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Equality under Threat by the Talented: Evidence from Worker-Managed Firms

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ABSTRACT

Equality under Threat by the Talented: Evidence from Worker-Managed Firms*

Are high-ability individuals more likely to quit egalitarian regimes? Does the threat of exit by talented individuals restrict the redistributive capacity of democratic organizations? This paper revisits that long-standing debate by analyzing the interplay between compensation structure and quit behavior in the distinct yet underexplored institutional setting of worker-managed firms. The study exploits two novel administrative data sources: a panel of Uruguayan workers employed in both worker-managed and conventional firms; and a linked employer–employee panel data set covering the population of Uruguayan worker-managed firms and their workers from January 1997 to April 2010. A key advantage of the data is that it enables one to exploit within-firm variation on wages to construct an ordinal measure of the worker ability type. The paper's four main findings are that (1) worker-managed firms redistribute in favor of low-wage workers; (2) in worker-managed firms, high-ability members are more likely than other members to exit; (3) the hazard ratio of high-ability members is lower for founding members and for those employed by worker-managed firms in which there is less pay compression; and (4) high-ability members are less likely to quit when labor market conditions in the capitalist sector are less attractive. This paper contributes to the study of the interplay between equality and incentives that permeates many debates in public finance, comparative economic systems, personnel and organizational economics.

JEL Classification: H00, J54, J62, M52, P0

Keywords: labor managed firms, redistribution, compensation structure, job mobility

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I. INTRODUCTION

The potential conflict between equality and the need for incentives is a major debate in economics and political philosophy.[1] Are high-ability individuals more likely to quit from egalitarian regimes? Is the redistributive capacity of democratic organizations restricted by the possibility that talented agents may exit? I revisit this long-standing debate by analyzing the relationship between compensation structure and quit behavior in a unique and underexplored institutional setting: worker-managed firms (WMFs).

Most economic activities in actual market economies are carried out by conventional firms (CFs) controlled by capital suppliers. In contrast, WMFs are defined as enterprises in which the workforce has ultimate control rights (Dow, 2003).[2] Worker-managed firms are democratic in the sense that members have equal political influence on economic decisions regardless of their capital contribution to the firm ("one person, one vote"). This type of firm captured the attention of such renowned economists Karl Marx, John Stuart Mill, Leon Walras, and Alfred Marshall. Since the late 1950s, an extensive theoretical literature has developed that seeks to understand the behavior of WMFs and to explain why they are relatively rare.[3]

One prominent explanation for the paucity of WMFs is that workplace democracy may result in substantial redistribution at the expense of high-ability workers. Median voter models suggest that, to the extent the median member is less productive than the average, most cooperative members can gain by reducing wage differences relative to differences in productivity (Kremer, 1997). Another explanation is that equality may provide insurance against unfavorable realizations of ability (Abramitzky, 2008). Irrespective of the precise mechanism behind egalitarian compensation policies in WMFs, both models predict that equality discourages the participation of high-ability members. However, the actual extent and effects of redistribution in WMFs have not been systematically studied.

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[1] The Rawlsian difference principle states that "social and economic inequalities, for example inequalities of wealth and authority, are just only if they result in compensating benefits for everyone, and in particular for the least advantaged members of society" (Rawls, 1971: 14–15). For a critique, see Cohen (1992).

[2] Other terms for WMFs that are used in the literature include worker cooperatives, producer cooperatives, and democratic firms.

[3] For a review of the literature see Bonin, Jones, and Putterman (1993), Dow and Putterman (2000), Dow (2003), and Putterman (2008). The most updated evaluation of the empirical literature is provided by Pencavel (2013).

This paper contributes to filling this gap by examining three interrelated questions. Do WMFs actually engage in redistributive compensation policies? Are high-ability members in WMFs more likely (than other members) to exit? Does the degree of equality affect the severity of brain drain? The empirical analysis is based on novel work history data from Uruguayan social security administrative records. To answer the first question, I use a panel of workers employed in both worker-managed and conventional firms. To address the second and third questions, I use a matched employer–employee panel data set that includes information on the total population of firms legally registered as producer cooperatives (PCs)—from which WMFs can be identified—and all their workers, both members and nonmembers. One major advantage of the latter data set is that I can observe the entire wage distribution (and other characteristics of the workforce) at each firm for any moment in time. This makes it possible to rank the ability of workers, including quitters, according to their position in the intrafirm wage distribution. Both panels are unusually long and extend over the period from January 1997 to April 2010.

The analysis yields four main results. First, I find a small wage premium associated with being employed in WMFs. Because there is mobility between worker-managed and conventional firms, identification rests on the variability provided by workers who switch between organizational types during the period—under the assumption that sorting is based on time-invariant characteristics. It is noteworthy that this wage gap decreases across the wage distribution. Quantile regression estimates confirm that WMFs do implement redistributive compensation policies. Second, estimates derived from duration models indicate that the high-ability members of WMFs exhibit a higher hazard rate of voluntary separation. Third, in WMFs that are more egalitarian, high-ability members are more likely to exit; in other words, the brain drain's severity depends on how compressed the pay scale is. I also find that the hazard ratio of high-ability members is lower in the case of founding members; this suggests that ideological commitment—which may be stronger for first-generation members—enables greater redistribution within WMFs. Fourth, the quit decisions of high-ability WMF members seem especially sensitive to labor market conditions in the capitalist sector. Higher unemployment rates and lower wages paid in the

conventional sector (relative to WMF members' compensation) significantly reduce the exit rate of high-ability members.

At a more general level, this paper contributes to the study of the interplay between equality and incentives that permeates many debates in public finance, development, comparative economic systems, human resources and organizational economics. First, it is related to a series of recent studies on equal-sharing rules and migration in communes, particularly in Israeli kibbutzim (Abramitzky, 2008, 2009, 2011). The paper adds to this literature in several ways. Kibbutzim studies have relied on self-reported measures of the degree of internal equality and have tested brain drain by comparing quitters to stayers in terms of education and skill levels, not in terms of their wages. Moreover, they have not investigated whether kibbutzim that shift away from equal-sharing rules do in fact reduce their brain drain. By contrast, I use matched organization–worker panel data that gives the entire wage distribution of each WMF and exploit within-firm variation in intrafirm wage dispersion to analyze how organizations use compensation policies to cope with brain drain. The interest in worker-managed firms rests on the fact that these organizations have existed (alongside investor-controlled firms) in most Western economies since the Industrial Revolution. Yet even though WMFs are thus a realistic organizational alternative to capitalist firms, they are usually found only in certain sectors (e.g., professional partnerships, taxis) and regions.[4] The paucity of WMFs, especially in labor-intensive sectors, remains a puzzle.

Second, the choice of a compensation structure and its effect on the retention of valuable employees is a core topic in personnel economics (Lazear and Shaw, 2007; Lazear and Oyer, 2013). Third, the paper is also related to the public economics literature on how mobility constrains redistributive taxation (Simula and Trannoy, 2010; Kleven, Landais, and Saez, 2013; Rothschild and Scheuer, 2013). The case of WMFs illustrates how egalitarian schemes are threatened when some individuals have attractive exit options and so can "vote with their feet".

[4] According to Arando et al. (2012), worker-managed firms account for 13% of economic activity in the northern Italian province of Emilia Romagna and 8% of industrial gross value added (and 4% of overall gross value) in the Basque Country, Spain, where the Mondragon Cooperative Corporation is located.

Fourth, this paper contributes directly to the literature on WMFs. In contrast to early neoclassical models, more recent theoretical approaches emphasize the role of labor discipline, credit market imperfections, and collective choice problems associated with determining distributional rules in heterogeneous WMFs (Bowles and Gintis, 1993, 1994; Kremer, 1997).[5] That being said, hardly any attention has been given to how members' heterogeneity and democratic governance actually interact in such firms (Pencavel, 2013). There is some extant research comparing worker-managed and conventional firms in terms of productivity, wage and employment adjustments, and firm demography; however, this study is one of the first to assess the extent and effects of redistribution in WMFs.[6] Participatory workplaces may use pay compression to enhance cohesiveness and teamwork, but the evidence presented here suggests that such greater equality has the negative side effect of brain drain (Levine, 1991). The role played by members' heterogeneity and distributional conflicts in organizations has received attention in development economics (Banerjee et al., 2001). In contrast to this literature, which investigates whether land inequality triggers distributional conflicts and inefficiencies in the context of agricultural cooperatives, I study WMFs that operate in nonagricultural sectors and in which political conflict is likely to be structured by differences in members' ability (Kremer, 1997).

The rest of the paper is organized as follows. Section II briefly discusses the related literature, and Section III provides contextual information on Uruguayan worker-managed firms and describes the data, and Section IV presents the main results. Section V concludes and discusses the implications of these results in terms of the organizational performance of WMFs.

II. DEMOCRATIC GOVERNANCE, COMPENSATION STRUCTURE, AND INCENTIVES UNDER WORKERS' CONTROL

There are few theoretical contributions that address the relationship between redistribution and incentives in worker-managed firms. Early neoclassical models assumed that the objective of a WMF is to maximize net income per worker, and they ruled out problems associated with democratic governance (see e.g. Ward, 1958). More recently, theoretical

[5] This theoretical literature will be discussed in Section II.

[6] See, for example, Craig and Pencavel (1992), Pencavel and Craig (1994), Craig and Pencavel (1995), Pérotin (2006), Burdin and Dean (2009), Fakhfakh, Pérotin, and Gago (2012), and Burdin (2013).

models have departed from the standard "complete contracting" assumptions in labor and credit markets. The main implication of these models is that WMFs should have a competitive advantage in regulating labor effort but should be credit constrained owing to the lack of workers' wealth for use as collateral (Bowles and Gintis, 1994). By assuming a homogeneous workforce, these models also eliminate the possibility of collective choice problems within WMFs and hence fail to account for the predominance of conventional firms even in labor-intensive industries (Kremer, 1997).

