*, rather than *at the margin*, and thus we ignore possible systematic differences between average and marginal incidence - which could lead us to underestimate the redistributive impact of transfer increases. Third, our analysis only considers the impact of income and consumption taxes and ignores the distributive impact of other taxes (e.g. on international trade) that are relatively important in some Latin American countries (especially in Central America). Fourth, our analysis is constrained by data limitations, particularly when broadening its focus to include also the distributive impact of indirect taxes and in-kind transfers. This requires combining data from different sources and, for some countries, different years. While we believe this is of little consequence for our qualitative conclusions, these data limitations may introduce some margin of error in the quantitative results. And fifth, our use of Western European countries as benchmark of comparison should also be taken as a reflection of data availability, rather than a normative statement about the desirability (or lack thereof) of the 'European model' as a specific choice between redistributive activism and efficiency costs. Related to this, we do not discuss the institutional constraints and implementation challenges that Latin American countries would have to overcome should they wish to expand significantly the role of taxes and/or transfers as redistributive vehicles[14].

The rest of the paper is structured as follows. In Section 4.2, we compare the distributional impact of taxes and transfers in Latin America and in Western Europe. Section 4.3 explores in detail why Latin America's fiscal policy is largely ineffective at reducing income inequality. Finally, Section 4.4 closes with some concluding thoughts.

4.2 Income inequality and fiscal redistribution in Latin America

The centerpiece of our analysis is an assessment of the redistributive performance of Latin America's fiscal systems. For this purpose, we document the facts concerning income inequality across Latin America before and after various fiscal interventions. We use a comparative perspective putting side by side the experience of two groups of countries from Latin America and Western Europe.

The country sample we use for this analysis is determined by the availability of suitable income distribution data. It comprises the six largest Latin American economies[15], which together account for 75% of the region's population and over 80% of its 2008 GDP, and up to 15 countries from Western Europe. However, not all of the latter countries possess complete information, and therefore some of the cross-regional comparisons presented below cover a smaller group of European economies. The Appendix describes in detail the data sources.

Before proceeding, it will be useful to define some terms. We use the term *market income* to refer to income before taxes and government transfers - thus its distribution is largely determined by market rewards to the private assets and efforts of individuals, and by the underlying distribution of those private assets. However, from a welfare perspective a more relevant measure is *disposable income* - i.e., household income after government cash benefits (e.g. pensions, unemployment insurance, social assistance transfers) have been received and direct taxes have been paid. Disposable income, rather than market income, is the relevant measure of the spending capacity of the different households. In addition, we use below two other definitions of household income. The first one is *gross income*, equal to market income plus government cash benefits or, equivalently, total household income before taxes. The second is what we shall call, for want of a better term, *post-tax*

[14] See for example Dethier (2007) for a discussion of the difficulties that developing countries would likely encounter if they tried to raise social security coverage to levels similar to those found in the developed world.

[15] Argentina, Brazil, Chile, Colombia, Mexico and Peru.

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income, which deducts from gross income all taxes, both direct and indirect, and thus offers a proxy for households' purchasing power. It allows a comprehensive view of the incidence of the overall tax burden, particularly in countries that rely heavily on indirect taxation.

Comparing the degree of inequity of the distribution of the various income definitions we can get a summary measure of the distributional effects of the three fiscal interventions that distinguish them - direct taxes, indirect taxes, and transfers. This is done in Table 4.1, which reports the value of the Gini coefficients for the four alternative measures of income. For the moment we focus on the first three columns of the table.

Income inequality is commonly measured in terms of the distribution of disposable income. The corresponding Gini coefficients appear in column 3 of the table, and they confirm that, by this definition of income, Latin America is much more unequal than Europe[16]. The Gini coefficients of disposable income average .50 for the Latin American countries considered, against just .31 for the European countries. In fact, the Latin American country with the lowest Gini coefficient in this sample (Chile) has inequality levels above those of the most unequal European country (Portugal)[17].

The distribution of disposable income is the combined result of the distribution of market income and the redistributive action of the state through direct taxes and cash transfers. Columns (1) and (2) in 4.1 help identify the role of each of these ingredients. The Gini coefficients of market income, shown in column 1, reveal a much less marked contrast across regions than that emerging from column (3). In fact, whereas the average Gini coefficient of market income for the Latin American countries, at .52, is only 2 percentage points above that of disposable income, the Gini coefficients of market income for the European countries are substantially higher than those for disposable income: the average of the former for the 15 countries in the sample is now .46, a whopping 15-percentage points higher than the average of the latter.

Thus, one important message from this analysis is that the equity gap between the two groups of countries is much narrower in terms of market income than in terms of disposable income. In other words, most of the difference between the levels of disposable income inequality in the two regions is due to the different impact of taxes and transfers: they reduce market income inequality considerably in Europe, and very little in Latin America. Indeed, before direct taxes and transfers are considered, several countries in Europe possess Gini indices comparable to those typically found in Latin America. This is the case, for example, of Ireland and the UK, whose Gini coefficients of market income are estimated at .53 and .52 respectively. Even the Nordic countries, commonly praised for their levels of equity, show relatively high inequality of market incomes. The Gini indices of Denmark, Finland, and Sweden are .49, .49 and .45 respectively. The most egalitarian countries in the sample in terms of market incomes are Austria and the Netherlands, with Gini coefficients of .38 and .39 respectively. These values are in sharp contrast with the Gini coefficients of disposable income of the same set of European countries, which are much lower. The difference is especially marked in the cases of Denmark or Ireland, where taxes and transfers lower the Gini coefficient by 20 and 19 percentage points respectively. Even the European countries that redistribute the least through the tax-benefit system (Italy, Greece and Portugal) still manage to lower their Gini indices by more than 10 percentage points.

The natural question is whether the dramatically different redistributive impact of the tax-benefit system in Latin America and Europe reflects mostly the action of transfers, taxes, or both. The redis-

[16]The Gini coefficient estimates in the table are close to those reported by de Ferranti *et al.* (2004) and Bank (2006). For disposable income, they report an average Gini coefficient of about .5 for Latin America and an average Gini coefficient of .31 for European countries.

[17]This result continues to hold in a broader sample of Latin American countries, because the lowest Gini in the region (Uruguay's) is .42.

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Table 4.1: Gini coefficients of the distribution of different income definitions

Country	Market Income	Gross Income	Disposable Income	Post Tax Income
	(1)	(2)	(3)	(4)
LAC	*0.52*	*0.51*	*0.50*	*0.51*
Argentina	0.49	0.49	0.48	0.48
Brazil	0.58	0.55	0.55	0.55
Chile	0.48	0.47	0.46	0.47
Colombia	0.54	0.54	0.53	0.54
Mexico	0.51	0.50	0.49	0.49
Peru	0.50	0.50	0.49	0.51
*EU15 **	*0.46*	*0.36*	*0.31*	*0.34*
Austria	0.38	0.30	0.25	N.A.
Belgium	0.47	0.36	0.29	0.30
Denmark	0.49	0.35	0.29	N.A.
Finland	0.49	0.36	0.32	0.35
France	0.42	0.35	0.31	0.34
Germany	0.43	0.33	0.28	N.A.
Greece	0.47	0.40	0.36	0.38
Ireland	0.53	0.39	0.34	0.35
Italy	0.48	0.41	0.37	0.40
Luxembourg	0.41	0.30	0.24	0.25
Netherlands	0.39	0.30	0.26	0.27
Portugal	0.49	0.43	0.38	0.41
Spain	0.47	0.40	0.35	0.37
Sweden	0.45	0.33	0.29	0.33
UK	0.52	0.39	0.34	0.36

Source: Author's calculation (see Data Appendix).

Notes: Income definitions considered are the following: (1) Market Income: Income before taxes and government transfers; (2) Gross Income: Market Income plus government cash transfers; (3) Disposable Income: Gross Income minus direct taxes; (4) Post Tax Income: Disposable Income minus indirect taxes.
$*$The average Gini coefficients computed for the European sample excluding Austria, Denmark and Germany (as in Column 4) for the Market, Gross and Disposable Income are 0.47, 0.37 and 0.32 respectively. Likewise, the average Gini coefficients computed for the European sample including only those countries considered in Table 4.4 for the Market, Gross, Disposable and Post-tax Income are 0.47, 0.37, 0.32 and 0.34 respectively.

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tributive effect of public transfers can be inferred comparing the Gini coefficients of gross income in column (2) of the table with those of market income. This again reveals a sharp contrast between the two regions. Public transfers contribute only slightly to lower inequality in Latin America - the Gini coefficients of gross income are at best 1-2 percentage points lower than those of market income, and in some countries (Argentina, Colombia, Peru) there is virtually no difference between the former and the latter. However, transfers play a big role in the European countries, where the difference between the Gini coefficients of the two income definitions averages 10 percentage points. This conceals substantial variation across countries: in Denmark, Finland, Ireland, the UK, and Sweden public transfers bring the Gini coefficient of gross income 12-14 percentage points below that of market income. In Belgium, Germany, and Luxembourg, the effect is sizable too - some 10-11 percentage points. At the other extreme, in Portugal transfers lower the Gini coefficient by just 6 percentage points.

As for the effect of direct taxes, the contrast between regions is less dramatic. The redistributive impact of direct taxes can be inferred by comparing the Gini coefficients of gross and disposable income shown in columns (2) and (3) of Table 4.1. Like in the case of transfers, the conclusion is that direct taxes reduce the levels of income inequality much more in European countries than in Latin America. For example, direct taxation lowers the Gini coefficient of household income by 6-7 percentage points in Austria, Belgium, Denmark and Luxembourg, and by an average 5 percentage points for the fifteen European countries considered. In turn, in the Latin American countries direct taxes are also progressive - with the exception of Brazil - but their impact on inequality is very weak: on average, the Gini coefficients decline by just 1 percentage point, with very little variation across countries.

Thus, the second important message that emerges from the previous discussion concerns the relative redistributive roles of direct taxes and transfer benefits. Among the Latin American countries considered, both are of roughly similar (and modest) magnitude, but in the European countries transfers play a much bigger role than taxes: of the 15 percentage-point difference between the average Gini coefficients of market and disposable income across European countries, about two-thirds (10 percentage points) are due to transfers[18].

As we shall discuss below, these two basic messages remain unchanged if we take a less-conventional, broader view of fiscal redistribution that encompasses also indirect taxation as well as in-kind transfers.

4.3 What limits fiscal redistribution in Latin America?

To summarize, the empirical evidence suggests that the redistributive impact of taxes and transfers is much bigger in Europe than in Latin America. But why exactly?

Engel *et al.* (1999) provide an analytical framework that helps answer this question. In their framework, the redistributive impact of a country's fiscal system is shaped by three factors. First, the volume of tax collection – as tax collection capacity ultimately determines the feasible volume of transfers. Second, the incidence of taxation. Third, the incidence of transfers. We next assess how Latin America fares in each of these three areas.

[18] This dominant role of transfers over taxes is stressed by Harberger (2003) and corroborated by the findings reported in OECD (2008).

