The Role of Self-Employment in Developing Economies: Lifetime Evidence from a Structural Search Model*

Gustavo Canavire-Bacarreza† Mauricio Tejada‡

August 19, 2026

Abstract

Workers in developing economies spend much of their careers in self-employment rather than in wage work, but the consequences are usually assessed from income observed at a single point in time. We ask three questions that require a career horizon: whether self-employment leads into wage employment, how much of the time spent in it reflects search frictions rather than preference, and what its removal would do to employment and to the distribution of income. We estimate an on-the-job search model in which self-employment is a third labor-market state reachable from any other, termination risk declines with income, and participation is endogenous, using short panels and repeated cross sections from Ecuador, a setting where half of employment is self-employment, to simulate fifty-year careers. The self-employed receive wage offers as often as the unemployed, and beginning a career in self-employment delays wage employment for years without lowering lifetime income. Yet workers spend 79 percent of their self-employment time in activities they would leave for a wage job paying their group's average wage. Removing self-employment triples unemployment, raises non-participation among unskilled women from 42 to 63 percent, and lowers lifetime income by 17 percent overall and 46 percent for unskilled women. Its distributional effect reverses with the population and horizon considered: inequality falls among workers, rises 14 percent once the jobless are counted, and rises 23 percent over entire careers.

Keywords: Self-employment; Labor-market search; Lifetime income inequality.
JEL Codes: J46, J64, D31.

---

*Tejada gratefully acknowledges financial support from ANID – MILENIO – NCS 2022_045. The views expressed herein are those of the authors and do not necessarily reflect the views of the World Bank, its Board of Directors, or the countries it represents. We thank Carolina Velez Ospina for her support with the data.

†Poverty and Equity Global Practice, World Bank, USA. E-mail: gcanavire@worldbank.org.
‡Department of Economics, Universidad Diego Portales, Santiago Chile. E-mail: mauricio.tejada@udp.cl.

1 Introduction

One worker in three in Latin America is self-employed, more than twice the share in the average OECD economy.[1] The pattern extends across the developing world. In low-income countries, workers spend only 20 to 50 percent of their working lives in wage employment and much of the remainder in self-employment rather than open unemployment (Breza and Kaur, 2025). The self-employed also differ widely in their reasons for entering the sector. Some run businesses chosen for their income prospects, autonomy, or flexibility. Others work for themselves because they cannot find a wage job, often in small-scale activities that require little capital and rarely grow (Maloney, 2004; Gindling and Newhouse, 2014; La Porta and Shleifer, 2014). Most assessments of these two groups rely on income observed at one point in time, but current income does not show how self-employment affects work and income over an entire career.

Cross-sectional comparisons miss two features of self-employment in developing economies. First, self-employment differs from its counterpart in developed economies, where it more often represents entrepreneurship and financial capital strongly shapes entry (Cagetti and De Nardi, 2006; Boeri et al., 2020). In developing economies, capital appears less important, while labor-market frictions, particularly the difficulty of finding and keeping a wage job, play a larger role (Meghir et al., 2015; Narita, 2020). Second, workers in developing economies move across jobs and labor-market states much more frequently than workers in richer economies, and transitions through low-income self-employment often do not lead to better jobs (Donovan et al., 2023). These patterns make self-employment a question about careers, not only current income. Our estimated model shows that current earnings understate its career value. Although self-employment pays less than wage work on average, it provides a path into wage employment, sustains labor-force participation, supports reservation wages, and reduces inequality over the lifetime.

Ecuador provides a useful setting for studying these career effects. Labor income is the main determinant of household welfare. Rising labor earnings accounted for almost 86 percent of the reduction in poverty between 2007 and 2014, while the collapse of employment during the 2020 pandemic explained nearly 90 percent of the subsequent increase. Close to half of all workers are self-employed, a share that is high even by Latin American standards, and self-employment is more common among women and the poor. Among poor workers, 97 percent work for themselves or in firms with fewer than ten employees, largely without contracts or social protection. The sector includes small formal businesses with growth prospects and own-account activities taken up when wage jobs are unavailable (Canavire-Bacarreza, 2025). Mobility between these activities and wage employment

[1] According to the International Labour Organization's ILOSTAT database, self-employment accounted for 36 percent of total employment in Latin America and the Caribbean between 2015 and 2025, compared with 15 percent in OECD countries.

differs sharply across workers. Poor and less-educated workers who enter self-employment rarely return to wage jobs, whereas better-educated workers move between the two states more frequently (World Bank, 2025). Ecuador also adopted the U.S. dollar as legal tender in 2000, allowing the survey to report incomes in a stable common unit, which facilitates structural estimation.[2]

To establish the career value of self-employment, this paper addresses three questions. First, does self-employment provide a route into wage employment, or does starting a career in self-employment produce a lasting disadvantage? Second, how much does self-employment contribute to lifetime income, and how much of the time workers spend in it reflects search frictions rather than preference? Third, what would happen without self-employment to labor-force participation, unemployment, accepted wages, and inequality, both in the cross section and over the lifetime?

To answer these questions, we extend the on-the-job search model of Flinn (2002) in three directions. First, we add self-employment as a third labor-market state alongside unemployment and regular wage employment. Individuals receive entrepreneurial income opportunities while in every labor-market state and can move directly between wage employment and self-employment. The model therefore treats self-employment as an alternative to wage work rather than as a residual state. Second, termination rates decline with income in both working states, capturing the lower probability that higher-paying jobs and businesses will end. Third, individuals decide whether to participate in the labor market, allowing the model to account for the large out-of-the-labor-force population in developing economies (Flabbi and Leonardi, 2010).

We estimate the model by simulated method of moments using ENEMDU, Ecuador's national employment survey. We pool the 2022–2024 cross sections and the 2021–2023 rotating panels and divide workers into four groups by gender and schooling. Identification follows the standard arguments of Flinn and Heckman (1982) and Flinn (2002). Spell durations identify total exit hazards, observed transitions separate these hazards by destination state, accepted incomes identify the offer distributions, and income growth among workers who remain in the same job identifies search on the job. We then simulate fifty-year careers and summarize them using three measures: discounted lifetime income, which combines earnings and mobility (Flinn, 2002; Flabbi and Mabli, 2010; Tejada, 2016); the probability of holding a regular wage job at each point in the career, conditional on the initial labor-market state (Esteban-Pretel et al., 2011); and the share of self-employment time spent in positions workers would leave for a wage job paying their group's average wage, which measures how much self-employment reflects search frictions rather than preference.

The parameter estimates show that self-employment does not isolate workers from the wage sector.

[2]Employment quality reinforces this motivation. Over the past decade, job creation has concentrated in the least productive sectors, formality has fallen by nearly thirteen percentage points, and standard measures of job quality have deteriorated most sharply among the self-employed. Among poor workers, about 92 percent of employment is low skilled, informal, or self-employed.

Self-employed workers receive wage offers every four to eight months, about as often as unemployed workers and considerably more often than wage employees. Entrepreneurial opportunities arrive most frequently among workers who are already self-employed, indicating that exposure to the market is an important source of new business opportunities. The estimated termination rates also imply longer-duration self-employment spells. Termination shocks arrive once every 12 to 21 years, depending on the group, compared with once every 9 to 16 years in wage employment. These longer spells come with lower potential income. Mean offered wages exceed mean potential self-employment incomes by 40 to 55 percent among unskilled workers and by 130 to 150 percent among skilled workers. The opportunity cost of remaining self-employed therefore rises sharply with schooling.

Unskilled women face the greatest constraints in accessing wage employment. They receive wage offers about every fifteen months while unemployed and every four years while employed. At the same time, they receive self-employment opportunities more frequently than any other group. Unskilled women consequently have both the largest self-employment share and the highest non-participation rate. Their outcomes show how limited access to wage offers can make self-employment the main route into work for workers with otherwise weak employment prospects.

Simulated careers show the long-run consequences of these differences. One year after entering the labor market, a worker whose first job is in self-employment holds a regular wage job with probability 0.13, compared with probabilities between 0.55 and 0.62 for a worker who starts in wage employment. This difference closes after roughly eight years for unskilled women, 12 to 15 years for men, and almost 30 years for skilled women. Starting in self-employment therefore entails a long-lasting disadvantage but not permanent exclusion from wage work. Moreover, workers who start in self-employment and those who start unemployed follow nearly identical paths into wage employment. Self-employment thus provides a route into regular work comparable to searching while unemployed.

Search frictions nevertheless account for most of the time spent in self-employment. Workers spend 79 percent of their self-employment time in positions they would leave for a wage job paying their group's average wage. Workers who spend most of their career spells in self-employment have higher cross-sectional income and welfare, consistent with positive selection into more valuable businesses, but they do not have higher lifetime income. Mobility also reduces the inequality observed at a single point in time. Within each worker type, the standard deviation of lifetime income is about 40 percent of its mean, compared with 65 to 69 percent for income measured in the cross section.

The counterfactual without self-employment shows its equilibrium contribution to participation, wages, and inequality. When entrepreneurial opportunities are unavailable, marginal participants, especially unskilled women, leave the labor force. Non-participation rises for every group and increases from 42 to 63 percent among unskilled women. Workers who would otherwise enter self-employment move mainly into unemployment rather than wage employment, raising the aggregate unemployment rate from about 5 to 13 percent. Removing the self-employment option also lowers reservation wages. Mean accepted wages fall by one-fifth on average and by nearly one-half among unskilled women because workers accept offers they would otherwise reject.