The problems faced by WMFs with an heterogeneous workforce have been discussed at length by Hansmann (1988, 1996). The author points out that the costs of collective decision making constitute the main disadvantage of this organization type vis-à-vis conventional enterprises. These costs are increasing in the members' heterogeneity. Whereas capital suppliers unanimously support the objective of maximizing profit, workers may have different attitudes regarding effort, investment decisions, wage levels, job security, and the provision of other workplace amenities. In a democratically controlled firm, workers must rely on some mechanism—typically, a majority voting rule—to aggregate their preferences. Unless the preferences of the median voter coincide with those of the mean voter (which is seldom the case), the resulting decisions may be inefficient in the sense of not maximizing organizational surplus (Hansmann, 1996). Organizational design in worker-managed firms may limit the diversity of preferences to ensure workforce homogeneity. Large inequalities among members may destabilize a cooperative governance structure even as internal rules adopted to minimize the costs of collective governance may also result in efficiency problems (Benham and Keefer, 1991).

The determination of the compensation structure is probably the most important collective choice problem faced by WMFs.[7] Kremer (1997) proposes a median voter model with heterogeneously productive members. He shows that, to the extent the median member is less productive than the average member, the majority of a cooperative's members are better-off if wage differences are reduced with respect to productivity differences. A WMF

[7] Earlier theoretical analyses of how distribution rules affect WMFs include Sen (1966) and Gui (1987).

will therefore set an egalitarian wage structure and then will have problems retaining high-ability members.[8]

Abramitzky (2008) develops a theoretical framework to study the effect of equal-sharing rules in Israeli kibbutzim.[9] In the first period, ex ante identical individuals make a sunk contribution to the kibbutz and set a sharing rule (the degree of equality). In the second period, individuals learn about their own productivity and decide whether or not to remain in the kibbutz. Equality provides insurance but discourages high-ability individuals from remaining in the organization (the brain drain effect). The model also predicts that, the higher the value of total assets, the greater the ability of the kibbutz to maintain an equal-sharing rule. Common property operates as a lock-in device because, as in collectively owned WMFs, members who leave the kibbutz have no claims on its assets.[10]

Unfortunately, there is little evidence on the extent and effects of redistribution in WMFs. There is some anecdotal evidence supporting the view that worker-managed firms implement more egalitarian wage distributions than do conventional firms. For instance, US plywood cooperatives have generally relied on equal-pay schedules regardless of seniority or tasks performed (Pencavel, 2001). Mondragon cooperatives located in the Basque Country, Spain, have strict regulations concerning maximum wage differentials (Dow, 2003), which is in line with survey evidence on WMFs located in the Italian province of Emilia Romagna (Estrin and Holmes, 1991; Bartlett et al., 1992). Most of the evidence on the incentive effects of egalitarian arrangements comes from the recent literature on kibbutzim. Abramitzky (2008) presents evidence supporting the basic implications of his model: more educated individuals and those employed in high-skilled occupations have a greater propensity to exit equal-sharing kibbutzim. Consistently with that model's prediction, wealth and ideological commitment is associated with greater

[8] Transferable membership rights may mitigate inefficient redistribution in WMFs. But membership markets are rare in practice, and most WMFs operate under collective ownership (Kremer, 1997).

[9] Kibbutzim resemble WMFs in many respects; for instance, both are managed by democratic principles (one person, one vote) and assets are held as common property. Of course, there are also important institutional differences: only the kibbutzim are communes, in which members share both production and consumption activities while working and living in one place.

[10] In this regard there is an important difference worth noting. Under Kremer's (1997) model, members vote on a wage schedule *after* their abilities are known (ex post redistribution); in Abramitzky (2008), however, contracts are written from a Rawlsian "original position": kibbutzim members choose the sharing rule behind a "veil of ignorance"—that is, *before* knowing their type (ex ante redistribution).

redistribution within kibbutzim. Abramitzky (2009) also documents adverse selection effects: equal-sharing kibbutzim attract individuals of lower ability.

III. CONTEXT AND DATA

III.A Worker-managed firms in Uruguay

In Uruguay, WMFs are those firms that are legally registered as producer cooperatives (PCs) in which the employee-to-member does not exceed 20%. Worker-managed firms are allowed to hire temporary employees in response to seasonal demand changes, but they must still comply with the legislated maximum level of hired workers in order to receive certain tax advantages—in particular, the exemption from paying the employer payroll tax to social security. The law also requires a minimum of six members to register a new cooperative firm.

Although their key organizational features are predetermined by law, WMFs have discretion over a broad range of associational rules. With respect to governance structure, WMFs must have a general workers' assembly that selects a council to supervise the daily operations (the council, in turn, usually selects the managers). Each member has only one vote, regardless of her capital contribution to the firm. Physical assets of WMFs can be owned by their members either collectively or individually. Under collective ownership, members do not own tradable shares but enjoy the right to usufruct as long as they work in the firm. Under individual ownership, members own capital shares that vary with the firm's value. Most Uruguayan WMFs operate under a collective ownership regime. As in other countries, membership markets are extremely rare in Uruguay.[11] A recent survey indicates that less than 10% of Uruguayan WMFs are owned by their workforce through individual shares (Alves et al., 2012). The activities of WMFs are financed via bank loans and/or retained earnings.[12] Previous studies have shown that Uruguayan WMFs exhibit a different adjustment process of wage and employment levels compared with conventional firms. The

[11] There is evidence that existing membership markets operate imperfectly, since share prices seem to be systematically undervalued (Craig and Pencavel, 1992). The role of membership markets has been extensively discussed in the literature (Sertel, 1982; Dow, 1986; Fehr, 1993).

[12] It is worth noting that capital markets play a minimal role in the financing and capitalization of conventional firms in Uruguay.

employment responses to idiosyncratic and macroeconomic shocks seem to be less elastic in WMFs than in conventional firms (Burdín and Dean, 2009; 2012).

III.B Worker-level panel data

To test whether redistribution actually takes place within WMFs, I use a random sample of Uruguayan workers who were registered in social security at least one month during the period from January 1997 to April 2010. The data were provided by Banco de Prevision Social, the agency in charge of social security affairs in Uruguay. Employers are obliged to deliver monthly information on their employees to the agency, which uses that information to calculate pension and social benefits.

The structure of the data is an unbalanced panel of workers extending from January 1997 to April 2010. The data contains information on daily wages, personal characteristics of the worker (gender, age, tenure), and attributes of the firm in which she works (firm size, industry). Each worker-month observation is tagged with a firm identification number so that job changes (and any other work history discontinuity) can be observed.[13] Most importantly, the data identifies the legal form of the firm for each worker's employment spell. Thus, workers employed by WMFs are identified as those working in a firm registered as a PC.[14] I restrict the sample to workers employed by nonagricultural private firms; public and rural workers are also excluded. Finally, I trim the data by excluding observations with daily wages corresponding to the top and bottom 1% of the wage distribution.

The descriptive statistics are presented in Appendix Table A.I. The resulting sample includes, on average, about 40,000 workers in each month. Those employed in WMFs amount to only some 3% of all workers. Average wages are higher in worker-managed than in conventional firms. However, the composition of the two groups is different: workers employed by WMFs are older than those employed by CFs, and in the latter case the average firm size is smaller. Proportionately fewer women are employed by WMFs than by

[13] The data contain information on both blue- and white-collar workers, including managers, but do not enable identification of workers' occupations. Survey evidence indicates that WMFs employ significantly fewer managers and supervisors irrespective of the sector in which they operate (Alves et al., 2012).
[14] In this data set, it is not possible to exclude PCs for which the employee-to member ratio is greater than 20%.

CFs, although female participation in the former has increased over the period. That change is driven, in part, by the change in the industry composition of Uruguayan WMFs: highly concentrated in the transport sector as recently as 1997, they have expanded into services and other sectors during the last decade.

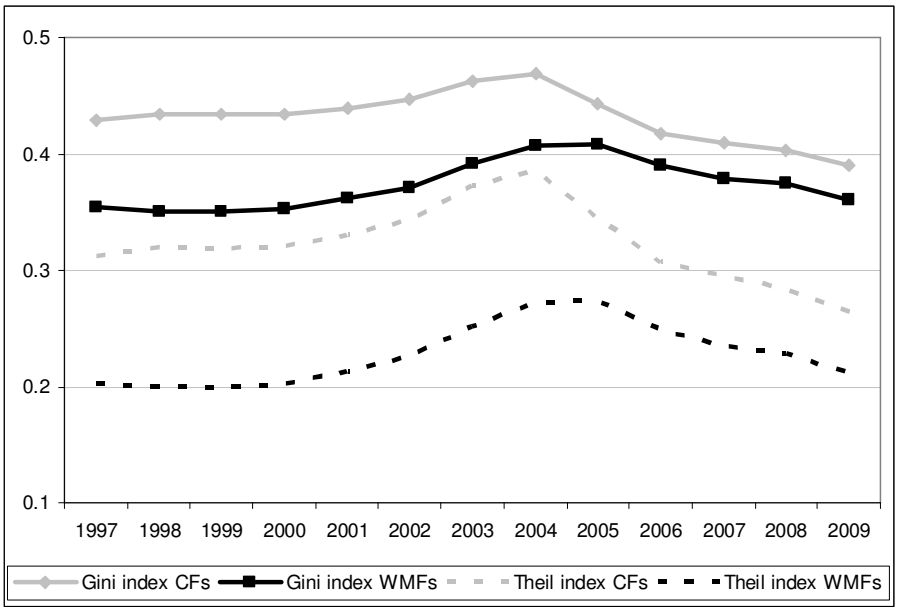
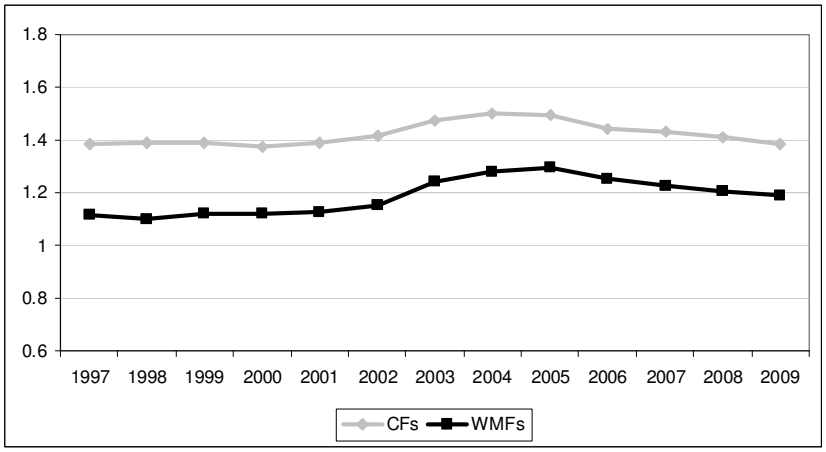