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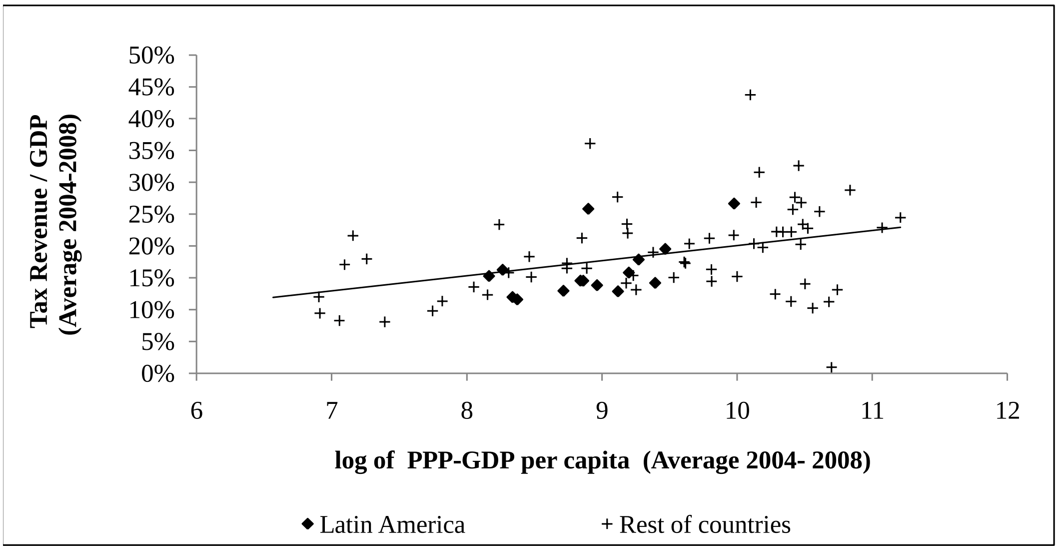
Figure 4.1: Central government tax revenue vs. per capita GDP



Source: Author's calculations

4.3.1 Tax collection

The overall volume of tax collection sets an upper bound on redistributive and other expenditures. This raises the question of whether Latin America's tax revenues are too low to permit adequate redistribution on the expenditure side. Of course, there is no such a thing as an optimal level of taxation that applies to all countries in all circumstances, and in setting the level of taxes countries have to trade off the excess burden of taxation against the social value of the public goods and services (such as education or infrastructure) and/or the distributional changes to be funded with them. These costs and benefits are likely to be affected by a host of country-specific factors.

Tax collection has been on the rise in Latin America in recent years. Since the early 1990s it has grown by an average of 2-3 percent of GDP (Cetrángolo & Gómez-Sabaini, 2007). But in spite of the rising trend, collection volumes remain well short of the international norm. Figure 4.1 shows that almost all Latin American countries lie below the regression line relating tax collection to the level of development, as measured by per capita GDP. Indeed, the median country in the region collects 3 percentage points of GDP less than would be expected given its level of development[19].

Why is tax collection so low in Latin America? Conceptually, there are two possible reasons: low statutory tax rates, and narrow tax bases. Of course, each of the two may play differently for different taxes. Table 4.2 offers a comparative perspective on tax rates across world regions, for both income and value added taxes. Latin America's income tax rates are at the low end of the spectrum, for both personal and corporate income. Indeed, income tax rates have declined across the region over the last two decades. The average top personal income tax rate fell from 49.5 percent in 1985 to under 30 percent in 2004. Likewise, the average top corporate tax rate declined from 43.9 percent in 1986 to 26.6 percent in 2004. As for VAT rates, they have followed the opposite trend: the regional average of countries' general rates rose from 12 percent in 1992 to 15 percent in 2004

[19] A similar result is obtained by Perry & Servén (2006) when they relate average collection during the 1990s (rather than in the late 2000s) to per capita GDP.

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Table 4.2: Comparative perspective on Tax Rates

Region	Tax Rate		
	Individuals	**Corporate**	**VAT**
East Asia Pacific	33.50	31.50	10.00
Latin America and the Caribbean	29.00	26.60	15.00
Middle East & North Africa	48.00	40.00	17.00
OECD	45.00	35.00	17.25
South Asia	39.50	41.00	15.00
Sub Saharan Africa	38.00	36.00	17.50

Source: Author's calculations based on Eichhorn (2006), World Bank's World Development Indicators 2006, OECD tax Database, Doing Business Database. Figures are group averages (of top rates in case of personal and corporate taxes) and refer to latest available year.

(Gómez-Sabaini, 2006). Still, Table 4.2 shows that they are not high by international standards.

What about tax bases? Narrow tax bases can be the result of plain tax evasion, or too generous tax concessions and loopholes. One rough way of gauging tax bases is by looking at tax productivity, defined as actual revenues (as a percentage of GDP) relative to the prevailing nominal rates[20]. This exercise is reported in Table 4.3. VAT productivity is comparable in Latin America to that observed in other world regions, including rich countries, although the regional average conceals significant cross-country variation[21]. For income taxes, however, Latin America places far behind industrial countries, even below Sub-Saharan Africa. This suggests that the problem behind Latin America's low income tax receipts is primarily one of narrow tax bases rather than tax rates.

Tax evasion is rampant across the region, although precise figures are obviously hard to calculate. Some recent estimates place evasion rates in the 45-60 percent range for both the personal and the corporate income tax (Jiménez *et al.*, 2010). Even for the VAT the estimates are quite large. Some rough calculations suggest that a 30 percent reduction in evasion would increase tax revenues by 17 percent in Argentina, 14 percent in Brazil and 12 percent in Chile (Pessino & Fenochietto, 2004).

The practice of tax evasion in Latin America is encouraged by weak tax administrations across the region. Indeed, opinion polls show that large majorities of individuals perceive tax collection as largely arbitrary and unfair – only 23 percent of those surveyed by Latinobarómetro in 2003 thought tax collection was 'impartial'. The culture of evasion is likely further promoted by the overwhelmingly negative views about the state's capacity to spend wisely - in the same survey, just 15 percent of respondents believed that tax revenues would be put to good use.

A key factor behind poor tax compliance (and hence tax collection) is informality, which is pervasive in Latin America. The size of the 'shadow economy' relative to the formal one is estimated to average around 40 percent, the highest figure across world regions, equaled only by Sub-Saharan Africa[22]. Maloney & Saavedra (2007) also argue that tax collection has a strong negative association

[20] In the case of the income tax, our productivity measure is based on the top marginal rate.

[21] Across the region, VAT productivity ranges from less than 25 percent in Mexico to over 50 percent in Honduras (Martner & Aldunate, 2006). The calculations in the table use GDP to approximate the VAT base. Using aggregate consumption instead yields similar qualitative results.

[22] See Eichhorn (2006) and Loayza & Sugawara (2009).

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Table 4.3: Tax Collection and Tax productivity

Region	Collection (% of tax revenue)		Tax Productivity	
	Income	VAT	Income	VAT
East Asia Pacific	45.24	30.51	0.19	0.42
Latin America and the Caribbean	28.15	60.87	0.05	0.23
Middle East & North Africa	49.64	24.38	0.07	0.09
ÒECD	38.36	41.22	0.16	0.39
South Asia	34.60	39.52	0.08	0.24
Sub Saharan Africa	24.71	43.97	0.09	0.34

Source: Author's calculations based on Eichhorn (2006), World Bank's World Development Indicators 2006, OECD tax Database, Doing Business Database.

with informality, with higher informality coming along with lower tax revenue, even when the level of development is held constant.

Informality, of course, is not exogenous, but the combined result of poor public services (that reduce the benefits of formality), weak tax administration, and the tax structure itself, in addition to other factors. Payroll taxes in particular are widely viewed as providing a major incentive to informality, and still account for a considerable fraction of total direct tax collection in Latin America - as high as 50 percent in Brazil (Martner & Aldunate, 2006). More broadly, the overall tax burden facing formal firms - combining corporate taxes and VAT along with payroll contributions– is higher in Latin America than in other world regions, with the exception of Sub-Saharan Africa[23], and this gives small firms a strong incentive to remain informal[24].

Aside from evasion, tax concessions - exemptions, deductions, and other loopholes - play a major role in narrowing income tax bases across Latin America. In many countries, households with above-average income levels are exempt from the personal income tax because of very high minimum personal exemption levels. The average level of the minimum taxable income is twice the region's per capita income, and in some countries it is much higher (Figure 4.2)[25]. High personal exemptions combined with a plethora of deductions severely reduce effective income tax rates, especially for the rich. As a result, tax rates highly progressive on paper collect little revenue in practice because the higher rates only kick in at extraordinarily high levels of income, so they are rarely, if ever, effective (Tanzi, 2000).

The limited collection of income taxes has prompted some experimentation with alternative taxes - such as gross asset taxes and taxes on financial transactions. The former is essentially a minimum tax, typically applied at a rate of 1 percent. While easy to administer and potentially effective against evasion, it is not devoid of efficiency problems (Gómez-Sabaini, 2006). In turn, taxes on financial transactions were originally introduced in several countries as emergency collection devices, but they have proven hard to replace and have tended to persist given their low administration cost and high revenue collection capacity - in spite of their distorting potential in terms of promoting financial

[23] See Bank (2006).
[24] On the role of taxes in encouraging informality in Latin America, see Levy (2008) and de Paula & Scheinkman (2009).
[25] Shome (1995) reports that these minimum exemptions are higher in Latin America than in other developing regions.

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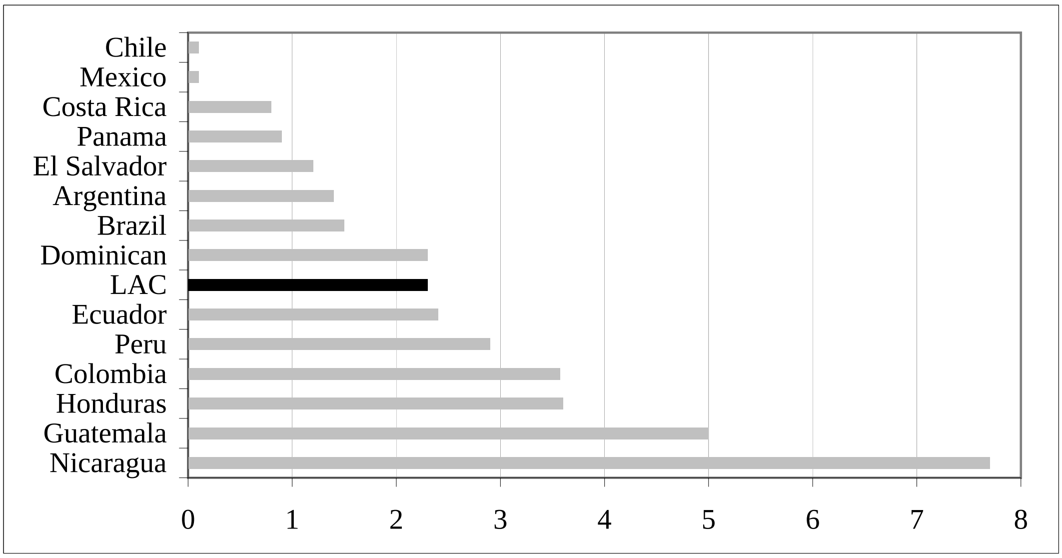
Figure 4.2: Personal exemption levels (multiples of GDP per capita)



Source: Stotsky & Woldemariam (2002).

disintermediation[26].

4.3.2 Tax incidence

The distributional implications of Latin America's tax structure have been widely debated. On *a priori* grounds, the large weight of indirect taxes in the region's total revenues leads to the presumption that overall taxation is likely to be regressive or neutral at best, given the conventional wisdom that income taxes are usually progressive while consumption taxes such as the VAT are not. In theory, multiple rates for VAT and similar taxes could achieve some progressivity - at the cost of potentially significant administrative complication - by granting more favorable treatment to food and other goods primarily consumed by the poor. Indeed, most Latin American countries make use of multiple VAT rates. In practice, however, their redistributive effect is generally modest, because much of the benefit from reduced VAT rates accrues to the rich, due to their higher levels of expenditure[27].