Lifetime income falls by 17 percent in the counterfactual, but the losses differ sharply across groups. Skilled workers lose about one-tenth of their lifetime income, while unskilled women lose almost one-half. As a result, the near equality in lifetime income between unskilled men and women in the benchmark (246 and 233, in present-value terms) changes into a large gap (204 and 126). Inequality also responds differently at the upper and lower ends of the distribution. Among workers, removing self-employment lowers the Gini index to between 0.80 and 0.89 of its benchmark value because the upper tail loses high-income business owners. At the lower end, however, the fall in reservation wages increases the ratio of the middle to the lower tail of accepted wages by factors ranging from two to four. Adding jobless workers to the calculation raises the aggregate Gini by 14 percent. Over the lifetime, the Gini rises by 23 percent and the Theil index by 49 percent. The estimated model therefore shows that self-employment provides both opportunity and insurance. It raises inequality among workers at the upper end of the income distribution, but it supports incomes and employment at the lower end, where its removal harms the most vulnerable workers.

Our modeling approach builds on the literature that brings self-employment and informality into frictional models of developing-country labor markets. Albrecht et al. (2009) study an economy in which low-productivity workers sort into informal self-employment; Bosch and Esteban-Pretel (2012) show that differences in job creation and destruction between formal and informal jobs account for their cyclical behavior; Meghir et al. (2015) model formal and informal firms competing for the same workers; and Bobba et al. (2022) and Bobba et al. (2021) link informality to schooling and human-capital decisions. Most models in this literature treat self-employment as a state that workers enter from non-employment and focus on its steady-state value. The paper most closely related to ours is Narita (2020), who estimates a life-cycle search model for Brazil in which workers can be wage earners, self-employed, or unemployed. Our model differs in three respects. It allows workers to search for entrepreneurial opportunities from every labor-market state, including wage employment, so frequent movements from jobs into businesses result from workers' choices; it allows termination risk to decline with income in both working states; and it evaluates the career rather than only the current state, including lifetime income, the consequences of the first job, and the distribution of lifetime resources with and without self-employment.

A second strand documents the different forms that self-employment takes in developing countries. Gindling and Newhouse (2014) classify the self-employed across 74 countries and find that most are low-earning own-account workers rather than employers. Maloney (2004) and Bosch and Maloney (2010) document frequent movements between self-employment and wage work in Latin America and argue that much of the sector is voluntary. Falco and Haywood (2016) show that self-employment in Ghana increasingly attracts skilled workers, while experimental evidence shows that grants and training can move workers into more productive self-employment (Blattman et al., 2014; Martínez A. et al., 2018). These studies establish that self-employment combines different types of activities and that workers move frequently between labor-market states. Their short panels, however, cannot measure the lifetime consequences of entering or remaining in self-employment. Even Donovan et al. (2023), who use rotating panels from 49 countries, observe each worker for only a short period. Our structurally estimated model uses these short-term observations to simulate complete careers and quantify outcomes that the panels cannot measure directly.

A third strand studies how self-employment affects the income distribution in Latin America. Amarante (2016) finds that self-employment income reduced measured inequality in five Latin American countries during 2002–2011, while Bargain and Kwenda (2011) document large earnings penalties in informal self-employment in Brazil, Mexico, and South Africa. These studies rely on cross sections of workers. Our counterfactual shows that the population and time horizon included in the calculation change the conclusion. Removing self-employment lowers inequality among current workers, but it raises inequality once jobless workers enter the calculation and raises it further when income is measured over entire careers. To construct these lifetime measures, we follow Flinn (2002) and Flabbi and Leonardi (2010) for developed economies and Tejada (2016) for Chile. To our knowledge, this paper provides the first structurally estimated search model of self-employment for Ecuador and the first analysis of how removing self-employment affects lifetime income and inequality in a developing economy. The paper's contribution is therefore threefold. It models self-employment as an alternative to wage work, estimates the model in an economy where self-employment accounts for about half of employment, and uses the counterfactual to separate its effects at the upper and lower ends of the lifetime income distribution.

Section 2 formalizes workers' movements among unemployment, wage employment, and self-employment, together with the participation decision. Section 3 explains how the ENEMDU data identify the model's parameters and evaluates the model's fit. Section 4 defines the lifetime measures and quantifies the career consequences of passing through self-employment. Section 5 measures the equilibrium consequences of removing self-employment. Section 6 concludes.

2 Model

To analyze labor-market dynamics in a developing economy such as Ecuador, we extend the continuous-time on-the-job search model of Flinn (2002) in three dimensions[3].

First, we add self-employment as a third labor market state alongside unemployment and regular wage employment. In our model, individuals draw entrepreneurial income opportunities, search for them from every state, and can move directly between regular employment and self-employment, so the model highlights the role of self-employment as an alternative to wage work rather than treating it as a residual. Second, rather than assuming constant termination rates, we allow them to depend on the wage in regular employment and on net income in self-employment, capturing the idea that better-paid jobs and higher-income businesses are less likely to be destroyed. Finally, we add a participation margin. Individuals make a one-time decision about whether to enter the labor market, which allows the model to account for the large out-of-the-labor-force population that characterizes developing economies.

As in Flinn (2002), our model maintains the reservation-value property that makes all decisions tractable and maps naturally into the durations, accepted incomes, and transitions across states that we use in the estimation. The model also captures rich labor-market dynamics while preserving this tractability. This section describes the environment, derives the value functions and optimal decision rules, and characterizes the steady-state equilibrium.

2.1 Environment

Time is continuous, and the economy is populated by a large number of infinitely lived individuals, each belonging to one of four types: skilled men, unskilled men, skilled women, or unskilled women. We index gender by $i = M, W$ and schooling by $j = S, U$. The future is discounted at rate ρ. At any given point in time, an individual can be in one of four labor market states: out of the labor force, unemployed, employed in a regular wage job, or engaged in entrepreneurial self-employment activity. We denote the proportion of individuals of type ij in each of these states by o_{ij}, u_{ij}, e_{ij}, and s_{ij}, respectively. We normalize the total population of each worker type to one, so that $o_{ij} + u_{ij} + e_{ij} + s_{ij} = 1$ for all i and j.

Individuals are heterogeneous in how much they value non-labor-market activities, as represented by a draw z from the exogenous distribution $Q_{ij}(z)$. Based on this valuation, individuals make a one-time decision about whether to participate in the labor market. As Flabbi and Leonardi (2010)

[3] In the model of Flinn (2002), individuals move between unemployment and regular employment; job termination occurs at a constant rate, as is standard since Flinn and Heckman (1982); and there is no extensive-margin decision. Esteban-Pretel et al. (2011) develop a similar on-the-job search model but with two types of jobs, regular and contingent, while maintaining constant termination rates and focusing only on the intensive margin of labor supply.

indicate, although this is a stylized way to model the participation decision because it abstracts from transitions into and out of the labor force, it greatly simplifies the analysis while still allowing for interesting equilibrium effects.

Conditional on participating in the labor market, unemployed individuals receive a flow of income b_{ij}, which may come from unemployment insurance, government subsidies, or other non-labor income sources. They also search for both a regular wage job and an independent entrepreneurial opportunity (self-employment). Meetings with potential employers occur at the Poisson rate λ_{ij}^{u}, and job offers consist of a wage w drawn from the type-specific exogenous distribution $G_{ij}(w)$. In turn, self-employment opportunities arrive at the Poisson rate δ_{ij}^{u}, with income net of costs y drawn from the type-specific exogenous distribution $R_{ij}(y)$.

While working in a regular wage job, individuals receive a flow of income w net of taxes. Jobs are terminated, sending the individual into unemployment, at the Poisson rate $\eta_{ij}^{e}(w)$, with $d\eta_{ij}^{e}(w)/dw < 0$. We assume that the termination rate decreases with the wage, reflecting the idea that higher-paid jobs are more costly to terminate and thus less likely to be destroyed. The model also allows for on-the-job search: employees meet new regular wage opportunities at the Poisson rate λ_{ij}^{e} and self-employment opportunities at the Poisson rate δ_{ij}^{e}. These rates differ from their unemployment counterparts, capturing differences in search intensity and effort across states. As before, offered wages w and net incomes y are drawn from $G_{ij}(w)$ and $R_{ij}(y)$.

In the self-employment state, individuals receive a flow of income y net of business costs. Entrepreneurial activity is subject to termination shocks at the Poisson rate $\eta_{ij}^{s}(y)$, with $d\eta_{ij}^{s}(y)/dy < 0$, which send the individual into unemployment and represent a complete loss of production capacity. As in regular employment, the termination rate decreases with net income, reflecting the idea that higher incomes are associated with a greater capacity to sustain the entrepreneurial business. Individuals in this state also search for regular wage jobs, with offers arriving at the Poisson rate λ_{ij}^{s}, and their activity may experience income shocks that lead to income upgrades, arriving at the Poisson rate δ_{ij}^{s}.

2.2 Value Functions

The environment described above maps naturally into a set of Bellman equations that characterize the model's dynamics under optimal individual decision-making. To set notation, let $NP_{ij}(z)$ be the value of an individual who decides not to participate in the labor market, U_{ij} the value of an unemployed individual searching for a job, $E_{ij}(w)$ the value of an individual working in a regular job and earning wage w, and $S_{ij}(y)$ the value of an individual engaged in self-employment and earning y net of costs.

The Bellman equation that determines the value of non-participation in the labor market for an individual of type i, j is:

$$\rho \, NP_{ij}(z) = z. \tag{1}$$

Equation (1) can be interpreted in the usual asset-pricing way: the return on the asset (the labor market state) is determined by the events that change its value. Here, a non-participating individual simply receives the flow utility z as the asset return, and once the non-participation decision is made, the individual remains out of the labor force forever.

The Bellman equation that determines the value of unemployment for an individual of type i, j is:

$$\rho \, U_{ij} = b_{ij} + \lambda_{ij}^{u} \int \max \{E_{ij}(w) - U_{ij}, 0\} \, dG_{ij}(w) + \delta_{ij}^{u} \int \max \{S_{ij}(y) - U_{ij}, 0\} \, dR_{ij}(y). \tag{2}$$

Equation (2) again reflects how the return on the asset, unemployment, is determined by future events. The individual receives the flow utility b_{ij} and, at the rates λ_{ij}^{u} and δ_{ij}^{u}, obtains value gains of $E_{ij}(w) - U_{ij}$ and $S_{ij}(y) - U_{ij}$, respectively, whenever the job offer or self-employment opportunity is acceptable; that is, whenever $E_{ij}(w) \geq U_{ij}$ or $S_{ij}(y) \geq U_{ij}$.