FIGURE I

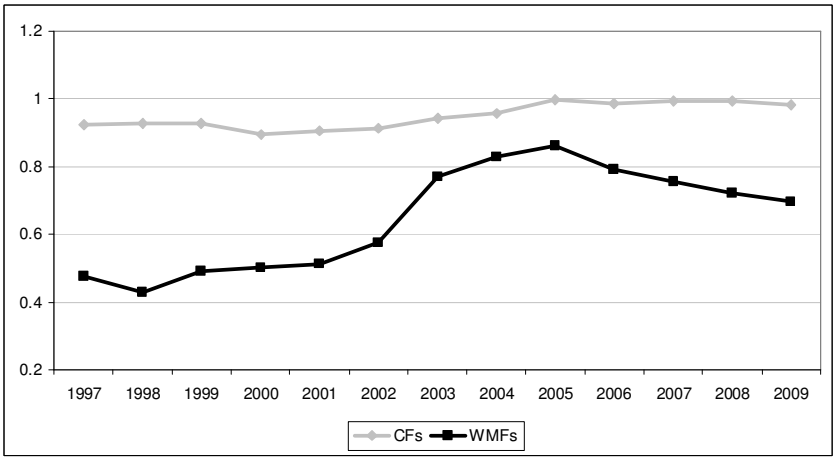
Wage Inequality in WMFs and CFs (daily wages), 1997–2009

To give a preliminary picture of the extent of redistribution within WMFs, I compute two standard inequality measures for workers employed by WMFs versus CFs.[15] Figure I plots the evolution of the Gini and Theil indexes of daily wages among the workers employed in each type of firms. As expected, wage inequality is systematically lower in WMFs. The Gini index is, on average, 9.3 percentage points (p.p) lower for workers employed by worker-managed than by conventional firms. It is interesting that this difference is even greater (14 p.p.) when wage inequality is measured by the Theil index, which is more sensitive to differences at the top of the wage distribution (Cowell, 2000).

[15] In each year, only workers between the ages of 20 and 55 are considered.



Panel A



Panel B

FIGURE II

Mean-to-Median Ratio and Wage Skewness in WMFs and CFs, 1997–2009

Notes: Panel A reports the mean-to-median ratio of daily wages. Panel B reports the Pearson's coefficient of wage skewness, computed as [3 × (mean − median)] ÷ (standard deviation).

Figure II provides further information that characterizes the wage distribution in WMFs and CFs. Worker-managed firms seem to reduce not only pay dispersion but also pay skewness, thus improving the median worker's compensation relative to the mean. Both the mean-to median wage ratio (Panel A) and the coefficient of wage skewness (Panel B) are systematically lower among workers employed by WMFs versus CFs. This is precisely the pattern one would expect from a WMF median voter model (Kremer, 1997).

III.C Matched organization–worker panel data

To investigate whether WMFs suffer from brain drain and whether this problem is related with the extent of internal redistribution, I exploit a matched employer–employee monthly panel data set. The data covers the entire population of Uruguayan firms registered as producer cooperatives and all their workers (members and nonmembers) during the period from January 1997 to April 2010. This data set, too, was provided by the Banco de Previsión Social and is based on the individual work histories used to calculate social benefits. Previous studies have pointed out that not all Uruguayan firms legally registered as PCs should be considered as WMFs (Burdin and Dean, 2009). Specifically, many PCs rely extensively on hired labor to carry out productive activities, which implies that—as in conventional firms—most of the workforce has no control over firm decisions. I therefore distinguish WMFs from the total population of PCs by using information of the employee-to-member ratio. I define WMFs as those PCs in which this ratio is lower than 20% at the time of entry. As mentioned previously, this is the maximum percentage allowed by the legal framework regulating the operation of WMFs in Uruguay. Estimates are performed using the subsample of WMFs just described.[16]

The main advantage of the data is that it is possible, for each WMF, to match the information on all its workers in each month with a unique identification number. Hence the structure of the data is that of a linked employer–employee panel data set. Firm-level information includes firm size (measured as total employment) and industry (5-digit SIC code). Worker-level information includes age, gender, job tenure, gross monthly wages, and number of days worked. Gross monthly wages are deflated by the Consumer Price Index and divided by the number of days worked in order to obtain the real daily wage for each worker. I also exclude workers whose daily wages are outside the 1%–99% range.

This ability to link firm and workers' information allows me to calculate different measures of workforce composition by firm (e.g., fraction of female workers, average age, age dispersion). Key to this study is that I can observe the entire wage distribution at any time and compute intrafirm pay dispersion indicators. The data enable me to observe each individual employment spell within WMFs and to locate workers' position in the firm's

[16] Results remain unchanged when the whole sample of workers employed in PCs is considered.

wage distribution. Among those workers who exit from WMFs during the period, I can also distinguish between voluntary quits and separation for other reasons (such as layoff, retirement, or death). Descriptive statistics on workers and firms are reported in Appendix Table A.II and Table A.III, respectively. The resulting sample includes, on average, roughly 10,500 workers and 270 producer cooperatives in each month. Information on the subsample of WMFs is also presented. It is worth noting that average wages in the individual-based data (Table A.II) are always higher than the average firm wage (Table A.III). This difference simply reflects the fact that larger PCs, which account for more workers, have higher average wages than smaller PCs; that is, the (unweighted) average firm wage is disproportionately influenced by small, low-wage PCs.

IV. RESULTS

IV.A Worker-managed firms redistribute in favor of low-wage workers

Section IV gives prima facie evidence that inequality is lower among workers employed by WMFs than among those employed by CFs. Of course, that naïve comparison may be affected by the different workforce and sectoral composition of each firm type. To provide more systematic evidence on redistributive policies in WMFs, I use the worker-level panel described in Section III.B and proceed as follows. First, in order to determine the sign and magnitude of the wage differential between workers employed in worker-managed and conventional firms, I estimate a standard Mincerian equation as follows:

$$\ln w_{ijt} = x_{ijt}\alpha + z_{jt}\beta + C_{ijt}\delta + \lambda_t + \eta_i + u_{ijt}\,, \qquad (1)$$

where ln *w* denotes the logarithm of real daily wages, the *x* are observed characteristics (gender, age, and tenure as well as quadratics in age and tenure) of the individual worker, the *z* are observed features (size, industry) of the enterprise *j* by which the individual is employed, and *C* is a dummy indicator variable that is set to 1 when worker *i* is employed by a WMF (and set to 0 otherwise); the λ_t are year fixed effects.[17] Unobserved factors affecting wages are represented by the terms *u* and η, where the latter denotes unobserved

[17] One drawback to using social security data is the lack of information on workers' education level.

factors that vary across individuals but are fixed for a given individual over time. The wage differential is captured by the coefficient δ.[18]

I estimate equation (1) via pooled ordinary least-squares (OLS) and fixed-effect (FE) regressions. The latter strategy is feasible because there is mobility of workers between WMFs and conventional firms. Under the assumption that selection into the WMF status is based on unobserved but time-invariant individual characteristics, fixed-effect regressions yield an unbiased estimate of the wage gap. The fraction of workers who switch between WMFs and CFs is roughly 4%.[19] It is well known (e.g., from the literature on unions) that FE estimates of a relatively persistent status—as when there are only a small number of switchers—are more susceptible to attenuation bias due to measurement errors (Freeman, 1984; Card, 1996). However, measurement errors are of less concern in this study because the estimates rely on administrative data that are extremely unlikely to reflect either misreporting or miscoding. The WMF status is measured by the legal form of the firm and hence is not likely to be misreported. Moreover, any change in a worker's WMF status corresponds to a change in the identification number of the firm employing that worker, which virtually eliminates miscoding.

Estimates are reported in Table I. Column 1 reports the results of the pooled OLS estimate, according to which a worker employed by a WMF earns 5.5% more than one employed by a CF; this difference is highly significant. However, an OLS estimate of equation (1) may be biased if C and η are correlated—that is, if unobservable factors affecting the choice between working for a WMF or a CF are correlated with the determinants of earnings (Pencavel, Pistaferri, and Schivardi, 2006). Column 2 reports the results from a fixed-effect regression that yields consistent estimates for δ under arbitrary correlation between C and η. The wage gap is still positive (2.7%) and significant at the 10% level.[20]

[18] There is no clear theoretical prediction regarding the sign of the wage differential. Pencavel, Pistaferri, and Schivardi (2006) adopt a similar empirical approach and find that, in Italy, being employed by a WMF is associated with a negative wage gap.

[19] Roughly, 70% of workers' transitions between WMFs and CFs correspond to CF-to-WMF switches. In Appendix Table A.IV, I report descriptive statistics on workers, distinguishing transition states ("Stayed in CF", "Stayed in WMF", "CF-to-CF mover", "WMF-to-CF" and "CF-to-WMF mover" .

[20] The Hausman test leads to a strong rejection of the null hypothesis that random effects yield consistent estimates ($p = 0.000$).