On the other hand, in a number of Latin American countries, the effective personal income tax rates of the upper income brackets are only a fraction of the statutory rates, negating much of the supposed progressivity of income taxes. In turn, payroll taxes are also usually regressive, since in most countries they are capped, so that the average rate declines with the level of income.

The few available assessments of the incidence of taxes in Latin America tend to find a neutral or regressive effect. Gómez-Sabaini (2006) (2006) summarizes the findings of individual studies of ten Latin American countries. In eight of them the overall effect of the tax system is found to

[26] See Tanzi (2000).

[27] See Tanzi (2000) and Ebrill *et al.* (2001) for discussion. Conceptually, what matters for redistribution is the change in VAT collected relative to the respective incomes of the rich and poor, compared with the change, similarly scaled, in the public transfers accruing to each of the two classes that were being financed out of the lost VAT collection. On the other hand, the regressive effects of consumption taxes may be much reduced in an intertemporal framework, as already mentioned.

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be regressive; in the other two it is approximately neutral. Those studies, however, use disparate methodologies, so that their results are not strictly comparable across countries.

The data assembled for this paper allow a more systematic look at the incidence of taxes across the six major Latin American economies. A summary view is provided in Figure 4.3). For each country, the left panel of the figure reports the share of total taxes paid by each quintile of the income distribution. The figure shows that the upper quintiles tend to pay more taxes in all cases. For example, the top quintile pays at least 50 percent of total taxes, and in some cases (Brazil, Mexico and Colombia) as much as 70 percent. However, this is hardly surprising given that the top quintile of the population also accounts for 55 to 65 percent of total income.

Indeed, a more informative measure of the tax burden is the effective tax rate that the different quintiles pay on their income (i.e., taxes paid relative to gross income). This is shown in the right panel of each figure. Relative to their income, the tax burden faced by poor households is not very different from that faced by households at the top of the distribution, and in Peru and Brazil the former face heavier tax burdens than the latter. However, different taxes have different incidence. Income taxes are generally progressive and, in some cases (e.g., Brazil, Chile or Peru), highly so as contributions come almost solely from the top quintile. Value added taxes, on the other hand, tend to be regressive.

How does this configuration of taxes affect income inequality in Latin America? We already reviewed the effect of direct taxation on the distribution of disposable income. But, as shown in Table 4.3, the bulk of Latin America's tax revenues arise from indirect taxes. Hence, a better way to assess the distributional effect of overall taxation is by comparing columns (2) and (4) of Table 4.1, which report the Gini coefficients of gross income and post tax income, respectively (with the latter calculated, as noted, by subtracting indirect taxes from conventional disposable income)[28]. The comparison shows that in Latin America the combined effect of direct and indirect taxes on income inequality, as captured by the Gini index, is minimal. Taxes contribute to a decline of the Gini index of 1 percentage point in Argentina and Mexico; an increase of 1 percentage point in Peru; and no noticeable change in Brazil, Chile and Colombia.

Closer inspection of columns (2), (3) and (4) of Table 4.1 further reveals that both direct and indirect taxes have very modest distributional effects in the Latin American countries considered - although their respective impacts generally are of opposite signs. Comparison of columns (2) and (3) indicates that direct taxes tend to be inequality-reducing, except in Brazil (where the Gini coefficient remains unchanged). On the other hand, comparison of columns (3) and (4) (i.e. the Gini coefficients of disposable and post tax income) indicates that indirect taxes cause a slight increase in inequality. However, their estimated impact is very modest in all countries. The largest effect occurs in Peru, where indirect taxation raises the Gini coefficient by two percentage points.

To provide a comparative perspective, Table 4.1 also reports a similar exercise for a subsample of European countries, based on data for the incidence of VAT receipts drawn from Baldini & Mantovani (2004). As already discussed, comparison of the average Gini coefficients in columns (2) and (3) indicates that direct taxation has a bigger redistributive impact in Europe than in Latin America. This pattern persists if we look at the redistributive impact of overall taxation, in spite of the regressive effect of indirect taxation. Comparison of columns (2) and (4) of Table 4.1 show that, in the European countries for which the requisite data is available the Gini coefficient of post-tax income is, on average, 3 percentage points lower than the Gini coefficient of gross income[29]. Put differently,

[28] In Table 4.1, data on post-tax income is not available for Austria, Denmark and Germany. However, it is easy to verify that this is of no consequence for the qualitative and quantitative conclusions from the comparisons across country groups described in the text.

[29] Similar results have been found for the United States by Pechman (1985, 1990), who analyzed the structure of the tax burden in the US, concluding that the tax system has very little effect on the distribution of income, and for the UK by Jones

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in the majority of European countries shown the regressive effect of indirect taxes offsets almost half of the progressive impact of direct taxes. There is some variation around this norm, however. For example, the estimated effect of overall taxation is strongly progressive in Belgium and Luxembourg.

In summary, two main messages emerge from our analysis of tax incidence. First, the progressive redistributive effect of direct taxation is weaker in Latin America than in Europe. But second, the impact of overall taxation on inequality is, on the whole, fairly small. This applies to both the Latin American countries under consideration and, with few exceptions, the European countries for which data is available. The simple reason is that, in virtually all countries examined, the regressive effect of indirect taxation offsets much of the progressive impact of direct taxes.

4.3.3 The redistributive effects of public spending

Section 4.2 showed that in European countries public transfers reduce sharply the degree of inequality of market incomes, while in Latin America public transfers achieve little on the inequality front. There are two potential (not mutually exclusive) reasons for this difference. The first is that the volume of transfers is much smaller in Latin America than in Europe. The second is that the targeting of the given volume of transfers is less progressive in the former than in the latter region.

Regarding the volume of transfers, Latin America spends, on average, less than half of what is spent in Europe. For example, Lindert & Shapiro (2006) find that on average the six Latin American countries considered in our study devote to transfers about 7.3 percent of GDP[30]. Of this total, about 6.3 percent of GDP goes to social insurance programs (such as pensions and unemployment insurance) and about 1 percent of GDP to social assistance programs such as conditional cash transfers, school meals, and scholarships. In contrast, according to Eurostat data EU-15 countries spend on average 14.7 percent of GDP on cash transfers, most of which for social insurance programs. Thus in principle, even if transfers were targeted just as progressively in Latin America as in Europe, we should expect them to have much less of an impact in the former than in the latter region, given the big difference in their total volume. Importantly, that difference reflects primarily the contrast between the levels of tax collection in the two country groups, rather than their differential use of tax revenues. Indeed, the ratio of total cash transfers to the tax collection of the general government is roughly similar in both sets of countries - it averages 35.9 percent in EU-15 countries, and 36.6 percent in the six Latin American countries considered[31].

But the problem is not only that Latin America spends relatively little on transfers. While in Europe public transfers are distributed in nearly-egalitarian way across income quintiles, in Latin America their targeting leaves much to be desired. This is shown in Figure 4.4, which reports for the Latin American countries the share of transfers (both cash and in-kind) flowing to each quintile of the income distribution, as well as the quintile's effective transfer receipt rate (i.e., transfers received relative to total gross income). The figure shows that the upper quintiles tend to receive more transfers in all countries shown. This reflects the dominant role of social insurance transfers (specifically, public pensions and unemployment insurance), which possess nearly universal coverage in Europe,

(2008), who using 2006-07 data estimates that the Gini coefficient of post tax income (.38) is 1 percentage point higher than that of gross income (.37).

[30] This average, however, conceals substantial heterogeneity among the group, which includes high spenders like Argentina and Brazil, moderate spenders (Chile, Colombia) and low spenders (Mexico and Peru).

[31] These figures are based on total tax collection and transfer data from ECLAC and Eurostat. There is of course considerable variation across countries in both groups. Among the Latin American countries, cash transfers range from a low of 25 percent of tax collection in Peru to a high of 41 percent in Brazil. In Europe, the range goes from 27 percent in Ireland to 45 percent in Germany.

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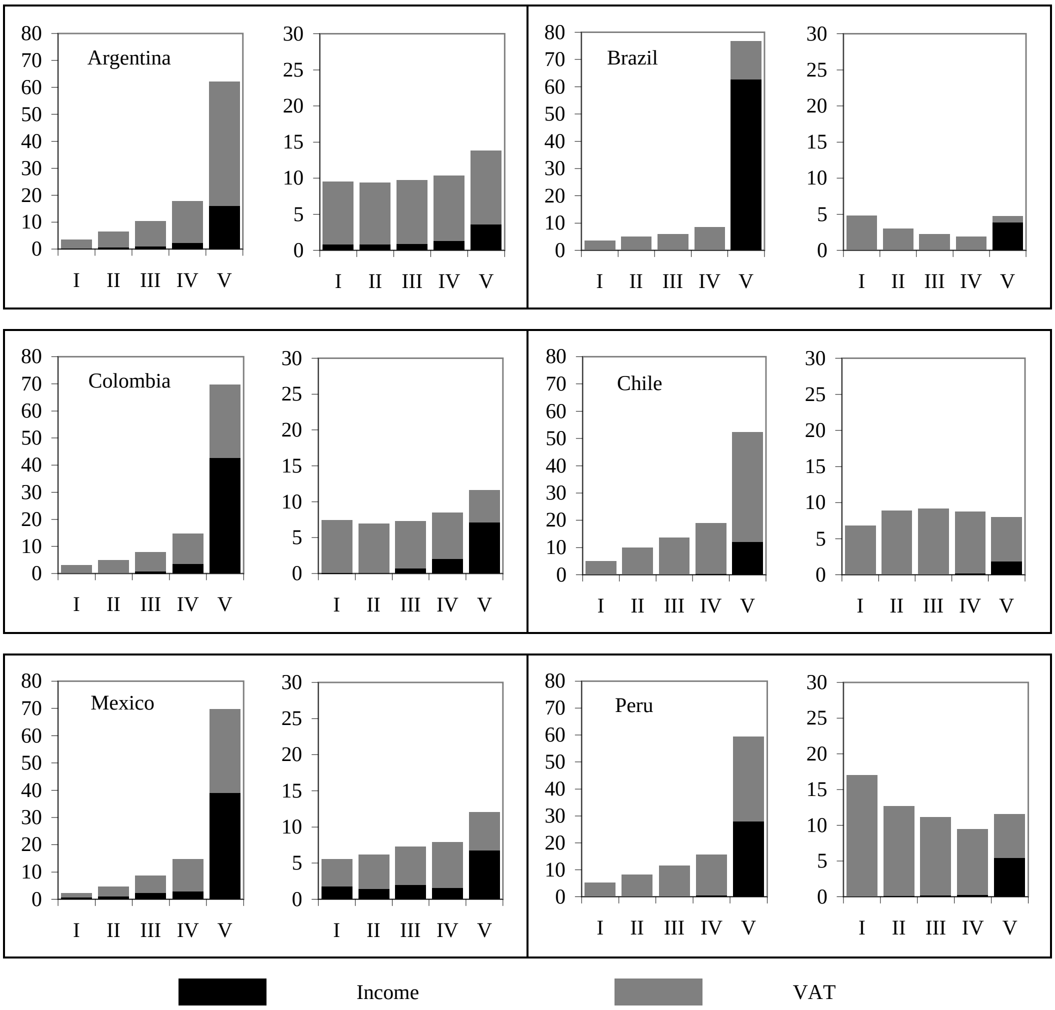
Figure 4.3: Taxation by income quintiles



Source: Author's calculation (see Appendix)
Note: Panels on the left are expressed as percentage of total contribution of all quintiles; panels on the right are expressed as percentage of the total gross income of each quintile