The Bellman equation that determines the value of a regular wage job for an individual of type i, j is:

$$\begin{aligned} \rho \, E_{ij}(w) = w + \eta_{ij}^{e}(w) \, [U_{ij} - E_{ij}(w)] + \lambda_{ij}^{e} \int \max \{E_{ij}(w') - E_{ij}(w), 0\} \, dG_{ij}(w') \\ + \delta_{ij}^{e} \int \max \{S_{ij}(y) - E_{ij}(w), 0\} \, dR_{ij}(y). \end{aligned} \tag{3}$$

The return to an individual employed in a regular job at wage w is determined by the income flow and by potential future changes in value. A termination shock, arriving at the rate $\eta_{ij}^{e}(w)$, generates the value loss $U_{ij} - E_{ij}(w)$. The arrival of a wage offer w', at the rate λ_{ij}^{e}, generates the value gain $E_{ij}(w') - E_{ij}(w)$ when the new wage is acceptable; that is, when $E_{ij}(w') \geq E_{ij}(w)$. Finally, the arrival of a self-employment opportunity with net income y, at the rate δ_{ij}^{e}, generates the value gain $S_{ij}(y) - E_{ij}(w)$ when $S_{ij}(y) \geq E_{ij}(w)$.

Finally, the Bellman equation that determines the value of a self-employment activity for an individual of type i, j is:

$$\begin{aligned} \rho \, S_{ij}(y) = y + \eta_{ij}^{s}(y) \, [U_{ij} - S_{ij}(y)] + \lambda_{ij}^{s} \int \max \{E_{ij}(w) - S_{ij}(y), 0\} \, dG_{ij}(w) \\ + \delta_{ij}^{s} \int \max \{S_{ij}(y') - S_{ij}(y), 0\} \, dR_{ij}(y'). \end{aligned} \tag{4}$$

The return to an individual engaged in self-employment with net income y is likewise determined by the income flow and by potential future changes in value. A business-loss shock, arriving at the rate $\eta_{ij}^{s}(y)$, generates the value loss $U_{ij} - S_{ij}(y)$. The arrival of an acceptable regular job offer with wage w, at the rate λ_{ij}^{s} and satisfying $E_{ij}(w) \geq S_{ij}(y)$, generates the value gain $E_{ij}(w) - S_{ij}(y)$. Finally, the arrival of a better self-employment opportunity with income y', at the rate δ^s_{ij} and satisfying $S_{ij}(y') \geq S_{ij}(y)$, generates the value gain $S_{ij}(y') - S_{ij}(y)$.

2.3 Equilibrium

The model has a reservation-value property: all decisions can be characterized by a set of reservation values, namely reservation wages for regular employment and reservation incomes for self-employment. While unemployed, there is a reservation wage $\tilde{w}_{ij}$ such that $E_{ij}(\tilde{w}_{ij}) = U_{ij}$, and any offer with $w \geq \tilde{w}_{ij}$ is accepted; similarly, there is a reservation net income $\tilde{y}_{ij}$ such that $S_{ij}(\tilde{y}_{ij}) = U_{ij}$, and any $y \geq \tilde{y}_{ij}$ is accepted. While employed at wage w, a new offer w' is accepted if it exceeds the contingent reservation wage $\tilde{w}^e_{ij}(w)$, which satisfies $E_{ij}(\tilde{w}^e_{ij}(w)) = E_{ij}(w)$, and a self-employment opportunity with income y is accepted if it exceeds the contingent reservation value $\tilde{y}^e_{ij}(w)$, which satisfies $S_{ij}(\tilde{y}^e_{ij}(w)) = E_{ij}(w)$. While self-employed with income y, a regular job offer w is accepted if it exceeds the contingent reservation wage $\tilde{w}^s_{ij}(y)$, which satisfies $E_{ij}(\tilde{w}^s_{ij}(y)) = S_{ij}(y)$, and a new business opportunity y' is accepted if it exceeds the contingent reservation value $\tilde{y}^s_{ij}(y)$, which satisfies $S_{ij}(\tilde{y}^s_{ij}(y)) = S_{ij}(y)$.

In a steady-state equilibrium, both the distribution of individuals across labor market states and the distributions of wages and net incomes within each state remain constant over time, so the flow of individuals out of each state equals the flow into it. Because the termination rates depend on incomes and acceptance decisions depend on the reservation values, the transition rates between states have to integrate over the earnings of the individuals currently in each state. Let $G^{ss}_{ij}(w)$ denote the steady-state distribution of wages among employed individuals of type ij, and let $R^{ss}_{ij}(y)$ denote the corresponding distribution of net incomes among the self-employed.[4] The flow-balance conditions for unemployment, regular employment, and self-employment are:

$$\left[\lambda^u_{ij}\bar{G}_{ij}(\tilde{w}_{ij}) + \delta^u_{ij}\bar{R}_{ij}(\tilde{y}_{ij})\right] u_{ij} = \left[\int \eta^e_{ij}(w)\, dG^{ss}_{ij}(w)\right] e_{ij} + \left[\int \eta^s_{ij}(y)\, dR^{ss}_{ij}(y)\right] s_{ij} \tag{5}$$

$$\left[\int \left(\eta^e_{ij}(w) + \delta^e_{ij}\bar{R}_{ij}(\tilde{y}^e_{ij}(w))\right) dG^{ss}_{ij}(w)\right] e_{ij} = \lambda^u_{ij}\bar{G}_{ij}(\tilde{w}_{ij})u_{ij} + \left[\lambda^s_{ij}\int \bar{G}_{ij}(\tilde{w}^s_{ij}(y))\, dR^{ss}_{ij}(y)\right] s_{ij} \tag{6}$$

$$\left[\int \left(\eta^s_{ij}(y) + \lambda^s_{ij}\bar{G}_{ij}(\tilde{w}^s_{ij}(y))\right) dR^{ss}_{ij}(y)\right] s_{ij} = \delta^u_{ij}\bar{R}_{ij}(\tilde{y}_{ij})u_{ij} + \left[\delta^e_{ij}\int \bar{R}_{ij}(\tilde{y}^e_{ij}(w))\, dG^{ss}_{ij}(w)\right] e_{ij} \tag{7}$$

In equation (5), the unemployed exit at the rate $\lambda^u_{ij}\,\bar{G}_{ij}(\tilde{w}_{ij}) + \delta^u_{ij}\,\bar{R}_{ij}(\tilde{y}_{ij})$, with $\bar{G}_{ij} = 1 - G_{ij}$ and $\bar{R}_{ij} = 1 - R_{ij}$, upon accepting either a job offer or a self-employment opportunity, and they are replaced by employed and self-employed individuals whose matches are terminated, at the average

[4] The invariant distributions G^{ss}_{ij} and R^{ss}_{ij} are equilibrium objects, determined by the analogous within-state requirement that, for every w, the flow of workers into jobs paying at most w equals the flow out of those jobs, and likewise for self-employment incomes. In estimation, we solve jointly for G^{ss}_{ij}, R^{ss}_{ij}, and the labor market states by simulation.

rates $\int \eta^e_{ij}(w)\, dG^{ss}_{ij}(w)$ and $\int \eta^s_{ij}(y)\, dR^{ss}_{ij}(y)$. In equation (6), the employed exit upon termination or upon accepting a self-employment opportunity, which occurs with probability $\bar{R}_{ij}(\tilde{y}^e_{ij}(w))$ for a worker earning w. Those who exit employment are replaced by unemployed individuals who accept a job offer and by self-employed individuals who accept a wage offer, which occurs with probability $\bar{G}_{ij}(\tilde{w}^s_{ij}(y))$ for a self-employed worker earning y. Equation (7) is the analog for self-employment.[5] Finally, the normalization $o_{ij} + u_{ij} + e_{ij} + s_{ij} = 1$ determines o_{ij} once u_{ij}, e_{ij}, and s_{ij} are determined.

Putting all these ingredients together, we are ready to state a formal definition of equilibrium in this model.

Definition. Given the mobility parameters $\left\{\lambda^u_{ij}, \lambda^e_{ij}, \lambda^s_{ij}, \delta^u_{ij}, \delta^e_{ij}, \delta^s_{ij}, \eta^e_{ij}, \eta^s_{ij}\right\}$ and the offer distributions $G_{ij}(w)$ for wages and $R_{ij}(y)$ for self-employment net incomes, a steady-state labor market equilibrium for each type ij is a set of value functions $\{U_{ij}, E_{ij}(w), S_{ij}(y)\}$, a set of reservation values $\left\{\tilde{w}_{ij}, \tilde{y}_{ij}, \tilde{w}^e_{ij}(\cdot), \tilde{y}^e_{ij}(\cdot), \tilde{w}^s_{ij}(\cdot), \tilde{y}^s_{ij}(\cdot)\right\}$, and an invariant distribution of labor market states and within-state earnings $(u_{ij}, e_{ij}, s_{ij}, o_{ij}, G^{ss}_{ij}, R^{ss}_{ij})$ such that: (i) the value functions solve the Bellman equations (1)–(4); (ii) the reservation values satisfy the indifference conditions characterized above; and (iii) the steady-state flow-balance conditions (5)–(7), together with the participation decision and the population normalization, hold.

3 Estimation

This section describes the data, presents the estimation method and the identification strategy, and discusses the resulting parameter estimates and the fit of the model to the data.