TABLE I
WAGE GAP BETWEEN WORKERS EMPLOYED IN WMFs AND CFs

	OLS	FE	FE	FE	FE
	(1)	(2)	(3)	(4)	(5)
Coop	0.055**	0.027*	0.028*	0.092**	0.091**
	(0.011)	(0.015)	(0.016)	(0.038)	(0.038)
Female	−0.230***				
	(0.005)				
Age	0.060***	0.210***	0.212***	0.212***	0.211***
	(0.001)	(0.002)	(0.002)	(0.002)	(0.002)
Age squared	−0.001***	−0.001***	−0.001***	−0.001***	−0.001***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Tenure	0.047***	0.032***	0.030***	0.030***	0.030***
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
Tenure squared	−0.001***	−0.001***	−0.001***	−0.001***	−0.001***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Firm size (in logs)	0.153***	0.122***	0.099***	0.100***	0.099***
	(0.001)	(0.001)	(0.002)	(0.002)	(0.002)
Year fixed effects	Yes	Yes	Yes	Yes	Yes
Industry fixed effects	Yes	Yes	Yes	Yes	Yes
Observations	5,264,811	5,264,811	3,533,031	3,445,097	3,445,097

Notes: The dependent variable is the log of daily wage. Column 1 reports pooled OLS estimates; Columns 2–5 report panel data fixed-effect estimates. The estimates reported in Columns 3–5 exclude workers employed in firms with fewer than six workers. Estimates in Columns 4 and 5 compare employees in CFs with members in WMFs (i.e., nonmembers are excluded). All estimates include a set of thirteen year dummies and six industry dummies. The estimates in Column 5 also include sectoral-specific year fixed effects. Standard errors (reported in parentheses) are clustered at the individual level. * significant at 10%; ** significant at 5%; *** significant at 1%

In Uruguay, there is a legal restriction on the minimum size of WMFs. More precisely, WMFs cannot be formed with fewer than six members—which helps explain why the average firm size is greater in worker-managed than in conventional firms (see Table A.I). For this reason I perform an additional FE estimate that excludes workers employed in micro-enterprises (i.e., firms employing fewer than six workers). The results, which are reported in Column 3 of Table I, remain unchanged. The estimates so far have compared all workers employed in WMFs (members and nonmembers) with those employed in CFs. Results are qualitatively similar when considering only WMF members. The wage gap is slightly higher (9%) and highly significant (see Column 4 of Table I). This is plausible given that WMF members' compensation includes distributed dividends. Finally, to account for heterogeneous time effects across sectors, Column 5 reports estimates that include sectoral-specific year fixed effects. Results are robust also to this modification.[21]

[21] I replicate the estimates reported in Column 4 when including both month and year fixed effects. Alternatively, I try adding a linear time trend. I also perform estimates using the log of hourly wages (instead

Having documented a positive wage premium associated with being employed in a WMF, I then ask whether this wage gap varies across the wage distribution. If WMFs actually implement redistributive policies, then we should expect the magnitude of the wage differential to be greater at the bottom of the wage distribution. In other words, the gain experienced by a worker who moves from a conventional firm to a worker-managed firm should be greater for low-wage than for high-wage workers. To perform this analysis, I use quantile regression to estimate the wag gap associated with being employed in a WMF at each quantile $\theta \in [0,1]$ of the distribution of the log of daily wages of worker *i* in firm *j* during month *t*:

$$Quant_{\theta}(w_{ijt} \mid \cdot) = Coop_{ijt}\gamma_{\theta} + X_{it}\beta_{\theta} + Z_{jt}\delta_{\theta}, \qquad (2)$$

where $Quant_{\theta}(w_{ijt} \mid \cdot)$ refers to the conditional quantile of the log of daily wages, X_{it} captures personal characteristics (gender, age, age squared, tenure, tenure squared), and Z_{jt} stands for firm attributes (firm size, industry); $Coop_{ijt}$ is a dummy variable set equal to 1 only if individual *i* is employed by a WMF. I perform separate quantile regression estimates by year, pooling monthly workers' records in each year.

Table II reports the results of Pooled quantile regression estimates for the 0.2, 0.4, 0.6, and 0.8 quantiles.[22] As expected, the wage premium associated with being employed in a WMF declines along the wage distribution and becomes negative at the top. The wage premium for the 0.2 quantile is 18% as compared with a wage penalty of 4% for the 0.8 quantile.[23] In Appendix Table A.V, I report the results of quantile regressions for each year separately, pooling monthly workers' records in each year. Interquantile differences appear to be quite stable over the period. For example: in 1997, the wage premium for the 0.2 quantile was 18% as compared with a wage penalty of 3% for the 0.8 quantile; in 2009, the respective

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of the daily wage) as the dependent variable. Results are robust to all these modifications. Estimates using daily wages are preferred because information on working hours is missing for nearly a fifth of the sample.

[22] In Appendix Table A.V, I report the results of quantile regressions for each year separately, pooling monthly workers' records in each year

[23] The null hypothesis of no interquantile differences is rejected in all cases.

figures were 16% and 4%. Compensation policies within Uruguayan WMFs seem to strongly favor workers at the bottom of the distribution.[24]

TABLE II
WAGE GAP ACROSS THE WAGE DISTRIBUTION.
QUANTILE REGRESSION ESTIMATES

	Pooled QR				Fixed effects QR			
	q20	q40	q60	q80	q20	q40	q60	q80
Coop	0.184***	0.111***	0.044***	-0.043***	0.041***	0.036***	0.026***	0.023***
	(0.002)	(0.002)	(0.002)	(0.002)	(0.001)	(0.001)	(0.001)	(0.001)
Age	-0.187***	-0.211***	-0.245***	-0.277***	0.014***	0.004***	-0.004***	-0.0146***
	(0.001)	(0.001)	(0.001)	(0.001)	(0.000)	(0.000)	(0.000)	(0.001)
Age squared	0.041***	0.050***	0.060***	0.078***	0.199***	0.208***	0.214***	0.218***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Female	-0.001***	-0.001***	-0.001***	-0.001***	-0.001***	-0.001***	-0.001***	-0.001***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Tenure	0.051***	0.049***	0.0475***	0.046***	0.038***	0.029***	0.0258***	0.0215***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Tenure squared	-0.001***	-0.001***	-0.001***	-0.001***	-0.001***	-0.001***	-0.001***	-0.001***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Firm size	0.162***	0.150***	0.145***	0.147***	0.120***	0.120***	0.121***	0.123***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Test of interquantile Differences								
20th = 40th	[0.000]				[0.000]			
20th = 80th	[0.000]				[0.000]			
40th = 80th		[0.000]				[0.000]		
Observations	5,264,811	5,264,811	5,264,811	5,264,811	5,264,811	5,264,811	5,264,811	5,264,811

Notes: The dependent variable is the log of daily wages. The *Coop* dummy variable is set equal to 1 only for workers employed in a PC. Firm size is measured as the log of total employment in each firm. All estimates include six industry dummies. Bootstrapped standard errors (reported in parentheses) are based on 200 replications. * significant at 10%; ** significant at 5%; *** significant at 1%

Pooled QR estimates compare individuals with different unobserved ability. For such reason, I implement the approach recently proposed by Canay (2011) to control for unobserved heterogeneity in a quantile setting.25 Table II also reports the results of the resulting Fixed Effect QR estimates. Consistent with Pooled estimates, the wage premium associated with being employed in a WMF is significantly declining in wages, reinforcing the idea that WMFs actually redistribute in favor of low-wage workers. Interestingly, the

[24] Studies based on survey data comparing WMFs and CFs do not find significant wage differences for low-wage occupations (unskilled workers) but confirm a significant wage penalty for managerial occupations (Estrin and Holmes, 1991; Alves et al., 2012). These studies rely on mean comparisons of firm-level data on wages grouped by occupation.

[25] For a detailed description of the procedure see Canay (2011).

wage gap seems to be partially driven by selection, particularly at the bottom of the wage distribution. The comparison between Pooled QR and FE-QR estimates indicates that the sign of the Pooled QR bias varies across the wage distribution, suggesting that the pattern of sorting into WMFs is heterogeneous. Results are partly in line with the prediction of a standard Roy model of selection (Roy, 1951; Borjas, 1987). As the compensation structure in WMFs is more compressed than in the capitalist sector and, hence, returns to ability are lower, one would expect that low-ability workers self-select into WMFs. This seems to be true at the top but not at the bottom of the wage distribution. Indeed, anecdotal and survey evidence suggests that WMFs rely on different recruitment channels than CFs (e.g. recommendations from incumbent members, trial periods) in order to ensure ideological commitment and screen out low-ability applicants, mitigating adverse selection effects (Benham and Keefer, 1991; Burdin, 2013b). In the next section, I investigate in greater detail the selection in exit from WMFs.

IV.B High-ability WMF members are more likely to quit

In this section I test whether redistributive policies implemented by WMFs affect workers' flows. Specifically, I analyze whether the hazard of voluntary separation is greater for high-ability workers. To perform this analysis, I use the linked organization–worker panel described in Section III.C. Because the study focuses on voluntary quits, I restrict the sample in several ways. First, I exclude workers older than 55 because they are probably considering retirement. Second, I do not consider separations caused by firm closures. Third, separations due to other reasons (e.g., layoffs, death) are treated as censored.[26] Finally, I drop left-censored spells—that is, individuals who were already working in a given firm at the beginning of the study period (January 1997). The problem of right-censored observations is handled by using duration analysis techniques.

In order to identify high-ability workers, I divide the workforce of each firm (at any moment in time) into two groups: those with wages above and those with wages below the firm's median wage. The intuition behind this procedure is to use the within-firm wage variation to rank workers according to their ability types. Controlling for other observables

[26] Voluntary quits constitute 72% of total worker separations. As expected, the fraction of voluntary quits increases (to 82%) when the analysis is restricted to members.

characteristics of the worker and the firm, I assume that the position of the worker in the internal wage scale is a reasonable proxy for her position in the ability distribution. This approach has been adopted in the literature on assortative matching between workers and firms, but it requires one to assume that workers' payoffs are increasing in their own types (Bartolucci and Devicienti, 2012).