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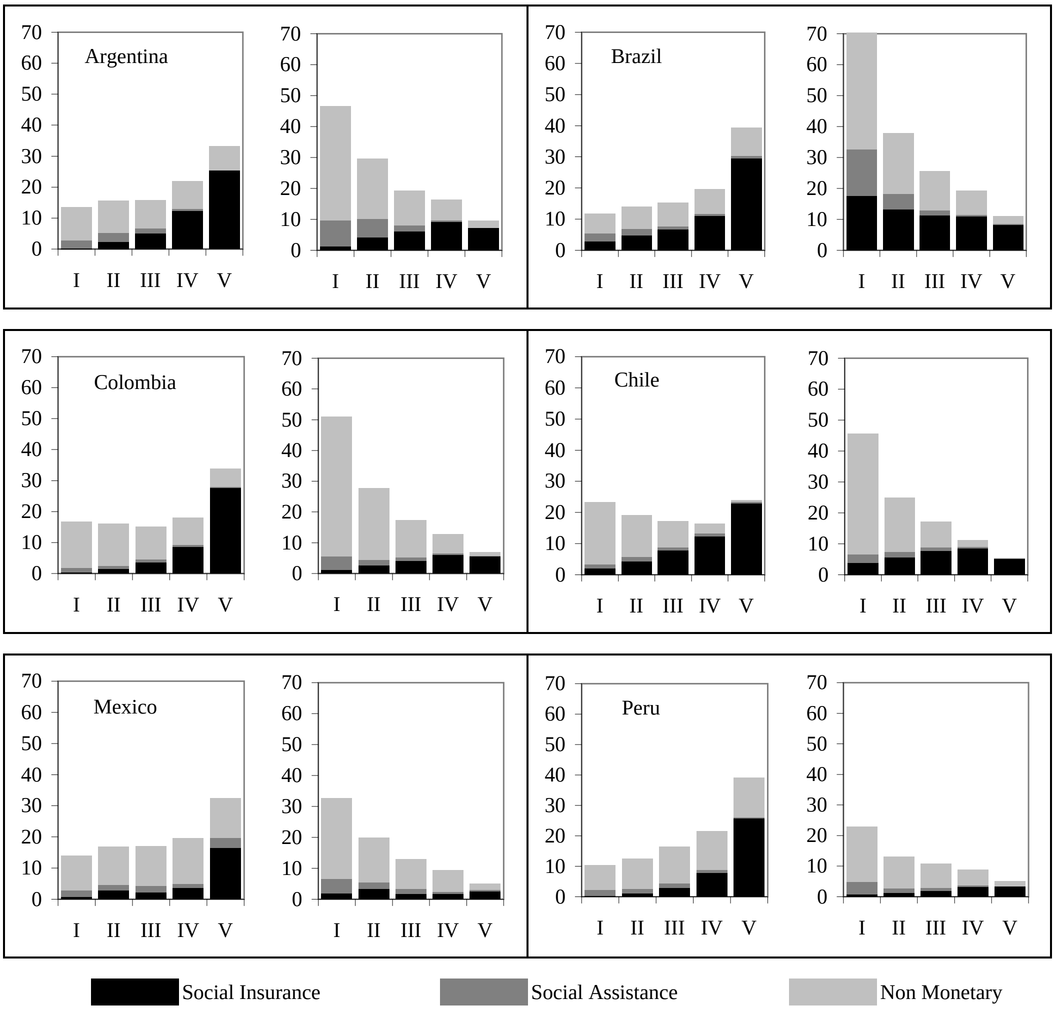
Figure 4.4: Transfers by income quintiles



Source: Author's calculation (see Appendix)
Note: Panels on the left are expressed as percentage of total receipts of all quintiles; panels on the right are expressed as percentage of the total gross income of each quintile

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but In Latin America typically accrue to the more affluent and thus tend to be regressive[32]. For example, in the case of pensions the data in de Ferranti *et al.* (2004) indicate that on average the top two quintiles of the population attract more than 80 percent of the spending, while the bottom quintile receives less than 3 percent[33]. Unemployment insurance is likewise regressive, although less markedly so. About 65 percent of the spending goes to the top two quintiles, with the bottom quintile obtaining less than 10 percent of the spending. Given that in Latin America poor households often work in the informal sector, it is not surprising that they have little access to these benefits aimed at formal sector workers[34].

This does not mean that all social programs are regressive. Indeed, there are programs, typically in the social assistance area, that are well targeted. Lindert & Shapiro (2006) show that Latin America's conditional cash transfer programs[35] are strongly progressive, with close to 75 percent of the resources accruing to the bottom two quintiles of the population. But in spite of their expansion in recent years, these programs are still quantitatively small and therefore their redistributive capacity is limited, at least in comparison with the large social insurance programs just reviewed. Thus, while social assistance transfers are systematically targeted towards the needy, their redistributive effect is largely outweighed by that of social insurance transfers.

So far, we have considered only cash transfers. It might be argued that in-kind transfers (such as free public education) could dramatically change the conclusions of the analysis, particularly if these transfers target the lower quintiles of the distribution. Indeed Figure 4.4 shows that in the Latin American countries considered the distribution of in-kind transfers among the different gross income quintiles is more progressive than that of cash transfers[36]. In fact, once in-kind transfers are taken into account, the figure clearly shows that, as percentage of their gross income, the total volume of transfers accruing to poorer households is bigger than that accruing to those at the top of the income distribution.

Quantifying the redistributive effect of transfers in kind requires taking a stand on how accurately their cost in terms of government spending reflects their value for the private individuals receiving them. There is clear evidence across Latin America, as well as in some European countries, that individuals often eschew public health and education services, available at low or zero cost, in favor of costlier private services perceived to be of higher quality. Even if we ignore this fact, and take the extreme view that the volume of public spending on health and education is an accurate measure of their value to individuals, does inclusion of these in-kind transfers lead to a more upbeat assessment of the redistributive performance of Latin America's fiscal systems? To answer this question, Table 4.4 reports the Gini coefficients of our various income definitions, recalculated after redefining pub-

[32]For the case of Brazil, this is stressed by Baer & Jr (2008).

[33]The biggest problem with Latin America's public pension systems from the point of view of equity probably is their low coverage: they are far from universal, and usually exclude workers in the informal sector and in agriculture. In addition, some special pension schemes for public sector employees are hugely regressive, as they offer extremely generous benefits to a privileged few, with the bulk of the cost financed by taxpayers. For example, the deficit of Brazil's civil service pension system, which covers only 13 percent of pensioners, amounted to nearly 4 percent of GDP in 2004. See Birdsall & Menezes (2008).

[34]It could be argued that public social insurance programs are not financed by direct taxation but by 'contributions' and therefore should be treated differently. For example, if pensions are viewed as an intertemporal transfer for an individual rather than as an intergeneration transfer at a point in time, the benefits of each household should be treated as deferred consumption. In practice, however, the link between contributions and benefits is tenuous at best, and the programs run sizable deficits financed by general taxation. Indeed, according to Lindert & Shapiro (2006) public pension subsidies financed through general taxation absorb about 5 percent of GDP in some Latin American countries, a figure higher than total spending on social assistance and, in some cases, higher than public spending on education and health.

[35]These include programs such as *Bolsa Escola* in Brazil, *Subsidio Unico Familiar* and *Solidario* programs in Chile, *Familias en Acción* in Colombia, and *Oportunidades* in Mexico.

[36]The incidence of in-kind transfers for the Latin American countries under consideration is computed using data from ECLAC (2007). The data covers only public expenditure on Health and Education.

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lic transfers to include both cash and in-kind transfers. On average, the latter represent 7.6 percent of GDP in the six Latin American countries under consideration. For comparative purposes, the same information is reported for the seven European countries for which the necessary data could be gathered; their in-kind transfers average 11.3 percent of GDP[37].

For ease of reference, the first column of Table 4.4 reproduces the Gini coefficients of market income from Table 4.1. The second column reports the Gini coefficients of gross income, inclusive of both cash and in-kind transfers. It is instructive to compare them with those in column (2) of Table 4.1, which considers only cash transfers. The former are substantially lower than the latter for all Latin American countries, by 4 percentage points on average. Comparing columns (1) and (2) of Table 4.4, we conclude that the broader concept of transfers contributes on average to a reduction of 5 percentage points in the Gini coefficient of Latin American market income, compared with the reduction of just 1 percentage point obtained in Table 4.1 when considering cash transfers alone. The main exception is Peru, where consideration of in-kind transfers makes little difference - they fail to enhance significantly the poor redistributive performance of cash transfers.

Thus, in-kind transfers do enhance fiscal redistribution in Latin America. However, their inequality-reducing effect is even bigger in the European countries shown in Table 4.4. Indeed, once in-kind transfers are taken into account, the Gini coefficient of gross income among the European countries shown averages just 0.31 - that is, 16 percentage points lower than the average Gini coefficient for market income (instead of the 10-point difference in Table 4.1 when only cash transfers were considered). Thus, in-kind transfers actually widen the gap between the redistributive performances of fiscal systems in Latin America and Europe.

To complete the picture, columns (3) and (4) of Table 4.4 show the distributional effects of direct and indirect taxes once in-kind transfers have been considered. Like in Section 4.3.2, on average both direct and indirect taxes exert little distributional impact in our sample of Latin American countries, while in Europe the progressive effect of direct taxation is reflected in a reduction of 5 percentage points of the Gini coefficient of disposable income relative to that of gross income (the same result found in Table 4.1). In turn, the difference between the Gini coefficients of gross and post-tax income averages 4 percentage points in the European countries considered, while it is nil on average among Latin American countries[38].

4.4 Concluding remarks

Latin America's high inequality extends to virtually all aspects of social and economic life, and is viewed by a large majority of the region's citizens as deeply unjust. High inequality undermines the stability and legitimacy of institutions and policies, and represents a powerful drag on Latin America's development prospects. These reasons put social equity at the top of the region's development agenda.

A close inspection of the international evidence on income inequality reveals that the big difference between Latin America and the more egalitarian countries of Western Europe lies not so much in the

[37]The data on total in-kind transfers of European countries is taken from Eurostat. Their incidence across income quintiles Figures for Europe are is based on data drawn from EUROMOD (2007); see the Appendix for further details. Unlike in Latin America, in the European countries considered the incidence across income quintiles of in-kind transfers is not much more progressive than that of cash transfers.

[38]The seemingly milder regressive effect of indirect taxation among European countries shown in Table 4.4, relative to that in Table 4.1, is mainly due to the different country sample. More precisely, the average Gini coefficients of Table 4.1, when calculated over the 7 European countries in Table 4.4, would equal 0.47, 0.37, 0.32 and 0.34 for market, gross, disposable, and post-tax-income, respectively.

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Table 4.4: Gini coefficient of the distribution of different income definitions (considering both cash and in-kind transfers)

Country	Market Income	Gross Income	Disposable Income	Post Tax Income
	(1)	(2)	(3)	(4)
LAC	*0.52*	*0.47*	*0.47*	*0.47*
Argentina	0.49	0.45	0.44	0.44
Brazil	0.58	0.52	0.51	0.52
Chile	0.48	0.42	0.41	0.42
Colombia	0.54	0.49	0.48	0.48
Mexico	0.51	0.48	0.47	0.47
Peru	0.50	0.49	0.48	0.49
EU7	*0.47*	*0.31*	*0.26*	*0.27*
Belgium	0.47	0.30	0.22	0.22
Germany	0.43	0.27	0.21	N.A.
Greece	0.47	0.35	0.31	0.32
Ireland	0.53	0.32	0.26	0.27
Italy	0.48	0.35	0.31	0.32
Netherlands	0.39	0.25	0.20	0.21
UK	0.52	0.32	0.27	0.28

Source: Author's calculation (see Data Appendix)
Notes: Income definitions considered in Columns (1) to (4) are the following: (1) Market Income: Income before taxes and government transfers; (2) Gross Income: Market Income plus government cash and in-kind transfers; (3) Disposable Income: Gross Income minus direct taxes; (4) Post Tax Income: Disposable Income minus indirect taxes.