3.1 Data

We estimate the model using data from the Ecuadorian *Encuesta Nacional de Empleo, Desempleo y Subempleo* (ENEMDU), a nationally representative household survey collected by the National Institute of Statistics and Censuses (INEC). The survey records individual demographics, labor-market status, occupation, labor income, hours worked, job tenure, and search duration, which together provide the cross-sectional information the model requires. We pool togheter the 2022, 2023, and 2024 waves of the survey. This data contributes with the distribution of individuals across labor-market states, the distributions of labor income, and the durations of unemployment, employment, and self-employment spells. Because Ecuador has used the U.S. dollar as legal currency since 2000, all

[5] Job-to-job transitions, which occur at the rate λ^e_{ij}, and transitions within self-employment, which occur at the rate δ^s_{ij}, do not appear because they leave the state shares unchanged and only shift the within-state distributions G^{ss}_{ij} and R^{ss}_{ij} toward higher earnings.

monetary variables are denominated in real term. For the dynamics, we use the year-to-year transitions matrices across labor-market states from ENEMDU's rotating-panel component, which re-interviews individuals roughly twelve months apart, matching the panels for 2021–2022 and 2022–2023.

To ensure some degree of homogeneity within gender ($i = M, W$) and schooling ($j = S, U$) groups, consistent with the model, we restrict our sample to individuals between 25 and 55 years old living in urban areas.[6] We define skilled workers as those with completed tertiary education and unskilled workers as those with at most a secondary degree. Additionally, we classify each individual into one of four labor-market states using the standard definitions based on the activity-condition and occupational-category variables in the survey: out of the labor force (o), unemployed (u), wage-employed (e), or self-employed (s). The self-employed comprise own-account workers and employers.

Labor income is measured as real hourly labor income: monthly labor earnings—wages for employees and earnings net of business costs for the self-employed—divided by usual monthly hours of work and deflated by the national consumer price index to December-2023 dollars. Durations are measured in months. We use completed job tenure for employees and the self-employed, and elapsed search duration for the unemployed. After the sample restrictions described above and data cleaning[7], the resulting sample contains roughly 49,000 to 86,000 observations, depending on the worker type. Table 1 reports descriptive statistics for the cross-section, and Table 2 reports the yearly transition matrix across labor market states conditional on participating in the labor market, that is, u, e, and s. Appendix A documents in detail the data sources, variable construction, and cleaning steps.

Labor market state distributions differ across the four worker types most sharply along two margins: labor-force attachment and the size of the self-employment sector. Non-participation is particularly important among unskilled women, about 42% of whom are out of the labor force, which contrasts sharply with 17% of skilled women and even more with the 5–7% observed among men. Women's non-participation is clearly associated with schooling, whereas there are no important differences among men by skill level. Self-employment, by contrast, is large for all worker types, ranging from 21% to 35% of the population and decreasing with schooling for both genders. Conditional on participation, the importance of self-employment is even more pronounced. Indeed, among those who work, the self-employed are a 57% majority of unskilled women, compared with 39% of unskilled men, 33% of skilled men, and 27% of skilled women. The low wage-employment share of unskilled women is therefore not merely a participation phenomenon, because those who do work are predominantly self-employed. Finally, unemployment is relatively low for all worker types (between 3% and 6%) and somewhat higher among the skilled.

Hourly income on average is higher in wage employment than in self-employment for every type:

[6] Labor markets in rural areas differ from the environment described in Section 2.

[7] We drop observations with missing labor-market state or occupation and, among the employed, those with missing or zero hourly income or missing tenure.

unskilled men, 3.17 versus 2.86 US$/hour; skilled men, 6.22 versus 4.74; unskilled women, 2.86 versus 2.31; skilled women, 5.67 versus 3.83. The gap in means widens with schooling, from 11% to 31% for men and from 24% to 48% for women, and is considerably larger for women than for men at both skill levels.

Self-employment income is also markedly more dispersed than wage income. Computing the coefficient of variation—the ratio of the standard deviation to the mean—for each type, wage employment ranges from 0.59 to 0.73, while self-employment ranges from 1.10 to 1.47, roughly 1.8 to 2.1 times higher within each type. Self-employment is therefore a lower-mean and higher-risk (variance) labor market state for every type. The skill gradient in dispersion, however, differs by gender and by state: among men, dispersion rises with schooling in both wage employment (0.60 to 0.73) and self-employment (1.26 to 1.47), so that skilled men are the most dispersed group in the entire income distribution; among women, dispersion rises only mildly with schooling in wage employment (0.59 to 0.62) but falls with schooling in self-employment (1.24 to 1.10), so that skilled women are the least dispersed group among the self-employed.

Finally, the skill premium is roughly two-to-one in wage employment for both genders and about 1.7 in self-employment, and within each skill level, men always earn more than women in both states.

Unemployment durations vary systematically by skill and gender. Unskilled workers remain unemployed for fewer months than skilled workers—about 3.0 versus 4.3 months for men, and 3.6 versus 4.4 months for women—consistent with the skilled being able to afford to wait longer for a better match. Skilled men and women remain unemployed for a similar length of time, whereas among the unskilled, men leave unemployment more frequently than women.

Employment durations reveal similar variability across gender and skill. Among the unskilled, men remain employed longer than women, 10.3 versus 7.1 years, while among the skilled the ranking reverses, though the gap narrows considerably: women remain employed a bit longer than men, 8.7 versus 8.2 years. Taken together, women leave regular jobs more frequently than men, a pattern driven mostly by the unskilled. Employment duration also rises with income: across skill level and gender, workers earning above the median remain employed longer than those earning below it, with above-median spells lasting from about 1.1 times as long (unskilled men) to nearly twice as long (skilled women) as below-median spells, suggesting more job retention as income rises.

Self-employment durations also vary across gender and skill, and in most cases self-employment stays longer on the job than employment for the same type of worker. Among the unskilled, self-employment lasts longer for men than for women, 12.5 versus 9.3 years, though self-employment stays longer relative to employment for women (32% longer) than for men (22% longer). Among the skilled, self-employment again lasts longer for men than for women, 9.2 versus 7.3 years; relative to employment, self-employment stays longer for men (12% longer) but shorter for women (17%

less)—the only type for which self-employment is the less durable state. The highest differences in duration are below the median: self-employment spells exceed employment spells by 18% for unskilled men, 52% for skilled men, 37% for unskilled women, and 17% for skilled women, whereas above the median the difference narrows to 23%, 8%, and 4% for unskilled men, skilled men, and unskilled women, respectively, and turns into a 24% shortfall for skilled women. In self-employment, therefore, low-income workers stay more, while at higher incomes employment becomes the relatively more durable state, particularly for skilled women. Overall, self-employment seems to be a more persistent state.

The yearly transition matrix confirms that both work states are persistent and that workers move across occupations. Indeed, wage and self-employment are both highly persistent, with one-year stay rates of 84–93% and 75–86% respectively. For employees, the probability of staying increases with schooling for both genders, from 84.5% to 88.5% among men and, more sharply, from 84.6% to 92.9% among women, so that skilled women are the most likely to stay employed of any type. For the self-employed, the probability of staying is essentially flat across skill for men, 78.2% versus 78.0%, but falls sharply with schooling for women, from 86.3% among the unskilled—the highest stay rate of any type—to 75% among the skilled—the lowest. Crucially, departures from self-employment almost never lead to unemployment—the self-employment-to-unemployment flow is only 2–5%—and overwhelmingly lead to wage work, so self-employment behaves as a stepping-stone rather than a revolving door into joblessness. The probability of staying unemployed, by contrast, ranges from 24% to 35% across types. Within a year, the probability that an unemployed worker moves into employment ranges from 32% to 50%, and into self-employment from 26% to 38%. By skill, the probability of moving into employment is higher for unskilled men than skilled men, 49.5% versus 41.7%, but higher for skilled women than unskilled women, 37.9% versus 32.1%, while the probability of moving into self-employment is roughly flat by skill for men, 26.3% versus 27.9%, and falls sharply with skill for women, from 37.9% to 26.8%. By gender, the probability of moving into employment is higher for men than women at both skill levels, though the gap narrows from 17 percentage points among the unskilled to under 4 among the skilled, whereas the probability of moving into self-employment is higher for women than men, a difference concentrated among the unskilled, where women's probability of moving into self-employment is nearly 1.5 times that of men.

From employment, the probability of moving into unemployment is low, ranging from 2.1% to 3.6% across types, while the probability of moving into self-employment is comparatively high, ranging from 5% to 12.8%. By skill, the probability of moving into self-employment falls with schooling for both genders, from 12.8% to 8.9% among men and, more sharply, from 11.8% to 5% among women, while the probability of moving into unemployment is roughly flat by skill for men, 2.7% versus 2.6%, but falls with schooling for women, from 3.6% to 2.1%. By gender, the probability of moving into unemployment is higher for women than men among the unskilled, 3.6% versus 2.7%, but the ranking reverses among the skilled, where men move into unemployment more than women, 2.6% versus 2.1%; the probability of moving into self-employment, by contrast, is higher for men than women at both skill levels, and the gap widens sharply with schooling, from 1 percentage point among the unskilled to nearly 4 among the skilled. This suggests that the probability of moving from employment to self-employment is not low, and that these transitions represent workers who are genuinely better off in self-employment.

From self-employment, the probability of moving into unemployment is low, ranging from 1.9% to 5.4% across types, while the probability of moving into employment is comparatively high, ranging from 11.3% to 19.9%. By skill, the probability of moving into unemployment rises with schooling for both genders, from 1.9% to 3.1% among men and, more sharply, from 2.4% to 5.4% among women, while the probability of moving into employment is roughly flat by skill for men, 19.9% versus 18.9%, but rises sharply with schooling for women, from 11.3% to 19.6%. By gender, the probability of moving into unemployment is higher for women than men at both skill levels, and the gap widens with schooling, from 0.5 percentage points among the unskilled to 2.3 among the skilled; the probability of moving into employment, by contrast, is much higher for men than women among the unskilled, 19.9% versus 11.3%, but the genders converge among the skilled, 18.9% versus 19.6%. The probability of leaving self-employment for a regular job is not negligible, suggesting that whenever the self-employed have the opportunity to move into employment, they do.