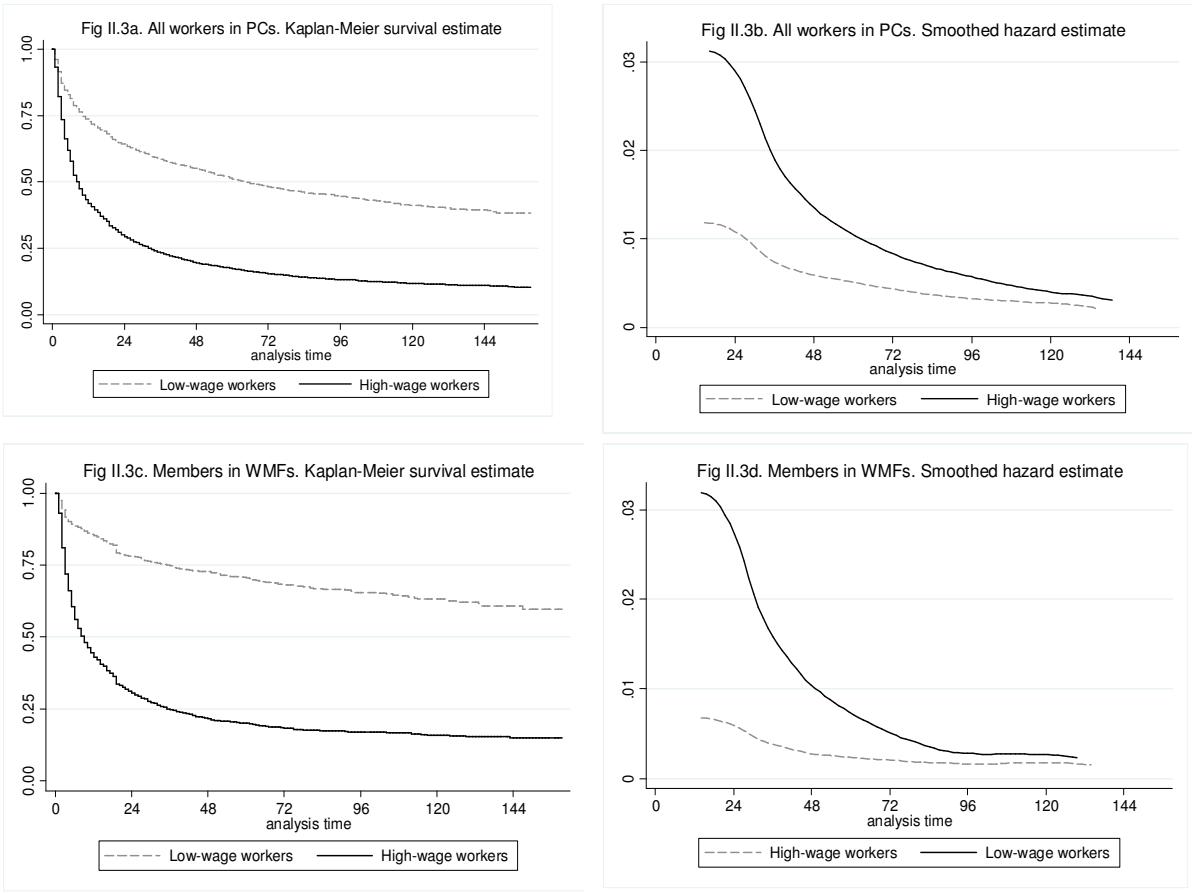

FIGURE III

Worker's Position in the Within-firm Wage Distribution and Job Duration

Notes: The *High-wage worker* indicator variable is set equal to 1 only for a worker whose daily wage is above the median daily wage in the firm that employs her. Figures IIIa and IIIb consider the full sample—that is, all workers (members and nonmembers) employed by PCs; in Figures IIIc and IIId the estimates are restricted to members of WMFs.

Figure III plots nonparametric estimates of the survival function and the hazard function for job separations while distinguishing between high- and low-wage workers. These functions are calculated for both the whole sample of workers employed in PCs (Figures IIIa and

IIIb) and the subsample of WMF members (Figures IIIc and IIId).[27] The hazard of job separation is systematically higher for high-wage workers in both cases. The log-rank test clearly rejects the null hypothesis that the survivor functions of the two types of workers are equal ($\chi_{(1)} = 2410$). In order to analyze the determinants of employment duration in WMFs (i.e., the time elapsed between workers' enrollment and voluntary separation), I estimate a proportional hazard model (Cox, 1972):

$$h_{ij}(t) = h_j(t)\exp(HighW_{it}\beta_1 + X_{it}\beta_2 + Z_{jt}\beta_3), \quad (3)$$

where $h_j(t)$ is the baseline hazard for firm *j* and where *t* is the number of months that individual *i* has been employed at firm *j*; the dummy variable $HighW_{it}$ is set equal to 1 for workers whose daily wage is above the firm's median daily wage, *X* is a vector of personal characteristics (gender, age, age squared), and *Z* is a vector of firm characteristics (firm size, proportion of female workers, workforce average age and its dispersion). The effect of a unit change in any covariate is to produce a constant proportional change in the hazard rate. The coefficient of interest is β_1.[28] To rule out potential unobserved firm-level confounding factors, I estimate stratified Cox models in which each firm has its own flexible baseline hazard function. This approach allows one to control for all time-invariant firm-level characteristics (Giuliano, Levine, and Leonard, 2011). Cox model estimates stratified by firm eliminate unobserved heterogeneity across firms but not across individuals within a firm. I account for unobserved individual-level heterogeneity by also estimating a parametric model in which each individual's duration depends on a random

[27] The Kaplan–Meier survivor function is defined as $\hat{S}(t_j) = \prod_{j|t_j<t}(1-d_j/n_j)$, where d_j is the number of failures occurring at time t_j and where n_j is the number at risk at time t_j (before any failures occur). The hazard function is calculated as a weighted kernel density using the estimated hazard contributions: $\Delta\hat{H}(t_j) = \hat{H}(t_j) - \hat{H}(t_{j-1})$, where t_j is the current failure time and $\hat{H}(t_j)$ is the estimated cumulative hazard. The Nelson–Aalen estimator of $\hat{H}(t_j)$ is defined as $\hat{H}(t_j) = \sum_{j|t_j\le t}(d_j/n_j)$; this is the sum of expected failures at each observed time. For further details on nonparametric survival analysis, see Jenkins (2005) and Cleves et al. (2008).

[28] The Breslow method is used for handling ties. I check the empirical plausibility of the proportional hazard (PH) assumption by means of graphical methods (Jenkins, 2005; Cleves et al., 2008). This assumption seems to be satisfied by the data; see Appendix Figure A.I. I also perform the test based on the Schoenfeld residuals for the variable *HighW* and do not reject the PH assumption ($p = 0.218$). The PH assumption is not rejected (at 5%) when the global test of the model is considered ($p = 0.0774$).

effect ("frailty") and the baseline hazard is assumed to have a log-normal distribution (Jenkins, 2005).[29]

Table II reports the results. All estimates are restricted to the subsample of members of WMFs. Column 1 reports the results of estimating equation (3) while controlling only for personal characteristics. In Column 2 the estimates control also for firm-level characteristics and include cohort fixed effects to account for common shocks (at the time of entry) that may affect subsequent job duration. Column 3 reports estimates of the parametric frailty model. Results are qualitatively similar across specifications. The hazard of job separation is systematically greater for high-ability workers. The results reported in Column 2 indicate that high-wage members are 3.7 times more likely than are low-wage members to exit.[30] Estimates reported in Column 3, which account for individual unobserved heterogeneity, indicate that the time ratio associated with being a high-ability worker is 0.23; this means that the status of high-ability worker reduces employment duration (survival time) within a WMF by 77%, or roughly 20 months.[31]

That the high-ability are more likely to exit provides further support for the idea that pay compression in WMFs is a deliberate policy. As Lazear and Shaw (2009) point out, there would be no reason for top workers to leave disproportionately (nor for bottom workers to stay disproportionately) if all workers were paid their competitive wage.[32]

[29] The log-normal distribution is consistent with the nonmonotonic pattern of duration dependence of the hazard observed in Figure III. Unlike the Cox model, the log-normal model does not rely on the PH assumption.

[30] By expressing the model in terms of the log of the hazard ratios, this effect is computed as exp(1.32).

[31] This effect is computed as $[1 - \exp(-1.484)] \times 100 = 77..32$. The mean employment duration for the subsample of WMF members is 27 months; thus, $(27 \times 0.77)/12 = 1.73$. Observe that, in Column 3, the covariate effects must be interpreted in terms of survival time ("accelerated failure time" metric) and not in terms of the hazard as in Cox model estimates ("proportional hazard" metric).

[32] I perform additional robustness checks as well. First, I estimate the Cox model considering all workers (members and nonmembers) in WMFs. Second, I consider the whole sample of workers employed in all PCs. Third, I exclude employment spells with time gaps. Fourth, I redefine high-ability workers as those whose wage is above the 80th percentile of the within-firm wage distribution. Finally, I estimate the Cox model defining covariates and the worker's position in the within-firm wage distribution at the time of entry. None of the described modifications alters the basic results. Estimates for these alternative regressions are available from the author upon request.