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extent of the inequality resulting from market forces, but in the redistributive power of the state. To put it differently, the gap between the two regions in terms of income inequality is much bigger *after* taxes and public transfers than *before* taxes and transfers, and this implies that a good deal of Latin America's excess inequality over international levels reflects the failure of the region's fiscal systems to perform their redistributive functions. And the magnitude of this failure is considerable: while in European countries fiscal redistribution through direct taxes and cash transfers leads to an average reduction of some 15 percentage points in the Gini coefficient of the distribution of income, in Latin America the reduction is on average just 2 percentage points. If indirect taxes- which in general are mildly regressive - are taken into account as well, the redistributive performance of the fiscal system declines slightly in both regions, but the big gap between them in terms of redistributive capacity remains.

The evidence from rich countries also shows that the bulk of the state's redistributive impact is due to the effect of public transfers. In Europe they account for over two-thirds of the overall redistributive effect. While direct taxes are modestly progressive, much of their effect is counteracted by the regressive impact of indirect taxes, so that on the whole taxes achieve little redistribution.

The paper has examined also the distributional impact of in-kind transfers. The requisite information is available only for a smaller sample of European countries, and only covers public expenditure on health and education. The analysis is based on the tentative assumption that the volume of in-kind transfers correctly measures their valuation by the individuals who receive them. The main conclusion is that in-kind transfers reinforce considerably the redistributive role of the state in Latin America - where their progressive impact is much bigger than that of cash transfers – but do so even more strongly in Europe. Thus, in-kind transfers actually widen the gap between both regions in terms of the redistributive performance of the state.

Why does Latin America do so poorly at fiscal redistribution? The paper has reviewed three potential explanatory factors, namely: (i) too low a volume of resources gets collected and transferred; (ii) tax collection is regressive; (iii) transfers are poorly targeted. All three are at play, to different extents in different countries, but on the whole the conclusion is that the prospects for significant fiscal redistribution lie mainly in increasing the volume of resources available for redistributive spending, and improving the targeting of expenditures. In contrast, the European experience suggests that even significant increases in the progressivity of Latin America's tax systems - which at present appear to be roughly neutral from the perspective of distribution – are likely to have only a modest effect on the distribution of income. In other words, from the perspective of inequality reduction, the overall volume of tax revenue, which shapes the capacity of the state to engage in redistributive spending, is likely to be a more important priority than the progressivity of the revenue-raising system.

These considerations offer some guidance for the design of reforms to make Latin America's fiscal systems more conducive to equity. In general, such reforms will likely pose significant institutional and implementation challenges that deserve separate analysis. But the evidence presented in this paper suggests that the specific reform priorities vary across countries. In some of them the top priority is to expand tax collection and thereby the volume of transfers; this is likely to be the case, for example, in Mexico and Peru. In most countries, there is significant scope for raising tax collection by reducing tax concessions and loopholes, and especially improving tax administration to reduce evasion, which is rampant across the region.

In other places, the biggest concern is instead to improve the targeting of transfers - indeed, major cash transfer programs, such as social insurance, are in some cases quite regressive, and their reform can make a big difference for overall inequality; this is the case, for example, of Brazil. In such circumstances, raising tax collection, without improving the targeting of the spending it finances, is

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unlikely to help much.

Of course, making the tax system more progressive will also help everywhere, although in general the quantitative impact on inequality is likely to be modest. This does not mean that the structure of the tax system is irrelevant, but only that tax choices should be primarily based on the efficiency and administration costs of different taxes - which lie beyond the scope of the incidence analysis in this paper.

Although we have focused on the public sector's distributive role, social equity is also affected by how well the state does at performing its other two classic objectives of efficiency and stabilization, because ultimately this affects the economic opportunities available to the poor and shapes the distribution of *market* incomes. Through these indirect channels, fiscal policy can also have a major impact on Latin America's inequality. Regarding the stabilization objective of fiscal policy, for example, one important, contribution to social equity relates to the prevention of crises. Macro-financial crises are almost invariably highly regressive because the costs of their resolution, in the form of resource transfers to better-off investors, end up being borne by all taxpayers; furthermore, the poor often are the most adversely affected at times of crises because they lack the assets to smooth out adverse income shocks. Excessive public indebtedness and procyclical fiscal policies have been key factors behind Latin America's vulnerability to crises. This means that fiscal prudence, possibly guided by formal fiscal rules that allow the operation of counter-cyclical policies - and particularly of counter-cyclical social expenditures – is also an essential part of a fiscal agenda to reduce inequality in Latin America.

4.5 Data and Methodology

Our analysis is based upon comparisons of Gini coefficients of different income definitions – market income, gross income, disposable income and after-tax income. To compute these Gini coefficients, we combine two types of information: first, the distribution of disposable income and, second, the incidence (or alternatively the distribution) of the taxes and transfers included in the various income definitions[39]. Such information is drawn from the following sources:

4.5.1 Data for Latin America

Market Income

We draw information on the distribution of disposable income by quintile for Argentina (2001), Brazil (1998), Colombia (2003), Mexico (2000) and Peru (2002) from the World Banks World Development Indicators (WDI). We obtain a measure of market income by quintile by adding direct taxes and deducting transfers for each quintile to the respective disposable income. The information on direct taxes and cash transfers by quintile comes from the sources described in 4.5.1 and 4.5.1.

Cash Transfers

The incidence of cash transfers is estimated on the basis of the information in Lindert & Shapiro (2006) for Argentina, Brazil, Chile, Colombia, Mexico and Peru.

Direct and Indirect Taxes

Direct and indirect tax incidence information is derived from Breceda & Saavedra (2009) for Argentina, Brazil, Colombia, Mexico and Peru.

[39]Given the organization of the original data sources, all the information for Latin American (European) countries considered in our sample is processed by income quintiles (deciles).

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In-kind Transfers

Data on incidence and distribution of public social expenditure on health and education is drawn from ECLAC (2007) for Argentina, Brazil, Chile, Colombia, Mexico and Peru.

Special Case: Chile

Income distribution by income decile as well as direct and indirect tax incidence by income decile is drawn from Table 5 of Engel *et al.* (1999). We infer the corresponding values by quintiles.

4.5.2 Data for European countries

Market Income, Cash Transfers and Direct Taxes

Data on market income, cash transfers and direct taxes comes from EUROMOD, which is a source of harmonized microdata on the different income components before and after redistribution (see `http://www.iser.essex.ac.uk/research/euromod/index/statistics`). We use information available for the year 2001.

Indirect Taxes

We draw simulated data on value added taxes (as % of disposable income) from Baldini & Mantovani (2004). We add this information to the main dataset described in 4.5.2 and thereby obtain a measure of post-tax income.

In-kind Transfers

The series of papers "Distributional effects of imputed rents in seven European countries", prepared as part of the project "Accurate Income Measurement for the Assessment of Public Policies" (see `http://www.iser.essex.ac.uk/research/euromod/research-and-policy-analysis-using-eur aim-ap`), examines the effects of imputed in-kind transfers in seven European countries (Belgium, Germany, Greece, Italy, Ireland, the Netherlands and the UK). From this source we take the data on the increase in disposable income due to in-kind transfers (education and health)[40]. The information corresponds to years between 2002 and 2005, depending on the specific country. We merge this information into the main dataset described in 4.5.2 and obtain measures of gross, disposable and after tax income when cash and in-kind transfers are jointly considered.

[40] We draw information from working papers prepared for each country. The specific references to these working papers as well as a compilation of the main results about the distributional effects of non-cash incomes in the joint sample of European countries under study can be found in Sutherland & Tsakloglou (2009).

Bibliography

ACEMOGLU, D., & ZILIBOTTI, F. 2001. Productivity Differences. *The Quarterly Journal of Economics*, **2**(116), 563–606. [59 and 71]

ACEMOGLU, D., JOHNSON, S., & ROBINSON, J. 2001. The Colonial Origins Of Comparative Development: An Empirical Investigation. *American Economic Review*, 1369–1401. [85]

ACEMOGLU, D., ANTRAS, P., & HELPMAN, E. 2007. Contracts and technology adoption. *The American Economic Review*, **3**(97), 916–943. [57]

AGHION, P., HARRIS, C., HOWITT, P., & VICKERS, J. 2001. Competition, imitation and growth with step-by-step Innovation. *Review of Economic Studies*, 467–492. [71]

AGHION, P., MEGHIR, C., & VANDENBUSSCHE, J. 2006. Growth, distance to frontier and composition of human capital. *Journal of Economic Growth*, 97–127. [59]

AHMAD, I., LEELAHANON, S., & LI, Q. 2005. Efficient Estimation of a Semiparametric Partially Linear Varying Coefficient Model. *The Annals of Statistics*, **33**(1), 258–283. [67]

ALBRECH, J., NAVARRO, L., & VROMAN, S. 2009. The Effects of Labor Market Policies in an Economy with an Informal Sector. *Economic Journal*, **119**(539), 1105–1129. [1]

ALESINA, A., & PEROTTI, D. 1996. Income distribution, political instability, and investment. *European Economic Review*, **6**(40), 1203–1228. [85]

ALESINA, A., & RODRIK, D. 1994. Distribution Politics and Economic Growth. *Quarterly Journal of Economics*, **2**(109), 465–490. [85]

ALVAREZ, F., & VERACIERTO, M. 1999. Labor-Market Policies in an Equilibrium Search Model. *Pages 265–316 of: NBER Macroeconomics Annual*, vol. 14. NBER. [1]

ARBACHE, J. 2004. Does trade liberalization always decrease union bargaining power? *Brazilian Association of Graduate Programs in Economics*, **1**(5), 99–121. [22]

ARELLANO, M., & BOVER, O. 1995. Another look at the instrumental variable estimation of error-components models. *Journal of Econometrics*, **1**(68), 29–51. [16 and 17]

AZARIADIS, C., & DRAZEN, A. 1990. Threshold Externalities in Economic Develpment. *Quarterly Journal of Economics*, **2**(105), 501–526. [58]

AZARIADIS, C., & STACHURSKI, J. 2005. Poverty Traps. *In:* AGHION, P., & DURLAUF, S. (eds), *Handbook of Economic Growth*, vol. 1B. North Holland Elsevier. [85]

BAER, W., & JR, A. GALVAO. 2008. Tax burden, government expenditures and income distribution in Brazil. *Quarterly Review of Economics and Finance*, 345–358. [99]

Goñi-Pacchioni, Edwin Antonio (2011), Essays on Labor and Development Economics
European University Institute DOI: 10.2870/27875

BALDINI, M., C. O'DONOGHUE, & MANTOVANI, D. 2004. *Modelling the redistributive impact of indirect taxes in Europe: an application of EUROMOD*. Working Paper EM7/01. EUROMOD. [95 and 104]

BANK, THE WORLD. 2006. *World Development Report 2006: Equity and Development*. The World Bank. [84, 88, and 93]

BARRO, R., & LEE, J. 2010. *A new data set of educational attainment in the world, 1950-2010*. Working Paper 15902. National Bureau of Economic Research. [60]