3.2 Estimation Method

We estimate the parameters of the model by the Simulated Method of Moments (SMM; Gouriéroux and Monfort, 2002), a method commonly used to estimate highly non-linear models such as ours. The model is estimated separately by gender $i = M, W$ and schooling level $j = S, U$, so that we allow for heterogeneity in every parameter across these groups. Let Θ_{ij} denote the set of parameters to be estimated for type ij.

Let $\mathcal{M}^D_{N_{ij}}$ denote a set of appropriately chosen moment statistics computed from our data sample of size N_{ij}, and let $\mathcal{M}_B(\Theta_{ij})$ denote the corresponding set of moment statistics computed from a simulated sample of size B drawn from the steady-state equilibrium implied by Θ_{ij}. The SMM estimator $\hat{\Theta}_{ij}$ minimizes the weighted quadratic distance between the two sets of statistics:

$$\hat{\Theta}_{ij}(\Omega_{ij}) = \arg\min_{\Theta_{ij}} \frac{1}{2}\left[\mathcal{M}^D_{N_{ij}} - \mathcal{M}_B(\Theta_{ij})\right]' \Omega_{ij} \left[\mathcal{M}^D_{N_{ij}} - \mathcal{M}_B(\Theta_{ij})\right], \tag{8}$$

where Ω_{ij} is a symmetric, positive-definite weighting matrix. We use a diagonal weighting matrix whose nonzero entries are the inverse of the bootstrapped variance of each moment in the sample. To compute $\mathcal{M}_B(\Theta_{ij})$ we set the annual discount rate to $\rho = 0.1$ and simulate the careers of $B = 15{,}000$

individuals, each starting in unemployment and lasting 50 years. We then build the simulated statistics from cross-sections taken 12 months apart over the portion of the horizon in which the distribution of labor market states has become invariant, so that the moments are computed at the steady-state equilibrium implied by Θ_{ij}.

For each type ij, the parameter vector Θ_{ij} collects sixteen parameters; to reduce the burden of notation, we omit the type subscripts on the individual parameters in the remainder of this subsection, as in Table 3. The mobility parameters are the arrival rates of regular-job and self-employment opportunities from each state, $\lambda^u, \lambda^e, \lambda^s$ and $\delta^u, \delta^e, \delta^s$, together with the termination-rate parameters η_0^e, η_1^e and η_0^s, η_1^s. Consistent with the model's assumption that termination rates decline with income (Jovanovic, 1979; Jolivet et al., 2006), we parametrize them as $\eta^e(w) = \eta_0^e + \eta_1^e \exp(-\eta_1^e w)$ and $\eta^s(y) = \eta_0^s + \eta_1^s \exp(-\eta_1^s y)$. The earnings parameters are those of the offer distributions. Following Flinn and Heckman (1982), we assume lognormality, with location and scale parameters (μ_e, σ_e) for the distribution of wage offers $G(w)$ and (μ_s, σ_s) for the distribution of self-employment income offers $R(y)$. Finally, b is the flow income received while unemployed, and γ parametrizes the distribution of the value of non-participation $Q(z)$, which, following Flabbi and Mabli (2010), we assume to be exponential with mean γ.

The vector of statistics $\mathcal{M}^D_{N_{ij}}$ comprises 31 moments that summarize the labor market behavior of each type. These include the shares of individuals in each labor market state (non-participation, unemployment, regular employment, and self-employment); the mean, standard deviation, and fifth percentile of income among the employed and among the self-employed; the mean and standard deviation of the durations of unemployment, employment, and self-employment; the nine elements of the monthly transition matrix across unemployment, employment, and self-employment, with the monthly matrix obtained from the yearly panel transitions[8]; the mean durations of employment and of self-employment below and above the within-type median income; and the mean income growth of those who remain employed and of those who remain self-employed.

3.3 Identification

Although all sixteen parameters are estimated jointly by minimizing the loss function in (8), the moment vector is constructed so that each block of moment statistics is informative about a specific piece of the parameter vector, following closely the identification arguments of Flinn and Heckman (1982) and, for environments with search on the job, Flinn (2002). As in these papers, durations provide information about total exit hazards, observed transitions allow us to decompose them into

[8] We recover the monthly matrix from the annual one via its eigendecomposition, $P = V\Lambda^{1/12}V^{-1}$, where P denotes the monthly transition matrix; see Section 3.3 for the time-homogeneity assumption and embeddability condition this requires.

competing exit states, and accepted incomes provide information about the underlying offer distributions (truncated at the reservation wages or earnings). Moreover, estimating the model for each gender and schooling type as a separate labor market, with no cross-type restrictions, accounts for observable heterogeneity. In this subsection we discuss the mapping from moment statistics to parameters; as in the previous subsection, we omit the type subscripts.

We begin with the duration moments. Because exits from each labor market state are governed by competing Poisson processes, spell durations identify total exit hazards. The total hazards out of unemployment, wage employment, and self-employment, respectively, are:

$$\begin{aligned} h_u &= \lambda^u \bar{G}(\tilde{w}) + \delta^u \bar{R}(\tilde{y}) \\ h_e &= \int_{\tilde{w}} \left[\eta^e(w) + \lambda^e \bar{G}(\tilde{w}^e(w)) + \delta^e \bar{R}(\tilde{y}^e(w))\right] dG^{ss}(w) \\ h_s &= \int_{\tilde{y}} \left[\eta^s(y) + \lambda^s \bar{G}(\tilde{w}^s(y)) + \delta^s \bar{R}(\tilde{y}^s(y))\right] dR^{ss}(y), \end{aligned}$$

where the integrals average income-specific hazards over the steady-state distributions of wages and net incomes. Spell durations measure tenure in the current income spell. A spell therefore ends not only when an individual leaves the state, but also upon an accepted within-state move: a job-to-job transition in employment or an upgrade to a better activity in self-employment. The acceptance terms thus enter the two hazards even though these moves leave state shares unchanged. We observe elapsed durations for spells in progress, sampled from the stock of individuals in each state. Under stationarity, these durations remain informative about the underlying hazards (Flinn, 2002). The distribution of individuals across $\{u, e, s, o\}$ provides additional steady-state information: because inflows equal outflows in each state, observed shares further restrict the mobility parameters.

The monthly transition matrix across u, e, and s further decomposes each total hazard by destination state. For example, for unemployed workers, the probabilities of moving to wage employment and self-employment separately identify $\lambda^u \bar{G}(\tilde{w})$ and $\delta^u \bar{R}(\tilde{y})$. The same logic applies when employment or self-employment is the source state. Because the rotating panel observes workers one year apart, observed transitions may combine several monthly moves, creating a time-aggregation problem. We mitigate this problem by recovering the monthly transition matrix from the annual matrix through its eigendecomposition, $P = V\Lambda^{1/12}V^{-1}$, where P denotes the monthly transition matrix.[9] This procedure assumes that monthly transition probabilities remain constant within each year.

Accepted wages and net incomes provide information about the underlying offer distributions $G(w)$ and $R(y)$ through the recoverability argument of Flinn and Heckman (1982). We assume

[9]This treats the underlying process as a time-homogeneous Markov chain: the same monthly transition matrix P compounds twelve times into the observed annual one, P^{12}, consistent with the model's own steady-state, time-homogeneous transitions. A well-defined monthly root in turn requires the annual matrix to be embeddable—real positive eigenvalues and a nonnegative root—which holds in our data for all four types.

lognormal offer distributions, which satisfy the recoverability condition: the distribution of accepted offers uniquely identifies the parameters of the underlying offer distribution and its truncation point. The mean and standard deviation of accepted wages and net incomes in each working state inform the location and scale parameters (μ_e, σ_e) and (μ_s, σ_s). The fifth-percentile moments locate the lower tail of the observed distributions and thus identify the truncation implied by the reservation values $\tilde{w}$ and $\tilde{y}$. These truncation points also identify the arrival rates: once $\bar{G}(\tilde{w})$ and $\bar{R}(\tilde{y})$ are known, the acceptance probabilities convert observed flows out of unemployment into λ^u and δ^u (Flinn, 2002). With search on the job, however, the accepted-income cross-section is not itself the truncated offer distribution but a mixture of order statistics, as workers climb the job ladder. We return to this initial-conditions problem below.

The conditional duration moments identify the income gradient of the termination rates, our main departure from the constant termination-rate specification of Flinn and Heckman (1982) and Flinn (2002). The $e \to u$ and $s \to u$ elements of the transition matrix pin down only the *average* termination rate over the corresponding accepted-income distribution and therefore cannot distinguish the level of the hazard from its slope. Under our two-parameter specification, the four conditional duration moments, mean employment and self-employment durations below and above the within-state median income, provide this separation. Below-median durations identify the excess low-income hazard η_1, which determines both its level and rate of decay. Above the median, where the exponential term has largely died out, durations identify the asymptotic floor η_0. Assigning η_1 this double role keeps the income gradient one-dimensional, so two conditional moments per labor market state are sufficient.

Identifying the on-the-job parameters raises the so-called initial-conditions problem. Repeated cross-sections provide no information about where each individual stands within an employment or self-employment spell. Observed income therefore comes from a mixture over spell orders whose weights depend on all the structural parameters. Flinn (2002) discusses two solutions. The first assumes that the process is in steady state and uses the implied point-sampled distribution; the second conditions on observed income and uses only the forward-looking portion of the likelihood. We adopt the steady-state solution because the distribution of labor-market histories in our prime-age sample is plausibly close to its steady state. This assumption would be less plausible for samples concentrated near labor-market entry or retirement. We therefore draw simulated cross-sections from the invariant region of simulated careers and compute their statistics exactly as we compute their data counterparts. The simulated and observed cross-sections thus embody the same mixture over spell orders.