TABLE II

WORKER'S POSITION IN THE WITHIN-FIRM WAGE DISTRIBUTION AND HAZARD OF EXIT IN WMFs. RESULTS FROM DURATION MODELS ESTIMATES.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
(A) *HighW*	1.320***	1.307***	−1.484***	1.895***	2.398***	2.453***	1.375***
	(0.0529)	(0.0529)	(0.0614)	(0.123)	(0.190)	(0.254)	(0.0690)
(B) *HighW* × *Coef. of variation*				−1.610***			
				(0.254)			
Coef. of variation				1.606***			
				(0.258)			
(C) *HighW* × *Mean-to-median ratio*					−0.995***		
					(0.149)		
Mean-to-median ratio					1.855***		
					(0.164)		
(D) *HighW* × *Sigma*						−4.184***	
						(0.809)	
Sigma						−4.064*	
						(2.230)	
(E) *HighW* × *Founding member*							−0.428***
							(0.119)
Founding member							−0.251**
							(0.119)
Hazard ratio/ Time ratio							
(A)	3.743***	3.695***	0.227***				3.955***
	(0.198)	(0.196)	(0.014)				(0.273)
Post-estimation: (A) + σ*(B)				2.482***			
				(0.177)			
Post-estimation: (A) − σ*(B)				5.054***			
				(0.434)			
Post-estimation: (A) + σ*(C)					3.274***		
					(0.177)		
Post-estimation: (A) − σ*(C)					4.129***		
					(0.268)		
Post-estimation: (A) + σ*(D)						1.491***	
						(0.278)	
Post-estimation: (A) − σ*(D)						6.386***	
						(0.975)	
Post-estimation: (A) + (E)							2.579***
							(0.254)
Worker-level controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Firm-level controls	No	Yes	Yes	Yes	Yes	Yes	Yes
Cohort fixed effects	No	Yes	Yes	Yes	Yes	Yes	Yes
Observations	183,523	183,514	183,514	163,151	163,151	112,235	96,722

Notes: Cox proportional hazard models stratified by firm—except for Column 3, which reports estimates from a shared "frailty" model in which the baseline hazard assumes a log-normal distribution. The HighW dummy variable is set equal to 1 for those workers whose daily wage is above the firm's median daily wage (and to 0 for other workers). All estimates control for worker-level characteristics (gender, age, age squared) and are restricted to WMF members. Estimates presented in Columns 2–6 also control for firm-level characteristics (firm size, average age of the workforce and its dispersion, fraction of female) and cohort fixed effects. The estimates presented in Column 3 include industry fixed effects; in Columns 4 and 5, the estimates are restricted to WMFs employing at least ten workers. In Column 7, estimates are restricted to WMFs (formed after January 1997) for which founding members can be identified. Robust standard errors (reported in parentheses) are adjusted for clustering at the individual level. * significant at 10%; ** significant at 5%; *** significant at 1%

IV.C High-ability workers are less likely to quit when redistribution within WMFs is less intense; founding members are also less prone to exit

One can certainly argue that high-ability workers are more mobile in any organizational setting and so not simply because of redistributive policies implemented by WMFs. Yet because matched organization–worker data is available only for WMFs, I am unfortunately not able to assess whether the brain drain is greater in WMFs than in CFs.

Despite this limitation, it is still possible to analyze how CF-to-CF transitions correlate with workers' pre-exit wage using the panel of workers described in section III.B. In Appendix Table IV, I report estimates from binary outcome models in which the dependent variable is a dummy that takes a value of one if the worker switched from a CF to another CF. I control for gender, age, tenure, quadratics in age and tenure and firm size in order to compare workers with similar observables characteristics. Results from both probit and logit regressions indicate that the probability of switching from a CF to another CF is negatively correlated with the worker's pre-exit wage level. Results are qualitatively similar when instead of the log of pre-exit daily wage I use wage quintiles to indicate the position of the worker in her sector-specific wage distribution. The probability of being a CF-to-CF switcher is lower for workers located at the top of their sector-specific wage distribution. At least from this exploratory analysis, there is no evidence that high-ability workers are more prone to leave their firms when the analysis is restricted to workers employed in the capitalist sector.

To assess whether the degree of equality within WMFs affects the outflow of high-ability members, I am able to exploit the observed within-firm variation in intrafirm wage dispersion among WMFs. As already mentioned, this procedure allows me to estimate models that control for unobserved differences across firms. The expectation is that a less compressed wage structure mitigates brain drain. To test this hypothesis, I estimate equation (3) while including a measure of intrafirm inequality and its interaction with the variable identifying high-wage members. Because measures of intrafirm inequality are not meaningful for small firms, I restrict the sample to WMFs employing at least ten workers.

In order to characterize the wage distribution within each WMF, I consider two measures: the coefficient of variation and the mean-to-median ratio of wages within the firm.[33] I expect the coefficient for the interaction term to be negative. If brain drain is driven by egalitarian wage policies implemented by WMFs then, ceteris paribus, high-wage workers should be less likely to exit WMFs in which redistribution is less pronounced.

The results reported in Columns 4 and 5 of Table II support this hypothesis. The interaction term is negative and statistically significant in both specifications, and the magnitude of the effect is sizable. I report the post-estimation of the hazard ratio (using a linear combination of parameter estimates) when the within-firm coefficient of variation in daily wages is one standard deviation above or below the mean. According to the values in Column 4 of the table, the hazard ratio of high-ability members is twice as high in a WMF for which the within-firm coefficient of variation in wages is one standard deviation (0.221) below the sample mean (0.392)

Results are qualitatively similar in estimates that include the mean-to-median firm wage ratio (see Column 5 of Table II). The hazard ratio of high-ability members is 1.26 times higher in a WMF for which the mean-to-median wage ratio is one standard deviation (0.117) below the sample mean (1.101). It should be emphasized that the mean-to-median wage ratio has a direct interpretation in terms of a WMF median voter model (Kremer, 1997). Higher values of the mean-to-median ratio indicate that the median voter commits *not* to engage in redistribution while taking into account participation constraints of the most productive members. A consistent feature of the findings reported here is that the brain drain effect is mitigated in those WMFs whose median member is less prone to leverage her pivotal position in the organizational political process to redistribute away from high-ability members.

Neither the coefficient for variation nor the mean-to median wage ratio take into account the observed heterogeneity among workers within firms. This is an important limitation when one considers that most theoretical predictions about the effect of egalitarian wage

[33] I compute the average of these variables over each individual employment spell. Hence, whereas the averages vary both between and within firms, they vary only between (not within) individuals. In this way I can estimate the Cox model stratified by firm.

policies in WMFs are derived from models in which members have different abilities but are observationally equivalent (Kremer, 1997). To provide further evidence on the interplay between intrafirm pay dispersion and brain drain in WMFs, I repeat the previous empirical exercise while using a conditional measure of intrafirm inequality. In order to compute a conditional measure of inequality at the firm level, I run a standard wage equation separately for each firm in each month, where the dependent variable is the log of real daily wage of individual *i* employed by firm *j* at time *t*. As explanatory variables, I consider the following worker characteristics: gender, age, job tenure, quadratics in age and tenure, and a dummy variable distinguishing between members and nonmembers. The standard errors of these wage regressions are taken as a conditional measure of intrafirm wage inequality, $Sigma_{jt}$; this variable captures the wage inequality (among workers employed by firm *j* at time *t*) that remains after controlling for observable personal characteristics.[34]

I next estimate equation (3) while including among the regressors both $Sigma_{jt}$ and its interaction with the variable identifying high-wage members. Results are reported in Column 6 of Table II. The hazard ratio of high-ability members is 4.3 times higher in a WMF whose intrafirm residual wage dispersion is one standard deviation below the mean. As expected, the more the wage structure of a WMF recognizes differences in workers' ability, the less the incentive of members in the upper tail of the internal wage distribution to exit the firm and migrate to the capitalist sector. The higher hazard rate of exit for top members in more egalitarian WMFs supports the argument that wage compression is a deliberate policy in this type of firm.

Finally, I analyze whether the hazard ratio of high-ability members varies with their status in the organization. Previous evidence from Israeli kibbutzim indicates a positive association between the degree of equality and the degree of members' ideology (Abramitzky, 2008). Ideology seems to play the role of relaxing the participation constraint by increasing the nonpecuniary value of staying in the kibbutz. It is unfortunate that I have no direct measure of a member's ideology. Nonetheless, it is possible to identify the founding members of WMFs formed after January 1997. It is reasonable to assume that the

[34] This procedure was suggested by Lazear (1989) and originally implemented by Winter-Ebmer and Zweimüller (1999).

ideological commitment of first-generation members is stronger than that in subsequent members. I estimate equation (3) while including an indicator variable for founding member and its interaction with the variable identifying high-wage members. These results are reported in Column 6 of Table II. On average, founding members are less likely to quit WMFs. A finding of particular interest is that the hazard ratio of high-ability members is 1.4 times lower in the case of founding members. This result confirms the intuitive notion that a WMF's redistribution policies are less constrained by the threat of brain drain when members are intrinsically motivated to join the firm.[35]

IV.D High-ability members are less likely to quit when outside options are less attractive

Finally, I analyze whether the hazard of exit of high-ability members varies according to changes in labor market conditions in the capitalist sector. To characterize the external labor market, I use three-month lagged values of both the monthly urban unemployment rate ($Unemp_{t-3}$) and the ratio of the median daily wage paid in the capitalist sector—computed for the specific 2-digit sector of the WMF in which the individual is employed—to the member's daily wage ($Ratiow_{it-3}$).[36] I then estimate equation (3) while including these variables and their interaction with the variable identifying high-wage members within WMFs.

Results are reported in Table III. As expected, the more (less) attractive are the external labor market conditions, the higher (lower) is the hazard of exit in WMFs. More precisely, estimates reported in Column 1–3 indicate that both an increase in the unemployment rate and a reduction in the 2-digit, sector-specific reference wage (relative to the member's current wage) significantly reduces the hazard of exit from WMFs. Column 4 of the table reports estimates that include labor market conditions interacted with the variable $HighW_{it}$, which identifies high-wage members within WMFs. Both interaction terms have the

[35] First-generation members may also have greater sunk investments in their firms. Therefore, I cannot rule out that a founding member's lower hazard of exit is due to lock-in effects associated with the collective ownership of a WMF's physical assets. Indeed, Abramitzky (2008) finds that the degree of equality is higher in wealthy kibbutzim and that higher wealth reduces the brain drain in equal-sharing kibbutzim.

[36] The monthly urban unemployment rate is based on official statistics published by the Uruguayan National Statistical Institute (www.ine.gub.uy). The 2-digit sector median daily wage in the capitalist sector is computed using the data set described in Section IV.A. Results remain unchanged when the values of both variables are lagged by six months.

expected sign and are highly significant. It is worth noting that the sensitivity of quit decisions to external labor market conditions also varies according to the member's position in the intrafirm wage distribution. When outside options in the capitalist sector become more attractive, the exit hazard increases more for high-ability than for low-ability members.