BARRO, R., & SALAIMARTIN, X. 1995. *Economic growth*. McGraw-Hill. [56]

BECKER, G., MURPHY, K., & TAMURA, R. 1990. Human Capital, Fertility, and Economic Growth. *Journal of Political Economy*, **5**(98), 512–537. [70]

BERTOLA, G., & ROGERSON, R. 1997. Institutions and Labor Reallocation. *European Economic Review*, **6**(41), 1147–1171. [1]

BETRAND, M, E. DUFLO, & MULLAINATHAN, S. 2004. How Much Should We Trust Differences in Differences Estimates. *Quarterly Journal of Economics*, **1**(119), 249–275. [15 and 16]

BIRDSALL, N., A. DE LA TORRE, & MENEZES, R. 2008. *Fair growth: economic policies for Latin America's poor and middle-income majority*. Center for Global Development. [99]

BIRDSALL, N., & DE LA TORRE, A. 2001. *Washington Contentious: Economic Policies for Social Equity in Latin America*. Carnegie Endowment for International Peace and Inter-American Dialogue. [85]

BLANCHARD, O., & DIAMOND, P. 1991. *The Cyclical Behavior of the Gross Flows of U.S. Workers*. Tech. rept. 1582. NBER Reprints. [35]

BLANCHARD, O., & GALI, J. 2008. *Labor Markets and Monetary Policy: A New-Keynesian Model with Unemployment*. Working Paper. CREI. [32]

BOERI, T., & BURDA, M.C. 1996. Active Labor market Policies, Job Matching and the Czech Miracle. *European Economic Review*, 805–817. [2]

BOSCH, M., & MALONEY, W. 2006. *Gross worker flows in the presence of informal labor markets. The Mexican experience 1987 - 2002*. mimeo. The World Bank. [3, 4, and 33]

BOSCH, M., & MALONEY, W. 2007. *Comparative Analysis of Labor Market Dynamics Using Markov Processes: An Application to Informality*. Discussion Paper 3038. IZA. [35]

BOSCH, M., & MALONEY, W. 2010. Comparative Analysis of Labor Market Dynamics Using Markov Processes: An Application to Informality. *Labour Economics*, **4**(17), 621–631. [10]

BOSCH, M., GOÑI PACCHIONI, E., & MALONEY, W. 2007. *Determinants of Rising Informality in Brazil: Evidence from Gross Worker Flows*. Discussion Paper 2970. IZA. [33, 35, and 36]

BOSCH, M., GOÑI PACCHIONI, E., & MALONEY, W. 2009. *Trade Liberalization, Labor Reforms and Formal-Informal Employment Dynamics*. mimeo. The World Bank. [22, 32, 33, and 35]

BOURGUIGNON, F. 2004. *The Poverty-Growth-Inequality Triangle*. mimeo. The World Bank. [85]

BRAVO-ORTEGA, C., & LEDERMAN, D. 2008. *Intellectual Property Rights, Human Capital and the Incidence of R&D Expenditures*. Serie de Documentos de Trabajo 277. Facultad de Economa y Negocios. Universidad de Chile. [64]

DOI: 10.2870/27875

BRECEDA, K., J. RIGOLINI, & SAAVEDRA, J. 2009. Latin America and the Social Contract: Patterns of Social Spending and Taxation. *Population and Development Review*, **4**(35), 721–748. [103]

CAI, Z., DAS, M., XIONG, H., & WU, X. 2006. Functional coefficient instrumental variables models. *Journal of Econometrics*, 207–241. [68 and 69]

CANOVA, F. 1999. *Testing for Convergence Clubs in Income Per-Capita: A Predictive Density Approach*. mimeo. Universitat Pompeu Fabra. [58]

CARLSSON, M., MESSINA, J., & NORDSTROM, O. 2009. *Wage Adjustment and Productivity: Evidence from Matched Employer-Employee Data*. mimeo. Sveriges Riksbank and The World Bank. [32 and 47]

CARNEIRO, A., GIMARAES, P., & PORTUGAL, P. 2009. *Real Wages and the Business Cycle: Accounting for Worker and Firm Heterogeneity*. mimeo. [47]

CARNEIRO, F. 1999. Insider Power in Wage Determination: Evidence from Brazilian Data. *Review of Development Economics*, **2**(3), 155–169. [37, 38, and 47]

CARNEIRO, F., & HENLEY, A. 1998. Wage Determination in Brazil: The Growth of Union Bargaining Power and Informal Employment. *The Journal of Development Studies*, **4**(34), 117–138. [38]

CARNEIRO, F., & HENLEY, A. 2001. *Modelling Formal vs. Informal Employment and Earnings: Micro-econometric Evidence for Brazila*. mimeo. Universidade Catolica de Brasilia and University of Wales Aberystwyth. [38]

CARNEIRO, F., & HENLEY, A. 2003. *Social Security Reforms and the Structure of the Labor Market: The case of Brazil*. Report to the Inter-American Conference on Social Security. [24]

CARROLL, C., & SAMWICK, A. 1997. The nature of precautionary wealth. *The Journal of Monetary Economics*, 41–71. [49 and 53]

CASELLI, F. 2004. Accounting for Cross-Country Income Differences. *In:* AGHION, P., & DURLAUF, S. (eds), *Handbook of Economic Growth*. [74]

CETRÁNGOLO, O., & GÓMEZ-SABAINI, J. 2007. *La tributación directa en América Latina y los desafíos a la imposición sobre la renta*. ECLAC. [91]

CHAMON, M. 1998. *Rising Wages and Declining Employment: The Brazilian Manufacturing Sector in the 1990s*. Texto Para Discussao 552. IPEA. [22]

CHEN, Y., & PUTTITANUN, T. 2005. Intellectual property rights and innovation in developing countries. *Journal of Development Economics*, 474–493. [59 and 71]

CHIN, J., & GROSSMAN, G. 2004. Intellectual property rights and north-south trade. *In:* JONES, R., & KRUEGER, A. (eds), *The political economy of international trade*. Basil Blackwell, Oxford. [59]

CHU, K, H. DAVOODI, & GUPTA, S. 2000. *Income Distribution and Tax and Government Social Spending Policies in Developing Countries*. Working Paper 00/62. IMF. [86]

COE, D., & HELPMAN, E. 1995. International R&D spillovers. *European Economic Review*, **39**(5), 859–887. [65]

CONLEY, T., HANSEN, C., & ROSSI, P. 2007. *Plausibly Exogenous*. mimeo. Graduate School of Business, University of Chicago. [71]

Goñi-Pacchioni, Edwin Antonio (2011), Essays on Labor and Development Economics
European University Institute DOI: 10.2870/27875

CORSEUIL, C., & FOGUEL, M. 2002. *Uma sugestao de deflactores para rendas obtidas a partir de algumas pequisas domiciliares do IBGE*. Texto Para Discussao 897. IPEA. [20 and 22]

COSTAIN, J., & REITER, M. 2008. Business Cycles, Unemployment Insurance, and the Calibration of Matching Models. *Journal of Economic Dynamics and Control*, 1120–1155. [32]

CROSBY, M. 2000. Patents, innovation and growth. *Economic Record*, 255–262. [71]

DARBY, M., HALTIWANGER, J., & PLANT, M. 1985. Unemployment Rate Dynamics and Persistent Unemployment under Rational Expectations. *Journal of Economic Dynamics and Control*, **4**(75), 614–637. [35]

DARBY, M., HALTIWANGER, J., & PLANT, M. 1986. *The Ins and Outs of Unemployment: The Ins Win*. Working Paper 1997. National Bureau of Economic Research. [35]

DARBY, M., HALTIWANGER, J., & PLANT, M. 1992. Gross Job Creation, Gross Job Destruction, and Employment Reallocation. *Quarterly Journal of Economics*, 819–863. [35]

DAVIS, S., & HALTIWANGER, J. 1990. *Gross Job Creation and Destruction: Microeconomic Evidence and Macroeconomic Implications*. Macroeconomics Annual. National Bureau of Economic Research. [35]

DE CARVALHO, I. 2002. *Old-age benefits and retirement decision of rural elderly in Brazil*. mimeo. MIT. [24]

DE FERRANTI, D., PERRY, G., FERREIRA, F., & WALTON, M. 2004. *Inequality in Latin America & the Caribbean: Breaking with History?* The World Bank. [85, 88, and 99]

DE PAULA, A., & SCHEINKMAN, J. 2009. *The informal sector*. PIER Working Paper 07-035. University of Pennsylvania. [93]

DEARDORFF, A. 1992. Welfare Effects of Global Patent Protection. *Economica*, 35–51. [59]

DEININGER, K., & OLINTO, P. 2000. *Asset distribution, inequality and growth*. Policy Research Working Paper Series 2375. The World Bank. [85]

DEOLALIKAR, A., & LARS-HENDRIK, R. 1989. Patenting by manufacturing firms in India: its production and impact. *Journal of Industrial Economic*, 303–314. [71]

DESDOIGTS, A. 1999. Patterns of Economic Development and the Formation of Clubs. *Journal of Economic Growth*, **3**(4), 305–330. [58]

DETHIER, J. J. 2007. *Social Security: What Can Developing Countries Learn from Developed Countries?* 2020 Focus Brief on the World's Poor and Hungry People. IFPRI. [87]

DONG, W., & YAOI, X. 2007. *Covariance Matrix Repairing in Gaussian Based EDAs*. 2007 IEEE Congress on Evolutionary Computation. [81]

DUONG, T. 2007. *Kernel density estimation for bivariate data*. mimeo. Department of Statistics, University of New South Wales Sydney Australia. [62]

DURLAUF, S. 2001. Manifesto for a growth econometrics. *Journal of Econometrics*, 65–69. [58]

DURLAUF, S., & JOHNSON, P. 1995. Multiple Regimes and Cross-Country Growth Behavior. *Journal of Applied Econometrics*, 365–384. [58]

EBRILL, L., KEEN, M., BODIN, J., & SUMMERS, V. 2001. *The modern VAT*. International Monetary Fund. [94]

 DOI: 10.2870/27875

ECLAC. 2007. *Social Panorama of Latin America*. ECLAC. [99 and 104]

EICHHORN, CH. 2006. *Optimal Policies in the Presence of Tax Evasion*. Inaugural-Dissertation. University of Munich. [92 and 93]

ENGEL, E., GALETOVIC, A., & RADDATZ, C. 1999. Taxes and income distribution in Chile: some unpleasant redistributive arithmetic. *Journal of Development Economics*, 155–192. [86, 90, and 104]

ENGERMAN, S., & SOKOLOFF, K. 2000. Institutions, Factor Endowments, and Paths of Development in the New World. *Journal of Economic Perspectives*, 217–232. [85]

FAN, J., & GIJBELS, I. 1995. Data-driven bandwidth selection in local polynomial fitting: Variable bandwidth and spatial adaption. *Journal of the Royal Statistical Society B*, **2**(57), 371–394. [69 and 80]

FAN, J., & ZHANG, W. 1999. Statistical Estimation in Varying Coefficient Models. *The Annals of Statistics*, **5**(27), 1491–1518. [67]

FAN, J., & ZHANG, W. 2000. Simultaneous Confidence Bands and Hypothesis Testing in Varying Coefficient Models. *Scandinavian Journal of Statistics*, 715–731. [67]

FARMER, R. 2006. *Shooting the Auctioneer*. mimeo. UCLA. [32]

FULLERTON, D., & ROGERS, D. 1993. *Who Bears the Lifetime Tax Burden?* Brookings Institution. [86]

GALOR, O. 1996. Convergence? Inferences from Theoretical Models. *Economic Journal*, 1056–1069. [58]