The steady-state assumption alone provides only weak discipline on the labor-market distribution along the income ladders, so we add another source of information. We match mean monthly income growth among workers who remain employed and among those who remain self-employed between panel waves. These moments are informative because the upgrade probabilities $\bar{G}(\tilde{w}^e(w))$ and $\bar{R}(\tilde{y}^s(y))$ decline as workers climb their respective income ladders. Mean income growth among stayers therefore provides direct information about the labor-market distribution along each ladder.[10] Together with the within-state resets in the duration hazards, these growth rates identify the on-the-job arrival rates λ^e and δ^s, neither of which is observed in the three-state transition matrix.

Three parameters remain: the discount rate, the unemployment income flow, and the parameter governing the distribution of the value of non-labor-market activities. The data do not allow us to jointly identify the first two parameters, so, following Moore et al. (2020), we fix $\rho = 0.1$ and estimate b. Given the other parameters, b shifts the reservation values $\tilde{w}$ and $\tilde{y}$ monotonically. The fifth-percentile moments identify the truncation points, while unemployment durations together with the flows out of unemployment identify the acceptance margins. Together, these moments identify the reservation values. Flinn (2002) exploits this mapping to recover b ex post by inverting the reservation-wage equation at his point estimates. Finally, following Flabbi and Mabli (2010), we assume that the distribution $Q(z)$ is exponential with mean γ. An individual enters the labor market if and only if the value of search exceeds the value of remaining outside, $z \leq \rho U$. Under this assumption, the non-participation rate is $o = \exp(-\rho U/\gamma)$. As discussed above, ρU is identified, so the non-participation rate identifies γ.

3.4 Results

Table 3 reports the estimated parameters for each of the four worker types. Table 4 presents the implied arrival times of offers and termination shocks, together with the mean and standard deviation of the distributions of wage and self-employment income offers. Throughout, we report the shares in unemployment, regular employment, and self-employment as fractions of labor-market participants and the non-participation rate as a fraction of each worker type's population.

Across the four worker types, the wage-offer arrival rate in self-employment, λ^s, is similar to the wage-offer arrival rate in unemployment, λ^u, and both exceed the wage-offer arrival rate in wage employment, λ^e. The exception is unskilled women, for whom the wage-offer arrival rate in self-employment, λ^s, far exceeds the corresponding rate in unemployment, λ^u. This ordering implies that self-employed workers receive wage offers at least as frequently as unemployed workers and more frequently than wage-employed workers. Table 4 shows that self-employed workers receive a wage offer every four to eight months. Unemployed workers wait about six to eight months for three worker types and about fifteen months for unskilled women. Among wage-employed workers, the corresponding

[10] Flinn (2002) uses life-cycle wage growth only for out-of-sample validation of his estimated model, whereas we include it among the estimation moments.

wait is about twelve to nineteen months for three types and forty-nine months for unskilled women. Self-employment therefore does not isolate workers from the wage sector; it can function as a search platform with access to wage offers comparable to that of open unemployment. The opposite ordering generally holds for self-employment opportunities. The arrival rate of self-employment opportunities while self-employed, δ^s, is the highest for every worker type, and the corresponding rate while wage-employed, δ^e, exceeds the rate while unemployed, δ^u, except among skilled women. This pattern indicates that entrepreneurial opportunities arise mainly through market exposure, especially while workers are already self-employed, rather than through active search from unemployment.

Termination shocks arrive less frequently in self-employment than in regular employment for every worker type. The mean interval between termination shocks is 21.2 versus 15.5 years for unskilled men, 17.3 versus 9.2 years for skilled men, 18.7 versus 10.0 years for unskilled women, and 12.3 versus 9.3 years for skilled women (Table 4). These estimates describe the duration of self-employment as a labor-market state rather than the duration of a single business. Self-employment includes heterogeneous activities, and the data classify individuals who close one business and start another as continuously self-employed.

The estimated termination parameters explain this pattern. The asymptotic termination-rate floors are small: the self-employment termination floor, η_0^s, ranges from 0 to 0.002, while the wage-employment termination floor, η_0^e, ranges from 0.001 to 0.005. As wages or net incomes rise, termination risk approaches these low floors, consistent with the model's assumption that higher-income matches are more costly to terminate. Most termination risk comes from the income-sensitive components governed by the termination shape parameters. Each shape parameter determines both the size of the income-sensitive component at zero income and its rate of decay. The wage-employment termination shape parameter, η_1^e, is nearly unchanged across schooling levels for men, rising from 0.056 to 0.057, but increases from 0.071 to 0.087 for women. The self-employment termination shape parameter, η_1^s, varies less across groups, ranging from 0.049 to 0.060. Skilled women also have the only positive self-employment termination floor, $\eta_0^s = 0.002$. This floor, together with their self-employment termination shape parameter, $\eta_1^s = 0.058$, is consistent with their having the shortest interval between self-employment termination shocks. This pattern indicates that the lowest-income wage jobs and self-employment activities face the highest termination risk.

For every worker type, the mean offered wage exceeds mean potential self-employment income, and the gap widens sharply with schooling. Among unskilled workers, mean wage offers exceed potential self-employment incomes by roughly 40 to 55 percent, 0.92 versus 0.66 US$/hour for men and 0.94 versus 0.62 for women. Among skilled workers, the gap rises to 130 to 150 percent, 2.21 versus 0.89 for men and 1.78 versus 0.77 for women. The pre-match skill premium, calculated from the underlying offer distributions, is correspondingly much larger for wages, about 2.4 times for men, than for potential self-employment income, about 1.35 times. This comparison indicates that schooling is associated with substantially larger potential income gains in wage employment than in self-employment. Potential self-employment income is also more dispersed than wage offers for every worker type, as reflected in its consistently higher coefficient of variation. Taken together, these patterns characterize self-employment as a lower-income activity than wage employment and imply a sizable opportunity cost for skilled workers who remain self-employed.

The most pronounced heterogeneity is between unskilled and skilled women. By a wide margin, unskilled women have the lowest wage-offer arrival rates in unemployment and wage employment, λ^u and λ^e. They wait about 14.9 months for a wage offer while unemployed, compared with 5.7 to 7.8 months for the other worker types, and more than four years, 49 months, while employed. Yet they have the highest wage-offer arrival rate in self-employment, λ^s, at 0.254, and relatively high arrival rates of self-employment opportunities in unemployment, wage employment, and self-employment, δ^u, δ^e, and δ^s. In the data, unskilled women also have the highest self-employment share, 54% of participants, compared with 25 to 37% for the other types, and the highest non-participation rate, 41.5%, compared with 5 to 17%. This combination of limited access to wage offers and abundant self-employment opportunities is consistent with a stronger preference for flexibility among unskilled women. Despite their high wage-offer arrival rate in self-employment, λ^s, unskilled women rarely move from self-employment to wage employment, implying that they reject most wage offers. The implied acceptance rates are 17%, 11%, 4%, and 12% for unskilled men, skilled men, unskilled women, and skilled women, respectively, with the rate for unskilled women by far the lowest. Skilled women display a different pattern. Once wage-employed, their arrival rate of self-employment opportunities, δ^e, is only 0.016, about one opportunity every five years, consistent with greater investment in wage employment than in flexible self-employment. A related asymmetry appears in offered wages. The mean offered wage is nearly identical across genders among unskilled workers, 0.94 for women and 0.92 for men, but about 20% lower for women among skilled workers, 1.78 compared with 2.21. Thus, among unskilled workers, the gender gap in accepted wages arises despite nearly identical mean offers and must reflect other margins, including arrival rates, reservation wages, and offer dispersion. Among skilled workers, part of the gap is already present in mean offers.

With respect to estimation precision, the bootstrapped standard errors reported in Table 3 indicate that the location and scale parameters of the wage-offer distribution, (μ_e, σ_e), the corresponding parameters of the self-employment income-offer distribution, (μ_s, σ_s), and the mean of the non-labor-market value distribution, γ, are the most precisely estimated, with relative standard errors generally between 1 and 8%. Cross-sectional income and participation moments tightly anchor these parameters. The wage-offer arrival rates across labor market states, $(\lambda^u, \lambda^e, \lambda^s)$, the arrival rates of self-employment opportunities, $(\delta^u, \delta^e, \delta^s)$, and the income-sensitive termination shape parameters,

(η_1^e, η_1^s), are moderately precise, with relative standard errors typically between 3 and 15%. The unemployment income flow, b, is among the least precisely estimated parameters, with relative standard errors ranging from 9% for skilled men to 41% for unskilled women.

Finally, we assess how well the estimated model reproduces the data. Table 5 compares the targeted moments statistics in the data with their counterparts generated by the estimated model. Overall, the model captures the central features of the data used to simulate labor market careers. Labor market state shares are matched within about three percentage points for every worker type. Mean wages differ from the data by about 5%, and mean self-employment incomes differ by about 10% for three of the four worker types. For unskilled women, the model overstates mean self-employment income by about a quarter. The model also closely reproduces within-state persistence probabilities and the principal flows out of unemployment, while capturing many of the longer, higher-income spell durations reasonably well.

Some discrepancies remain. Simulated unemployment durations are generally about twice as long as those observed in the data, suggesting that the constant arrival-rate specification does not fully capture short unemployment spells. The model understates the dispersion of wages and self-employment incomes by 20 to 37% in most cells and generates lower transitions between wage employment and self-employment. The fit is also weaker for low-income self-employment durations, which the model understates by 60 to 70%, and for income growth among self-employed stayers, particularly skilled workers. The largest discrepancies concern durations and earnings dispersion, while the model closely matches labor market state shares, within-state persistence, and the principal flows out of unemployment. Because the model tends to compress earnings dispersion, the inequality measures derived from simulated careers may be conservative. The fit of the main cross-sectional moments and transitions provides a reasonable basis for the lifetime and counterfactual exercises that follow.

4 Lifetime Measures

4.1 Definitions

An advantage of a fully estimated structural model is that it allows us to simulate complete labor market careers. We use these simulations to construct an income measure that accounts for earnings in each labor market state and mobility across states over the life cycle. Following Flinn (2002) and Flabbi and Mabli (2010) for developed economies, and Tejada (2016) for Chile, we summarize each simulated career by its discounted lifetime income.