TABLE III
LABOR MARKET CONDITIONS AND HAZARD OF EXIT IN WMFs.
RESULTS FROM DURATION MODELS ESTIMATES

	(1)	(2)	(3)	(4)
HighW	1.468***	1.531***	1.530***	1.709***
	(0.061)	(0.070)	(0.070)	(0.067)
Ratiow		0.207**	0.210***	0.095
		(0.083)	(0.082)	(0.066)
Unemp	−0.039***		−0.039***	−0.012
	(0.014)		(0.014)	(0.014)
HighW × *Unemp*				−0.089***
				(0.011)
HighW × *Ratiow*				0.256***
				(0.094)
Worker-level controls	Yes	Yes	Yes	Yes
Firm-level controls	Yes	Yes	Yes	Yes
Observations	163,949	159,628	159,628	158,917

Notes: Cox proportional hazard models stratified by firm. The *HighW* dummy variable is set equal to 1 for those workers whose daily wage is above the firm's median daily wage (and to 0 for other workers); *Ratiow* is the ratio of the member's daily wage to the median daily wage corresponding 2-digit sector of the WMF in which the individual is employed; and *Unemp* is the monthly urban unemployment rate. All estimates include *Ratiow* and *Unemp* (lagged three months) and are restricted to WMF members. In addition, all estimates control for worker-level characteristics (gender, age, age squared) and firm-level characteristics (firm size, average age of the workforce and its dispersion, fraction of female). Robust standard errors (reported in parentheses) are adjusted for clustering at the individual level. * significant at 10%; ** significant at 5%; *** significant at 1%

V. CONCLUSION AND DISCUSSION

In this paper I study the extent and effects of redistribution in WMFs. The analysis is based on a panel of Uruguayan workers and a linked employer–employee panel data set covering the country's entire population of WMFs and their workers. The analysis supports four basic findings. First, workplace democracy is associated with substantial redistribution among workers. There is only a small wage premium associated with being employed in a WMF, and this gap declines significantly with increasing wage. Second, WMFs suffer from brain drain: the separation hazard of high-ability members is more than three times higher than that of low-ability members. Third, in WMFs there is a relationship between the extent

of pay compression and the severity of brain drain: I find that high-ability workers are less likely to exit a WMF whose wage structure is less compressed. I also find that the status of founding member is generally associated with a lower hazard of exit and significantly reduces the hazard of high-ability members, suggesting that the presence of intrinsically motivated workers enables greater redistribution. Finally, I find that the quit behavior of high-ability members varies as a function of labor market conditions in the capitalist sector.

It is beyond the scope of this paper to analyze the relationship between pay compression and organizational performance in WMFs. The brain drain effects documented here suggest a plausible mechanism to account for a potential negative relationship between pay compression and performance. Another possible explanation, which is suggested by tournament theory, is that a compressed wage structure reduces the expected gains from internal promotions and hence does not provide enough incentive to increase workers' efforts (Lazear and Rosen, 1981). Nevertheless, panel data evidence on the relative efficiency of WMFs indicates that they perform as well as (or even bettter than) conventional firms in terms of productivity (Craig and Pencavel, 1995; Fakhfakh, Pérotin, and Gago, 2012; Pencavel, 2013). Burdin (2013a) also shows that Uruguayan WMFs are less likely to dissolve than are conventional firms. Note that experiments on team production in which selection plays no role—because random assignment guarantees that the allocation of subjects to organizational types is fully exogenous—also find positive performance effects associated with workplace democracy (Frohlich et al., 1998; Mellizo, Carpenter, and Matthews, 2011).

Those experimental and nonexperimental studies suggest that other beneficial effects associated with pay compression are at work in WMFs. First, a greater degree of equality may result in higher productivity through greater teamwork (Lazear and Shaw, 2007). Conversely, competition for promotions within firms may erode workplace cooperation and cohesiveness (Lazear, 1989; Levine, 1991). Third, pay dispersion may exacerbate rent-seeking behavior within firms: workers may withhold information to increase their influence and persuade managers, wasting time and organizational resources rather than producing (Milgrom and Roberts, 1990). Fourth, pay comparisons within organizations reduce the job satisfaction for workers at the bottom of the wage distribution (Card et al.,

2012). Finally, pay dispersion may reduce the effectiveness of peer pressure as a mechanism for overcoming free-rider problems in team production (Kandel and Lazear, 1992). For instance, pay differences may increase social distance between members and reduce the psychological costs (guilt, shame) incurred by those who deviate from group effort norms.

Because group sociology may influence team pay practices (Encinosa, Gaynor, and Rebitzer, 2007), it is possible for pay equality—and other policies that increase the degree of empathy among members—to facilitate mutual monitoring within WMFs. The costs of equality associated with brain drain and inferior management quality may be outweighed by other labor discipline benefits, such as higher motivation of shop-floor workers and lower supervision costs.[37] Further research is needed to investigate the efficiency-enhancing effects of pay compression in democratically controlled workplaces.

[37]Survey evidence indicates that supervision intensity is significantly lower in Uruguayan WMFs than in conventional firms. WMFs rely more on mutual monitoring among co-workers to ensure workplace discipline. Interestingly, egalitarian WMFs have lower supervision intensity and rely more on horizontal monitoring than non-egalitarian WMFs (Burdin, 2013b).

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Appendix

TABLE A.I
DESCRIPTIVE STATISTICS. PANEL OF WORKERS

	1997		2001		2005		2009	
	CF	WMF	CF	WMF	CF	WMF	CF	WMF
Number of workers	36,117	1,305	33,944	1,092	38,148	1,138	46,667	1,220
Fraction female	0.43	0.36	0.45	0.38	0.45	0.42	0.45	0.44
Age	36.34	41.21	37.32	42.66	37.59	43.51	38.07	43.09
	(12.63)	(10.57)	(12.27)	(10.75)	(12.16)	(10.80)	(12.18)	(11.13)
Tenure	5.26	9.12	5.80	10.33	5.39	10.81	4.87	10.15
	(6.67)	(8.12)	(6.82)	(8.64)	(6.97)	(9.28)	(6.62)	(9.76)
Monthly wage	13,829	25,138	13,118	22,632	10,779	17,880	13,376	21,210
	(13,260)	(17,546)	(13,398)	(15,693)	(11,181)	(15,551)	(12,428)	(16,028)
Daily wage	523.55	922.90	497.15	911.43	416.68	668.41	519.10	804.54
	(469.08)	(592.96)	(469.11)	(620.86)	(394.11)	(520.80)	(434.06)	(543.61)
Hourly wage	89.60	156.81	87.65	143.53	71.99	115.60	89.60	131.68
	(84.72)	(103.06)	(86.34)	(90.87)	(75.82)	(92.15)	(85.70)	(95.99)
Firm size	3.74	5.78	3.81	5.68	3.81	5.45	3.94	5.69
	(1.96)	(1.77)	(2.02)	(1.76)	(2.01)	(1.74)	(2.03)	(1.76)
Fraction in Manufacturing	0.29	0.29	0.23	0.26	0.24	0.29	0.22	0.22
Fraction in Transport	0.07	0.30	0.08	0.31	0.08	0.27	0.08	0.25
Fraction in Services	0.32	0.40	0.36	0.41	0.36	0.41	0.35	0.49

Notes: Summary statistics reported in October of each year. Wages are measured as pesos uruguayos deflated by the official Consumer Price Index (CPI). Firm size is measured as the log of total employment in each firm.

TABLE A.II

DESCRIPTIVE STATISTICS: LINKED EMPLOYER-EMPLOYEE PANEL DATA. WORKER-LEVEL INFORMATION

	1997	2001	2005	2009
All workers employed in PCs				
Observations	9,634	9,533	10,265	12,706
Fraction female	0.31	0.36	0.41	0.45
Fraction members	0.40	0.42	0.45	0.41
Average age	41.08	42.59	42.83	41.88
Average job tenure	9.20	9.85	9.81	8.75
Gross monthly wage	25,538	23,675	17,154	19,355
Daily wage	982	1,004	679	805
Fraction in Manufacturing	0.37	0.29	0.30	0.26
Fraction in Transport	0.31	0.30	0.25	0.21
Fraction in Services	0.30	0.39	0.42	0.48
Only those workers in WMFs				
Observations	3,270	3,202	3,898	4,417
Fraction female	0.15	0.14	0.24	0.27
Average age	42.23	44.02	44.61	43.94
Average job tenure	7.46	8.77	8.22	8.11
Gross monthly wage	23,757	22,594	16,243	17,629
Daily wage	944	890	666	811
Fraction in Manufacturing	0.17	0.08	0.15	0.13
Fraction in Transport	0.79	0.76	0.57	0.50
Fraction in Services	0.04	0.13	0.22	0.28

Notes: Summary statistics are reported in October of each year. Wages are measured as pesos uruguayos deflated by the official Consumer Price Index (IPC).

TABLE A.III

DESCRIPTIVE STATISTICS: LINKED EMPLOYER-EMPLOYEE PANEL DATA. FIRM-LEVEL INFORMATION

	1997	2001	2005	2009
All PCs				
Number of firms	241	262	285	309
Firm size (log of employment)	2.69	2.57	2.63	2.63
Firm average wage	11,027	9,785	7,153	9,259
Coef. of variation (daily wages)	0.25	0.27	0.32	0.32
Fraction female	0.23	0.28	0.35	0.39
Average age	42.10	43.11	43.35	43.77
Age dispersion (S.D.)	9.63	9.47	9.57	9.84
Average job tenure	4.33	5.18	5.22	5.45
Job tenure dispersion	2.33	2.90	3.26	3.69
Fraction in Manufacturing	0.25	0.18	0.19	0.18
Fraction in Transport	0.44	0.40	0.33	0.26
Fraction in Services	0.26	0.34	0.38	0.42
WMFs				
Number of firms	145	160	187	203
Firm size (log of employment)	2.50	2.37	2.52	2.54
Firm average wage	10,257	8,922	6,671	8,844
Coef. of variation (daily wages)	0.15	0.18	0.24	0.26
Fraction female	0.19	0.22	0.30	0.33
Average age	43.11	44.50	44.11	44.18
Age dispersion (S.D.)	9.50	9.44	9.53	9.74
Average job tenure	4.00	5.12	4.79	5.21
Job tenure dispersion	1.90	2.70	2.90	3.37
Fraction in Manufacturing	0.25	0.19	0.19	0.20
Fraction in Transport	0.59	0.53	0.40	0.31
Fraction in Services	0.14	0.20	0.27	0.33

Notes: Summary statistics are reported in October of each year. Wages are measured as pesos uruguayos deflated by the official Consumer Price Index (IPC). S.D. = standard deviation.