GALOR, O., & WEIL, D. 2000. Population, technology, and growth: From Malthusian stagnation to the demographic transition andbeyond. *American Economic Review*, 806–829. [70]

GERTLER, M., & TRIGARI, A. 2009. Unemployment Fluctuations with Staggered Nash Wage Bargaining. *Journal of Political Economy*, **1**(117), 38–86. [32 and 47]

GEWEKE, J., MARSHALL, R., & ZARKIN, G. 1986. Mobility indices in continuous time Markov chains. *Econometrica*, **6**(54), 1407–1423. [36]

GINARTE, J., & PARK, W. 1997. Determinants of Patent Rights: A Cross-National Study. *Research Policy*, 283–301. [60 and 71]

GOLDBERG, P., & PAVCNIK, N. 2003. The Response of the Informal Sector to Trade Liberalization. *Journal of Development Economics*, 463–496. [2, 3, 4, 10, 16, and 18]

GÓMEZ-SABAINI, J. 2006. Evolución y situación tributaria actual en América Latina: Una serie de temas para la discusión. *In: Tributación en América Latina: en busca de una nueva agenda de reformas*. [92, 93, and 94]

GOULD, D., & GRUBEN, W. 1997. The role of intellectual property rights in economic growth. *In:* DEV GUPTA, S., & CHOUDHRY, N. (eds), *Dynamics of Globalization and Development*. Kluwer Academic Publishers. [71]

GRIFFITH, R., REDDING, S., & REENEN, J. VAN. 2004. Mapping the Two Faces of R&D: Productivity Growth in a Panel of OECD Industries. *Review of Economics and Statistics*, **4**(86), 883–895. [56, 59, 65, and 71]

 DOI: 10.2870/27875

GRILICHES, Z. 1986. Productivity, R&D and basic research at the firm level in the 1970's. *American Economic Review*, **1**(76), 141–154. [64]

GRILICHES, Z. 1995. R&D and Productivity: Econometric Results and Measurement Issues. *In:* STONEMAN, P. (ed), *Handbook of the Economics of Innovation and Technological Change*. Blackwell, Oxford, England and Cambridge, Massachusetts. [64]

GRILICHES, Z., & MAIRESSE, J. 1990. R&D and Productivity Growth: Comparing Japanese and US Manufacturing Firms. *Pages 317–340 of:* HULTEN, C. (ed), *NBER Studies in Income and Wealth*. NBER and University of Chicago Press. [34]

GRILLICHES, Z., & LICHTENBERG, F. 1984. R&D and Productivity Growth at the Industry Level: Is there Still a Relationship? *In:* GRILLICHES, Z. (ed), *R&D, Patents and Productivity*. University of Chicago Press. [65]

GROSSMAN, G., & HELPMAN, E. 1991. Endogenous Product Cicles. *Economic Journal*, 1214–1229. [59]

GUISO, L., PISTAFERRI, L., & SCHIVARDI, F. 2005. Insurance within the firm. *The Journal of Political Economy*, **5**(113), 1054–1087. [53]

HAEFKE, C., SONNTAG, M., & VAN RENS, T. 2008. *Wage Rigidity and Job Creation*. Discussion Paper 3714. IZA. [32, 39, 45, and 47]

HALL, B., & MAIRESSE, J. 1995. Exploring the relationship between R&D and productivity in French manufacturing firms. *Journal of Economterics*, **65**(1), 263–293. [65 and 82]

HALL, R. 1995. *Lost Jobs*. Tech. rept. 1. [35]

HALL, R., & JONES, C. 1999. Why do some countries produce so much more output per worker than others? *The Quarterly Journal of Economics*, **114**(1), 83–116. [74]

HALL, R., & MILGROM, P. 2008. Employment Efficiency and Sticky Wages: Evidence Flows in the Labor Market. *American Economic Review*. forthcoming. [32]

HANSEN, B. 2000. Sample Splitting and Threshold Estimation. *Econometrica*, 575–603. [58]

HARBERGER, A. 2003. *Reflections on Distributional Considerations and the Public Finances*. mimeo. University of California, Los Angeles. [86 and 90]

HASTIE, T., & TIBSHIRANI, R. 1993. Varying-Coefficient Models. *Journal of the Royal Statistical Society. Series B (Methodological)*, **4**(55), 757–796. [67]

HAY, D. 2001. The post 1990 Brazilian trade liberalization and the performance of large manufacturing firms: productivity, market share and profits. *The Economic Journal*, 624–641. [10]

HEATHCOTE, J., STORESLETTEN, K., & VIOLANTE, G. 2008. *The macroeconomics implications of rising wage inequiality in the United States*. Tech. rept. 14052. [53]

HENLEY, A., ARABSHEIBANI, G., & CARNEIRO, F. 2009. On defining and measuring the informal sector: Evidence from Brazil. *World Development*, **5**(37), 992–1003. [3]

HOPENHAYN, H. 2004. Labor market policies and employment duration:the effects of labor market reform in Argentina. *Pages 497–516 of:* HECKMAN, J. J., & PAGÉS, C. (eds), *Law and Employment lessons from Latin America and the Caribbean*. NBER and University of Chicago Press. [2]

IBGE. 1991. *Para Compreender a PME*. IBGE. [2 and 33]

 DOI: 10.2870/27875

JACK, W. 2000. *Health Insurance Reform in four Latin American countries. Theory and practice*. Policy Research Working Paper 2492. The World Bank. [24]

JIMÉNEZ, J, GÓMEZ-SABAINI, J., & PODESTÁ, A. 2010. *Tax gap and equity in Latin Amrica and the Caribbean*. Fiscal Studies 16. ECLAC. [92]

JONES, C., & WILLIAMS, J. 1998. Measuring the Social Return to R&D. *The Quarterly Journal of Economics*, 1119–1135. [63, 64, and 65]

JONES, F. 2008. *The effects of taxes and benefits on household income, 2006-07*. Tech. rept. Office for National Statistics, UK. [95]

KOURTELLOS, A. 2001. *Econometric Modeling of Heterogeneity and Nonlinearity With an Application to Cross-Country Growth Data*. mimeo. University of Wisconsin. [58]

KOURTELLOS, A. 2002. *Modeling Coefficient Heterogeneity in Cross-Country Growth Regression Models*. mimeo. Department of Economics, Virginia Tech. [58]

KRAY, A. 2008. *Instrumental Variables Regressions with Honestly Uncertain Exclusion Restrictions*. mimeo. The World Bank. [71]

KREBS, T., KRISHNA, P., & MALONEY, W. 2010. *Income Dynamics, Mobility and Welfare Developing Countries*. mimeo. The World Bank. [53]

KUCERA, D., & RONCOLATO, L. 2008. Informal Employment: Two Contested Policy Issues. *International Labour Review*, **4**(147), 321–348. [20]

KUCZYNSKI, P., & WILLIAMSON, J. 2003. *After the Washington Consensus: Restarting Growth and Reform in Latin America*. Institute for International Economics. [85]

KUGLER, A. 1999. The Impact of Firing Costs on Turnover and Unemployment: Evidence from the Colombian Labour Market Reform. *International Tax and Public Finance*, 389–410. [2]

KUGLER, A. 2004. The Effect of Job Security Regulations on Labor Market Flexibility: Evidence from the Colombian Labor market reform. *Pages 183–228 of:* HECKMAN, J. J., & PAGÉS, C. (eds), *Law and Employment lessons from Latin America and the Caribbean*. NBER and University of Chicago Press. [1, 2, 14, and 29]

KUGLER, A., & PICA, G. 2008. Effects of Employment Protection on Worker and Job Flows: Evidence from the 1990 Italian Reform. *Labour Economics*, 78–95. [2]

KUME, H., PIANI, G., & DE SOUZA, C. BRZ. 2003. A política brasileira de importacão no período 1987-98: descricão e avaliacão. *In:* CORSEUIL, C., & KUME, H. (eds), *A abertura comercial brasileira nos anos 1990: impactos sobre emprego e salário*. IPEA. [10, 11, and 12]

LACH, S. 1995. Patents and productivity growth at the industry level: a first look. *Economics Letters*, 101–108. [71]

LEDERMAN, D., & MALONEY, W. 2005. *R&D and Development*. mimeo. The World Bank. [56, 59, 60, 61, 64, 65, 71, and 73]

LEDERMAN, D., & SAENZ, L. 2005. *Innovation and Development around the World, 1960-2000*. Policy Research Working Paper 3774. The World Bank. [59]

LEMOS, S. 2009. Minimum wage effects in a developing country. *Labour Economics*, 992–1003. [24]

Goñi-Pacchioni, Edwin Antonio (2011), Essays on Labor and Development Economics
European University Institute DOI: 10.2870/27875

LEVY, S. 2008. *Good intentions, bad outcomes: social policy, informality and economic growth in Mexico*. Brookings Institution. [93]

LINDERT, K., SKOUFIAS E., & SHAPIRO, J. 2006. *Redistributing Income to the Poor and the Rich: Public Transfers in Latin America and the Caribbean*. SP Discussion Paper 0605. The World Bank. [96, 99, and 103]

LIU, C., & STENGOS, T. 1999. Non-Linearities in Cross-Country Growth Regressions: A Semi-parametric Approach. *Journal of Applied Econometrics*, **5**(14), 527–538. [58]

LOAYZA, N., SERVÉN L., & SUGAWARA, N. 2009. Informality in Latin America and the Caribbean. *In:* LOAYZA, N., & SERVÉN, L. (eds), *Business regulation and macroeconomic performance*. The World Bank. [92]

LÓPEZ, H., & SERVÉN, L. 2006a. *A normal relationship? Poverty, growth, and inequality*. Policy Research Working Paper Series 3814. The World Bank. [85]

LÓPEZ, H., & SERVÉN, L. 2006b. *Too poor to grow*. Policy Research Working Paper Series 5012. The World Bank. [85]

LUCAS, R. 1988. On the Mechanics of Economic Development. *Journal of Monetary Economics*, 3–42. [57]

MALONEY, W., G. PERRY O. ARIAS P. FAJNZYLBER A. MASON, & SAAVEDRA, J. 2007. *Informality: Exit and Exclusion*. The World Bank. [92]

MALONEY, W., & NUNEZ, E. 2001. *Minimum Wages in Latin America*. Policy Research Working Paper. The World Bank. [24]

MALONEY, W., & RODRIGUEZ-CLARE, A. 2007. Innovation Shortfalls. *Review of Development Economics*, **4**(11), 665–684. [58]

MALONEY, W. F. 1999. Does Informality Imply Segmentation in Urban Labor Markets? Evidence from Sectoral Transitions in Mexico. *The World Bank Economic Review*, **2**(13), 275–302. [10]

MANSFIELD, E. 1965. Rates of return from industrial research and development. *The American Economic Review*, **2**(55), 310–322. [64]

MARTNER, R., & ALDUNATE, E. 2006. *Política fiscal y protección social*. Gestión Pública 53. ECLAC. [92 and 93]

MASKUS, K. 2000. *Intellectual Property Rights in the Global Economy*. [59 and 71]

MASKUS, K., & MCDANIEL, C. 1999. Impacts of the Japanese patent system on productivity growth. *Japan and the World Economy*, 557–574. [71]

MASKUS, K., & PENUBARTI, M. 1995. How trade-related are intellectual property rights? *Journal of International Economics*, **3**(39), 227–248. [71]

MEGHIR, C., & PISTAFERRI, L. 2004. Income Variance Dynamics and Heterogenity. *Econometrica*, **1**(71), 1–32. [49 and 53]