Consider an individual k of type ij, where i indexes gender and j indexes schooling. The individual's simulated career contains L_{kij} labor market spells indexed by $\ell = 1, \ldots, L_{kij}$. Let t_ℓ denote the elapsed time at the beginning of spell ℓ, with $t_1 = 0$ and $t_{L_{kij}+1} = 600$ months (50 years). Spell ℓ therefore lasts $t_{\ell+1} - t_\ell$. The discounted lifetime income of individual k is

$$LI_{kij} = \sum_{\ell=1}^{L_{kij}} e^{-\rho t_\ell} \int_{t_\ell}^{t_{\ell+1}} z_{kij}^{\ell}\, e^{-\rho(v-t_\ell)}\, dv, \tag{9}$$

where ρ is the model's discount rate, v is the elapsed time since the beginning of the career, and z_{kij}^{ℓ} is the flow income received during spell ℓ. Flow income equals unemployment income, b_{ij}, during unemployment, wage w during regular employment, and net income y during self-employment. Within each integral, $e^{-\rho(v-t_\ell)}$ discounts income received at time v to the beginning of spell ℓ. The factor $e^{-\rho t_\ell}$ then discounts the value of that spell to the beginning of the career.

Thus, LI_{kij} is the present value of the income stream generated over the individual's career. It reflects both income within each labor market state and the transitions that determine when, and for how long, that income is received. This dynamic content distinguishes lifetime income from cross-sectional income. The steady-state cross section records participants' income at a single point in time, whereas LI_{kij} captures the long-run resources available over the working life (Tejada, 2016). Comparing the two distributions shows whether mobility across labor market states smooths or amplifies the income dispersion observed at a point in time.

We complement lifetime income with a long-run measure of welfare. Following Flinn (2006) and Flabbi and Leonardi (2010), we use stationarity to define welfare for each type as the expected value among labor market participants. This measure averages the state-specific values U_{ij}, $E_{ij}(w)$, and $S_{ij}(y)$ using the invariant distribution across labor market states and the within-state earnings distributions. Let $o_{ij}, u_{ij}, e_{ij}, s_{ij}$ denote the invariant population proportions of the four labor market states, where $u_{ij} + e_{ij} + s_{ij} = 1 - o_{ij}$. Welfare per participant is

$$\mathcal{W}_{ij} = \frac{1}{1-o_{ij}} \left[u_{ij}\, U_{ij} + e_{ij} \int E_{ij}(w)\, dG_{ij}^{ss}(w) + s_{ij} \int S_{ij}(y)\, dR_{ij}^{ss}(y) \right], \tag{10}$$

where $G_{ij}^{ss}(w)$ and $R_{ij}^{ss}(y)$ are the steady-state distributions of wages among employed workers and net incomes among self-employed workers, respectively, as defined in Section 2.3. The factor $1/(1-o_{ij})$ conditions the distribution of labor market states on participation, so $\mathcal{W}_{ij}$ is the expected value of a participant of type ij.

We also use the value functions to construct an approximate measure of the extent to which self-employment reflects search frictions rather than a preference for self-employment. Consider a self-employed worker of type ij earning net income y. If $E_{ij}(\bar{w}_{ij}) > S_{ij}(y)$, the worker would accept a regular job paying the group-specific average wage if such an offer arrived. The worker remains self-employed because no such offer has arrived. By contrast, if $S_{ij}(y) > E_{ij}(\bar{w}_{ij})$, the worker would reject the offer and therefore prefers the current self-employment position to the benchmark regular job. For each simulated career containing at least one self-employment spell, we define the proportion of time in self-employment spent in low-value positions as

$$\Pi_{kij} = \frac{\sum_{\ell=1}^{L_{kij}} (t_{\ell+1} - t_\ell) \mathbf{1}\{q_{kij}^\ell = s\} \, \mathbf{1}\{S_{ij}(y_{kij}^\ell) < E_{ij}(\bar{w}_{ij})\}}{\sum_{\ell=1}^{L_{kij}} (t_{\ell+1} - t_\ell) \mathbf{1}\{q_{kij}^\ell = s\}} \tag{11}$$

where q_{kij}^ℓ denotes the labor market state during spell ℓ, y_{kij}^ℓ denotes net income during a self-employment spell, and $t_{\ell+1} - t_\ell$ is the spell duration. Thus, Π_{kij} is the proportion of the individual's time in self-employment spent in positions that the worker would leave for a regular job paying the average wage. We denote the mean of this individual-level measure for type ij by Π_{ij}. A larger value indicates that search frictions account for a greater proportion of the group's self-employment histories.

We use these measures to compare the full sample with subsamples defined by their self-employment histories. The first subsample comprises individuals whose first job after the initial unemployment spell is self-employment. The second comprises individuals who spend the majority of their labor market spells in self-employment. These comparisons show whether beginning a career in self-employment or remaining predominantly attached to it is associated with lower lifetime income and welfare, and whether low-value positions are more common in such careers.

Finally, following Esteban-Pretel et al. (2011), we use the simulated careers to assess whether self-employment serves as a stepping stone toward regular employment or as a state in which workers tend to remain. For a given horizon τ, we compute the probability of holding a regular job τ periods later, conditional on the starting point. Given the large simulated samples, we estimate this probability as the fraction of workers in regular employment at the end of the corresponding time window.

We consider two starting points. The first comparison conditions on the initial job taken after unemployment, contrasting careers that begin in self-employment with those that begin in regular employment. This comparison shows whether beginning in self-employment lowers the subsequent probability of holding a regular job and how long any gap persists. The second comparison conditions on the state from which the search for regular employment begins, contrasting unemployment with self-employment. This comparison shows whether self-employment provides a route into regular employment or instead delays entry into that state.

4.2 Results

For each of the four worker types Table 6 reports the mean and standard deviation of cross-sectional income, lifetime income LI_{kij}, and the individual low-value proportion Π_{kij}, together with the welfare measure $\mathcal{W}_{ij}$. It reports these measures for the full simulated sample and the two self-employment subsamples. Figures 1 and 2 plot the probability of holding a regular job over the first 30 years of the career, conditional on the starting point.

Panel B of Table 6 shows that schooling generates the largest differences in lifetime income. Lifetime income averages 246 for unskilled men and 233 for unskilled women, compared with 502 for skilled men and 429 for skilled women. Thus, skilled workers earn roughly twice as much over their careers as unskilled workers. Gender gaps are smaller but vary considerably by schooling. Women's lifetime income is about 13, or 5 percent, lower than men's among unskilled workers, whereas it is about 73, or 15 percent, lower among skilled workers. Welfare follows the same ranking as in the cross-section, from skilled men to skilled women, unskilled men, and unskilled women ($\mathcal{W}_{MS} > \mathcal{W}_{WS} > \mathcal{W}_{MU} > \mathcal{W}_{WU}$). Lifetime outcomes therefore preserve the between-group differences observed in the steady-state cross-section.

A lifetime perspective changes the assessment of inequality within worker types. For all types of workers, the coefficient of variation is about 0.40 for lifetime income, compared with between 0.65 and 0.69 for cross-sectional income. Mobility across labor-market states therefore smooths a substantial share of the dispersion observed at a point in time. As discussed in Section 4.1, the lifetime measure averages income across the unemployment, wage-employment, and self-employment spells in a worker's career. It thereby removes transitory variation that raises cross-sectional inequality, leaving the more persistent differences across worker types.

Conditioning on a first job in self-employment is associated with little change in mean outcomes, as shown in the "Start as SE" columns of Table 6. All differences relative to the overall sample are within 3 percent. Lifetime income is virtually unchanged for unskilled men, at 246.6 compared with 246.5 overall, and in the aggregate, at 335.1 compared with 335.0. Among skilled workers, lifetime income is modestly lower for those who start in self-employment, by about 3 percent for men (489 compared with 502) and 2 percent for women (420 compared with 429). By contrast, it is about 3 percent higher for unskilled women (240 compared with 233). Welfare differs by less than 1 percent for every worker type and is marginally higher in the aggregate (40.3 compared with 40.0). The distribution of lifetime income is also similar in the two samples. Its aggregate coefficient of variation is 0.54 in both, although within-type dispersion is slightly higher among workers who start in self-employment, particularly skilled women. Between-group differences narrow modestly: the gender gap falls from about 5 to 3 percent among unskilled workers and from 15 to 14 percent among skilled workers, while schooling gaps also decline. Thus, starting in self-employment changes neither average lifetime outcomes nor their overall dispersion substantially.

Workers who spend the majority of their spells in self-employment have higher cross-sectional income and welfare within every worker type, as shown in the "Most time as SE" columns of Table 6.

Relative to the overall means, their cross-sectional income is between 7 and 20 percent higher, and their welfare is between 5 and 16 percent higher. Lifetime income, however, is between 1 and 4 percent lower in every type, with the largest difference among skilled women. Higher cross-sectional income and steady-state welfare therefore do not translate into higher lifetime income.

Lifetime income is also more dispersed among workers whose majority of spells are in self-employment. Within worker types, its coefficient of variation ranges from 0.41 to 0.50, compared with 0.39 to 0.42 in the overall sample. In the aggregate, it rises from 0.54 to 0.57. The combination of higher cross-sectional income and steady-state welfare, slightly lower lifetime income, and greater dispersion is consistent with positive selection into durable self-employment based on the value of the current business. Because higher-income self-employment activities face lower termination rates, workers operating more valuable businesses are more likely to accumulate careers with a majority of self-employment spells, whereas workers in lower-value activities are more likely to leave for wage employment or unemployment. This selection does not imply that having a majority of spells in self-employment raises lifetime income.