TABLE A.IV

DESCRIPTIVE STATISTICS ON WORKERS BY TRANSITION STATES

	% Female	Daily wage	Age	Tenure	Firm size	% Manufacturing	% Transport	% Services
Stayed in CFs	0,45	533,32 (497,19)	45,80 (13,27)	13,31 (9,24)	4,27 (2,09)	0,27	0,08	0,37
Stayed in WMFs	0,34	971,89 (638,47)	50,22 (10,02)	18,33 (9,42)	6,24 (1,66)	0,31	0,33	0,35
CF-to-CF movers	0,44	371,23 (370,28)	39,70 (11,28)	4,66 (5,91)	5,13 (1,85)	0,24	0,08	0,32
WMF-to-CF movers	0,50	672,30 (553,70)	46,18 (10,66)	10,71 (8,99)	6,06 (1,65)	0,16	0,13	0,60
CF-to-WMF movers	0,54	565,08 (528,96)	43,76 (10,35)	8,45 (8,74)	6,23 (1,38)	0,16	0,12	0,57

Notes: Wages are measured as pesos uruguayos deflated by the official Consumer Price Index (CPI). Firm size is measured as the log of total employment in each firm. All variables measured at the time of entry. In the case of movers, tenure is measured at the maximum tenure reached previous to the first transition.

TABLE A.V

WAGE GAP ACROSS THE WAGE DISTRIBUTION. RESULTS OF QUANTILE REGRESSIONS. PERIOD 1997-2009 (POOLED OLS)

	1997				2000			
	q20	q40	q60	q80	q20	q40	q60	q80
Coop	0.175***	0.095***	0.021***	−0.033***	0.192***	0.107***	0.037***	−0.028***
	(0.006)	(0.005)	(0.006)	(0.007)	(0.007)	(0.005)	(0.006)	(0.007)
Age	0.037***	0.055***	0.066***	0.081***	0.044***	0.059***	0.068***	0.088***
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
Age squared	−0.001***	−0.001***	−0.001***	−0.001***	−0.001***	−0.001***	−0.001***	−0.001***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Female	−0.213***	−0.251***	−0.288***	−0.319***	−0.182***	−0.222***	−0.253***	−0.271***
	(0.003)	(0.002)	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)
Tenure	0.056***	0.049***	0.045***	0.043***	0.063***	0.054***	0.0485***	0.045***
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
Tenure squared	−0.001***	−0.001***	−0.001***	−0.001***	−0.001***	−0.001***	−0.001***	−0.001***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Firm size	0.205***	0.183***	0.171***	0.166***	0.194***	0.173***	0.158***	0.149***
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
Test of interquantile Differences								
20th = 40th	[.000]				[.000]			
20th = 80th	[.000]				[.000]			
40th = 80th		[.000]				[.000]		
Observations	389,190	389,190	389,190	389,190	389,055	389,055	389,055	389,055

Notes: The dependent variable is the log of daily wages. The *Coop* dummy variable is set equal to 1 only for workers employed in a PC. Firm size is measured as the log of total employment in each firm. All estimates include six industry dummies. Bootstrapped standard errors (reported in parentheses) are based on 200 replications. *** significant at 1%.

TABLE A.V (CONTINUED)

WAGE GAP ACROSS THE WAGE DISTRIBUTION. RESULTS OF QUANTILE REGRESSIONS. PERIOD 1997-2009

	2003				2006				2009			
	q20	q40	q60	q80	q20	q40	q60	q80	q20	q40	q60	q80
Coop	0.142***	0.053***	–0.023***	–0.107***	0.159***	0.110***	0.040***	–0.040***	0.160***	0.114***	0.059***	–0.039***
	(0.006)	(0.006)	(0.006)	(0.009)	(0.007)	(0.005)	(0.007)	(0.009)	(0.008)	(0.005)	(0.005)	(0.006)
Age	0.039***	0.053***	0.064***	0.081***	0.039***	0.044***	0.054***	0.072***	0.031***	0.038***	0.045***	0.065***
	(0.001)	(0.001)	(0.001)	(0.002)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
Age squared	–0.001***	–0.001***	–0.001***	–0.001***	–0.001***	–0.001***	–0.001***	–0.001***	–0.001***	–0.001***	–0.001***	–0.001***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Female	–0.153***	–0.195***	–0.239***	–0.277***	–0.167***	–0.185***	–0.226***	–0.264***	–0.202***	–0.211***	–0.240***	–0.263***
	(0.003)	(0.003)	(0.003)	(0.004)	(0.002)	(0.002)	(0.002)	(0.003)	(0.002)	(0.002)	(0.002)	(0.003)
Tenure	0.061***	0.060***	0.053***	0.050***	0.039***	0.042***	0.044***	0.043***	0.038***	0.044***	0.048***	0.049***
	(0.001)	(0.001)	(0.001)	(0.001)	(0.005)	(0.005)	(0.001)	(0.001)	(0.000)	(0.000)	(0.000)	(0.001)
Tenure squared	–0.001***	–0.001***	–0.001***	–0.001***	–0.001***	–0.001***	–0.001***	–0.001***	–0.001***	–0.001***	–0.001***	–0.001***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Firm size	0.188***	0.170***	0.159***	0.156***	0.125***	0.125***	0.129***	0.141***	0.106***	0.107***	0.112***	0.122***
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
Test of interquantile Differences												
20th = 40th	[.000]				[.000]				[.000]			
20th = 80th	[.000]				[.000]				[.000]			
40th = 80th		[.000]				[.000]				[.000]		
Observations	340,130	340,130	340,130	340,130	429,504	429,504	429,504	429,504	492,771	492,771	492,771	492,771

Notes: The dependent variable is the log of daily wages. The *Coop* dummy variable is set equal to 1 only for workers employed in a PC. Firm size is measured as the log of total employment in each firm. All estimates include six industry dummies. Bootstrapped standard errors (reported in parentheses) are based on 200 replications. *** significant at 1%.

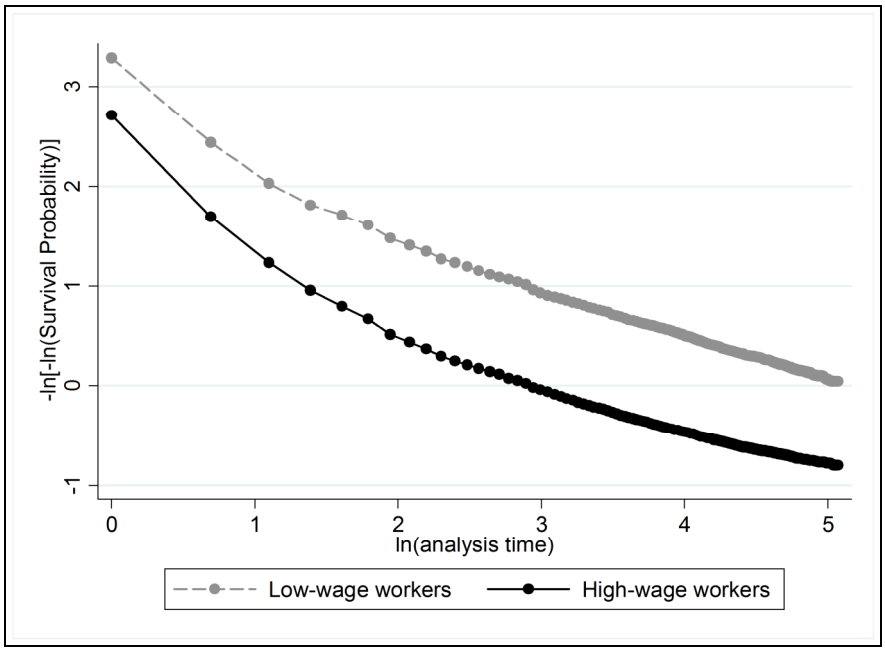

FIGURE A.I

Graphical Check of the Proportional Hazard assumption

Notes: This graph plots the transformation $-\ln[-\ln\{\hat{S}(t)\}]$ versus $\ln(t)$ for high- and low-wage members employed by WMFs, where $\hat{S}(t)$ is the Kaplan–Meier estimate of the survivor function. The proportional hazard assumption is not violated when the curves are parallel.

TABLE A.VI

REGRESSION ANALYSIS OF SWITCHERS BETWEEN CONVENTIONAL FIRMS

	(1)	(2)	(3)
	Probit	Logit	Logit
Log Wage	-0.242***	-0.404***	
	(0.007)	(0.012)	
Wage Quintile 2			-0.035
			(0.022)
Wage Quintile 3			-0.172***
			(0.024)
Wage Quintile 4			-0.291***
			(0.026)
Wage Quintile 5			-0.333***
			(0.030)
Female	-0.122***	-0.201***	-0.152***
	(0.009)	(0.016)	(0.016)
Age	0.090***	0.152***	0.147***
	(0.004)	(0.006)	(0.006)
Age squared	-0.001***	-0.002***	-0.002***
	(4.94e-05)	(8.37e-05)	(8.34e-05)
Tenure	-0.133***	-0.220***	-0.225***
	(0.002)	(0.004)	(0.004)
Tenure squared	0.002***	0.004***	0.004***
	(0.0001)	(0.0002)	(0.0002)
Firm size	0.263***	0.443***	0.418***
	(0.003)	(0.005)	(0.004)
Observations	94,680	94,680	94,680
Pseudo R2	0,2052	0,2054	0,1982

Notes: The dependent variable is a dummy that takes a value of one if the worker experienced a CF-to-CF transition. Log Wage is the log of daily wage. Wage Quintile indicates the position of the worker in her sector-specific wage distribution. Firm size is measured as the log of total employment in each firm. All variables measured at the time of entry (before exit). In the case of movers, tenure is measured as the maximum tenure reached previous to the first transition. All estimates include a set of thirteen cohort dummies and six industry dummies. Standard errors reported in parentheses. * significant at 10%; ** significant at 5%; *** significant at 1%.