MENEZES-FILHO, N., & MUENDLER, M. 2007. *Labor Reallocation in Response to Trade Reform*. Working Paper Series 1936. CESifo. [2 and 10]

MENEZES-FILHO, N., MUENDLER, M., & RAMEY, G. 2008. The structure of worker compensation in Brazil with a comparison to France and the United States. *The Review of Economics an Statistics*, **2**(90), 324–346. [34 and 39]

 DOI: 10.2870/27875

MENZIO, G. 2005. *High Frequency Wage Rigidity*. mimeo. University of Pennsylvania. [32]

MESSINA, J., & VALLANTI, G. 2007. Job Flow Dynamics and Firing Restrictions: Evidence from Europe. *The Economic Journal*, 279–301. [2]

MOEN, E., & ROSEN, A. 2006. *Incentives in Competitive Search Equilibrium and Wage Rigidity*. Discussion Paper 5554. CEPR. [32]

MORTENSEN, D., & PISSARIDES, C. 1994. Job Creation and Job Destruction in the Theory of Unemployment. *Review of Economic Studies*, **3**(61), 397–415. [26]

MORTENSEN, D., & PISSARIDES, C. 2003. Taxes, Subsidies and Equilibrium Labor Market Outcomes. *In:* PHELPS, E. (ed), *Designing Inclusion: Tools to Raise Low-end Pay and Employment in Private Enterprise*. Cambridge University Press. [1]

MUENDLER, M. 2004. *Trade, Technology, and Productivity: A Study of Brazilian Manufacturers, 1986-1998*. mimeo. UCSD and CESifo. [10, 11, 34, and 45]

MUSGRAVE, R. 1959. *The Theory of Public Finance: Study in Public Economy*. McGraw-Hill. [86]

NASSIF, A., & PIMENTEL, F. 2004. Estructura e competitividade da Industria Brasileira: o que mudou? *Revista do BNDES*, **22**(2), 3–19. [11]

NELSON, R., & PHELPS, E. 1966. Investment in humans, technological diffusion and economic growth. *American Economic Review*, **2**(56), 69–75. [74]

NERI, M., GONZAGA, G., & CAMARGO, J. 2000. *Efeitos Informais do Salrio Mnimo e Pobreza*. Texto para Discussao 724. IPEA. [24]

NEUMARK, D., CUNNINGHAM, W., & SIGA, L. 2004. The effects of the minimum wage in Brazil on the distribution of family incomes: 1996-2001. *Journal of Development Economics*, **80**(1), 136–159. [24]

NEWEY, W., POWELL, J., & VELLA, F. 1999. Nonparametric estimation of triangular simultaneous equations models. *Econometrica*, 565–603. [69]

OECD. 2008. *Growing Unequal? Income Distribution and Poverty in OECD Countries*. OECD. [86 and 90]

OLIVEIRA, A. 1999. *Relatório Metodológico: Microdados da Pesquisa Mensal de Emprego*. mimeo. CEDEPLAR and UFMG. [2 and 33]

OULTON, N., & YOUNG, G. 1996. How high is the social rate of return to investment? *Oxford Review of Economic Policy*, **2**(12). [69]

PACK, H. 1994. Endogenous Growth Theory: Intellectual Appeal and Empirical Shortcomings. *Journal of Economic Perspectives*, **1**(8), 55–72. [58]

PAES DE BARROS, R., & CORSEUIL, C. 2004. The Impact of Regulations on Brazilian Labor Market Performance. *Pages 273–350 of:* HECKMAN, J. J., & PAGÉS, C. (eds), *Law and Employment lessons from Latin America and the Caribbean*. NBER and University of Chicago Press. [2, 11, 12, 18, and 35]

PAN-AMERICAN-HEALTH-ORGANIZATION. 2005. *Brazil health system profile*. PAHO. [24]

PARK, W. 2001. Intellectual Property and Patent Regimes. *In: Economic Freedom of the World: 2001 Annual Report*. Freedom House: New York. [60 and 64]

 DOI: 10.2870/27875

PARK, W. 2008. International patent protection: 19602005. *Research Policy*, 761–766. [60 and 64]

PECHMAN, J. 1985. *Who Paid the Taxes: 1966-85?* The Brookings Institution. [95]

PECHMAN, J. 1990. The Future of Income Tax. *American Economic Review*, 1–20. [95]

PERRY, G., MALONEY, W., ARIAS, O., FAJNZYLBER, P., MASON, A., & SAAVEDRA, J. 2007. *Informality: Exit and Exclusion*. The World Bank. [48]

PERRY, G., O. ARIAS H. LÓPEZ W. MALONEY, & SERVÉN, L. 2006. *Poverty reduction and growth: virtuous and vicious circles*. The World Bank. [85 and 91]

PESSINO, C., & FENOCHIETTO, R. 2004. *Efficiency and Equity of the Tax Structure in Argentina, Brazil and Chile: Analysis and Policy Considerations*. Center for Global Development. [92]

PINHEIRO, A., & DE ALMEIDA, G. BACHA. 1994. *Padrões Sectoriais da Protecão na Economia Brasileira*. Texto para discussão 355. IPEA. [11 and 12]

PISSARIDES, C. 2000. *Equilibrium Unemployment Theory*. Cambridge: MIT Press. [27]

PISSARIDES, C. 2009. The Unemployment Volatility Puzzle: Is Wage Stickiness the Answer? *Econometrica*, **5**(77), 1339–1369. [32 and 37]

PRIES, M., & ROGERSON, R. 2005. Hiring Policies, Labor Market Institutions and Labor Market Flows. *Journal of Political Economy*, **4**(113), 811–839. [1]

PRIMO BRAGA, C., & FINK, C. 2000. International transactions in intellectual property and developing countries. *International Journal of Technology Management*, **1**(19), 35–56. [59]

PRITCHETT, L. 2001. Where has all the education gone? *The World Bank Economic Review*, **3**(15), 367–391. [75]

QUAH, D. 1997. Empirics for Growth and Distribution: Stratification, Polarization, and Convergence Clubs. *Journal of Economic Growth*, 27–59. [61]

RAMOS, L. 2002. *A evolução da informalidade no Brasil metropolitano: 1991-2001*. Texto para discussão 914. IPEA. [3, 4, and 33]

RAMOS, L., & BRITO, M. 2003. *O funcionamento do mercado de trabalho metropolitano brasileiro no período 1991-2002: tendências, fatos estilizados e mudanas estruturais*. Nota tecnica. Mercado de trabalho. IPEA. [3 and 33]

RAMOS, L., & FERREIRA, V. 2005a. *Padrões Espacial e Setorial da Evolução da Informalidade No Período Pós-Abertura Comercial*. mimeo. IPEA. [3, 25, and 33]

RAMOS, L., & FERREIRA, V. 2005b. *Padrões Espacial e Setorial da Evolucão da Informalidade no Brasil 1991-2003*. Texto para discussão 1099. IPEA. [3, 25, and 33]

RAMOS, L, & REIS, J. 1997. *Grau de formalizaão, nível e qualidade do emprego no mercado de trabalho metropolitano do Brasil*. Mercado de Trabalho - Conjuntura e Análise 5. IPEA. [3 and 33]

ROMER, P. 1990. Endogenous Technological Change. *The Journal of Political Economy*, **5**(98), S71–S102. [57]

RUDANKO, L. 2008. *Labor Market Dynamics under Long Term Wage Contracting*. mimeo. Boston University. [32]

Goñi-Pacchioni, Edwin Antonio (2011), Essays on Labor and Development Economics
European University Institute
DOI: 10.2870/27875

SCHUMPETER, J. 1934. *The Theory of Economic Development*. [57]

SEDLACEK, G., DE BARROS, R. PAES, & VARANDAS, S. 1990. Segmentaão e Mobilidade no Mercado de Trabalho: A Carteira de Trabalho em São Paulo. *Pesquisa e Planejamento Econômico*, **1**(20), 87–103. [2 and 33]

SHIMER, R. 2005a. The Cyclical behavior of Equilibrium Unemployment and Vacancies. *American Economic Review*, **1**(95), 25–49. [35]

SHIMER, R. 2005b. *Reassessing the Ins and Outs of Unemployment*. mimeo. University of Chicago. [37]

SHIMER, R. 2009. *Labor Markets and Business Cycles*. Lectures in Macroeconomics. CREI. forthcoming. [32]

SHOME, P. 1995. Recent Tax Policy Trends and Issues in Latin America. *In: Policies for Growth: The Latin American Experience*. International Monetary Fund. [93]

SMEEDING, T. 2006. Government Programs and Social Outcomes: Comparison of the United States with Other Rich Nations. *In:* AUERBACH, A., CARD, D., & QUIGLEY, J. (eds), *Public Policy and Income Distribution*. Russell Sage Foundation. [86]

SOARES, F. 2005. *The Impact of trade liberalization on the informal sector in Brazil*. Working Paper 7. International Policy Centre for Inclusive Growth. [10]

SOLON, G., BARSKY, R., & PARKER, J. 1994. Measuring the Cyclicality of Real Wages: How Important is Composition Bias? *Quarterly Journal of Economics*, **1**(109), 1–25. [39 and 45]

SOUZA, P., & BALTAR, P. 1979. *Salario Minimo e a Taxa de Salario no Brasil*. Working Paper 9. Planejamento Economico. [24]

STOTSKY, J., & WOLDEMARIAM, A. 2002. *Central American Tax Reform: Trends and Possibilities*. Working Paper 02/227. IMF. [94]

SUTHERLAND, H., A. DECOSTER M. MATSAGANIS, & TSAKLOGLOU, P. 2009. *Accurate income measurement for the assessment of public policies.Final Report*. Working Paper EM7/09. EUROMOD. [104]

TANZI, V. 2000. *Taxation in Latin America in the Last Decade*. Working Paper 76. Center for Research on Economic development and Policy Reform. Stanford University. [93 and 94]

THE-WORLD-BANK, & IPEA. 2002. *Brazil Jobs Report*. Policy Briefing. The World Bank and IPEA. [3]

THOMPSON, M., & RUSHING, F. 1996. An empirical analysis of the impact of patent protection on economic growth. *Journal of Economic Development*, 61–79. [71]

THOMPSON, M., & RUSHING, F. 1999. An empirical analysis of the impact of patent protection on economic growth: an extension. *Journal of Economic Development*, 67–76. [71]

ULKU, F. 2005. Dressing the Wound: Organizing Informal Sector Workers. *Monthly Review*, **1**(57). [4]

ULYSSEA, G. 2005. *Informalidade no mercado de Trabalho brasileiro: uma Resenha da literatura*. Texto para discussão 1070. IPEA. [3 and 33]

VERAS, F. 2004. *Some stylized facts of the informal sector in Brazil in the 1980s and 1990s*. Texto para discussão 1020. IPEA. [3 and 33]

YOUNG, A. 1993a. Substitution and Complementarity in Endogenous Innovation. *The Quarterly Journal of Economics*, **3**(108), 775–807. [57]

YOUNG, A. 1993b. Substitution Invention and Bounded Learning by Doing. *Journal of Political Economy*, **3**(101), 443–472. [57]

ZENOU, Y. 2008. Job Search and Mobility in Developing Countries. Theory and Policy Implications. *Journal of Development Economics*, 336–355. [1]

ZHANG, W., & LEE, S. 2000. Variable bandwidth selection in varying-coefficient models. *Journal of Multivariate Analysis*, 116–134. [69]

Goñi-Pacchioni, Edwin Antonio (2011), Essays on Labor and Development Economics
European University Institute DOI: 10.2870/27875