Aggregating across worker types reverses the pattern for cross-sectional income and welfare. Among workers whose majority of spells are in self-employment, mean cross-sectional income is 3.77, compared with 3.98 overall, and welfare is 37.0, compared with 40.0. These lower aggregate means arise even though both outcomes are higher within every worker type. The reason is compositional: the aggregate is a weighted average of the type-specific means, and unskilled workers and women, who have lower income and welfare on average, account for a larger share of the group whose majority of spells are in self-employment. Their greater weight more than offsets the higher means observed within each type.

The aggregate lifetime-income gap combines this composition effect with small differences within types. Lifetime income is 283 among workers whose majority of spells are in self-employment, compared with 335 overall, a decline of about 15 percent. Within each worker type, however, the decline is only between 1 and 4 percent. Most of the aggregate gap therefore reflects the concentration of careers with a majority of self-employment spells among lower-income worker types, while a smaller part reflects lower lifetime income within types. These aggregate comparisons describe who selects into persistent self-employment and should not be interpreted as the effect of self-employment on an otherwise identical worker.

Panel C of Table 6 shows that self-employed workers spend a large share of their time in self-employment in positions they would leave for a regular job paying the group-specific average wage. The mean low-value share (Π_{ij}) ranges from 63 to 85 percent across worker types and equals 79 percent in the aggregate. It is highest among skilled women and men, at 85 and 83 percent, and lowest among unskilled women, at 63 percent. This ranking partly reflects the higher value of the average regular job among skilled workers, which sets a higher benchmark for their self-employment positions to exceed.

Beginning a career in self-employment changes this pattern little. The aggregate low-value share is 78 percent among workers who start in self-employment, compared with 79 percent overall, and its dispersion is similar in the two samples. By contrast, the share falls within every worker type among those whose majority of spells are in self-employment, to about 65 percent for skilled and unskilled men, 59 percent for unskilled women, and 76 percent for skilled women. This pattern is consistent with workers in more valuable businesses remaining self-employed longer. Unlike income and welfare, the within-type and aggregate comparisons point in the same direction.

Considerable heterogeneity nevertheless remains among workers whose majority of spells are in self-employment. Across worker types, the standard deviation of the low-value share ranges from 29 to 36 percentage points, compared with 27 to 32 percentage points overall. In the aggregate, it rises from 41 to 47 percentage points. Thus, even within this group, some workers spend relatively little of their self-employment time in low-value positions, while others spend a substantial share of their self-employment time in such positions.

Figure 1 shows, for each worker type, the probability of wage employment during the first 30 years after labor-market entry, conditional on whether the worker's first job is in self-employment or wage employment. In the first year, this probability is about 0.13 among workers who start in self-employment, compared with 0.55 to 0.62 among those who start in wage employment. The gap narrows for every worker type, and the two paths eventually converge or become nearly indistinguishable. Catch-up is slow, however, and for skilled women takes almost the entire 30-year horizon. Thus, an initial spell of self-employment is better characterized as a long-lived scar than as a permanent dead end. It reduces the probability of wage employment for a substantial part of the working life, but the disadvantage eventually fades. This persistence is consistent with the long duration of self-employment spells: entering self-employment early in a career can shape employment outcomes for a decade or more without permanently excluding the worker from wage employment.

Figure 2 compares workers who begin their careers in self-employment with those who begin unemployed. The two paths start much closer together than in Figure 1, with an initial gap of about 0.10 to 0.25, and converge within roughly 20 years. Workers who start in self-employment therefore reach wage employment only slightly more slowly than those who start unemployed. The larger initial gap in Figure 1 reflects its different benchmark: workers who start in wage employment are already in the state whose probability is being measured, whereas both groups in Figure 2 begin outside wage employment. In this sense, self-employment provides an alternative route into regular employment, as in Esteban-Pretel et al. (2011). The similar paths suggest that workers use self-employment to search for regular jobs much as they use unemployment.

The figures also reveal substantial differences across worker types in the duration of the initial disadvantage and the level to which the paths converge. In Figure 1, the gap closes fastest for unskilled women, within roughly eight years, and slowest for skilled women, near the end of the 30-year horizon. Both groups of men fall between these extremes, with convergence after roughly 12 to 15 years. Figure 2 shows the same broad ordering, but all gaps close within about 20 years. The long-run probability of wage employment is highest for skilled workers, at about 0.67 to 0.68, followed by unskilled men, at about 0.60, and unskilled women, at about 0.43. This ranking mirrors the steady-state shares of wage employment among participants.

Taken together, the results show that self-employment is a heterogeneous labor-market state. Workers whose majority of spells are in self-employment are positively selected into relatively valuable businesses and have higher cross-sectional income and steady-state welfare, although not higher lifetime income. Most self-employment time, however, is spent in positions workers would leave for a regular job paying the group-specific average wage, indicating that search frictions account for much of the time spent in self-employment. Starting a career in self-employment leaves a long-lasting scar on the probability of wage employment, but this disadvantage eventually fades and does not translate into lower lifetime income or welfare on average. Finally, the similar employment paths of workers who start unemployed and self-employed suggest that both states serve as platforms for obtaining regular jobs.

5 Counterfactual Scenario

5.1 Definitions

In this section, we ask what the labor market and the distribution of income would look like if self-employment were unavailable. We construct a counterfactual economy without self-employment, solve for its equilibrium, and compare it with the estimated benchmark. Specifically, we set the three self-employment arrival rates to zero, $\delta_{ij}^{u} = \delta_{ij}^{e} = \delta_{ij}^{s} = 0$, so that no self-employment opportunities arrive from any labor-market state. Removing self-employment offers from the value functions (2)–(4) produces an equilibrium with a zero self-employment share, in which unemployed and wage-employed workers search only for wage offers. We hold every other parameter at its estimated value. The experiment therefore isolates the role of self-employment as both a route into work and an outside option that supports the value of labor-market participation.

Shutting down self-employment substantially changes the structure of the labor market. We therefore solve the new equilibrium and resimulate careers, allowing reservation wages $\tilde{w}_{ij}$, the steady-state distribution across labor-market states, spell durations, and the distribution of simulated careers to adjust. We then recompute the cross-sectional labor-market outcomes, the distribution of lifetime income LI_{kij} defined in (9), and the inequality measures. Tables 7 and 8 report each counterfactual value and, in parentheses, its ratio to the benchmark. Because the estimated model understates the dispersion of wages and self-employment incomes, the inequality changes should be interpreted as a lower bound on the distributional role of self-employment.

5.2 Results

The counterfactual shows that self-employment supports labor-market participation and provides an outside option for workers with limited access to wage jobs. Removing it pushes workers out of the labor force, roughly triples unemployment, lowers mean accepted wages, and reduces lifetime income for every worker type. Its effect on inequality depends on the population and time horizon considered. Within every worker type, removing self-employment lowers inequality among current workers, although the aggregate Gini remains unchanged because differences across group means increase. Once unemployed workers are included, and especially when income is measured over entire careers, self-employment lowers inequality. These effects are largest among the groups that rely most on self-employment in the benchmark. The self-employment share among participants is about 54 percent for unskilled women, 39 percent for unskilled men, 29 percent for skilled men, and 26 percent for skilled women. The burden of removing self-employment therefore falls most heavily on unskilled women.

Table 7 first shows how removing self-employment changes the allocation of workers across labor-market states. Non-participation rises for every worker type by factors ranging from 1.18 to 1.73. The largest change occurs among unskilled women, whose non-participation increases from 42 to 63 percent. For marginal participants, who are disproportionately unskilled women, self-employment provides an important source of work. Once this option disappears and wage prospects worsen, the value of searching falls below the value of remaining out of the labor force. Among those who continue to participate, the absence of self-employment raises unemployment rather than producing a comparable expansion in wage employment. The aggregate unemployment rate increases from about 5 to 13 percent. The proportional increase is largest among the groups with the highest initial self-employment shares, reaching factors of 4.1 for unskilled women and 3.3 for unskilled men, compared with 1.6 for skilled women. Although wage employment becomes the only working state and its share among participants rises mechanically, the main response is higher unemployment because workers lose one of the two routes out of joblessness.

Removing self-employment also lowers accepted wages. Mean accepted wages fall for every worker type, by 22 percent in the aggregate and by 44 percent for unskilled women. The mechanism is the loss of the outside option. Without self-employment, the reservation wage $\tilde{w}_{ij}$ falls, and workers accept wage offers they would previously have rejected.

Unemployment spells also become substantially longer. Mean unemployment duration increases by factors ranging from 1.6 to 3.8, with the largest increases among unskilled workers. Without self-employment, only an acceptable wage offer can end an unemployment spell, whereas a self-employment opportunity provides an additional route out of unemployment in the benchmark. Mean employment duration remains essentially unchanged in the aggregate but falls slightly for women, including a 13 percent decline for skilled women. Two effects largely offset each other: removing transitions into self-employment lengthens wage-employment spells, while lower accepted wages raise termination rates and shorten them.

These labor-market changes reduce lifetime income for every worker type. Mean lifetime income falls by 17 percent in the aggregate. The losses are smallest for skilled workers, at about 9 to 10 percent, and largest for unskilled women, whose lifetime income falls by 46 percent. Self-employment therefore substantially narrows the gender gap among unskilled workers. In the benchmark, mean lifetime income is about 246 for unskilled men and 233 for unskilled women, a gap of about 5 percent. Without self-employment, these values fall to 204 and 126, respectively, widening the gender gap to about 38 percent. Self-employment thus accounts for much of the near equality in lifetime income between unskilled men and women.

Table 8 shows that the effect on cross-sectional inequality depends on the part of the income distribution considered. Among workers, excluding the unemployed, removing self-employment lowers the Gini and Theil indices within every worker type. The Gini falls to between 0.80 and 0.89 of its benchmark value, while the Theil index falls to between 0.63 and 0.78. The largest declines occur among skilled men. Self-employment creates high-income business opportunities that widen the upper part of the income distribution within worker types, so removing these opportunities lowers inequality among workers. At the aggregate level, however, the Gini remains unchanged because wider differences in mean income across worker types offset the declines within types.

The lower part of the worker income distribution moves in the opposite direction. The P_{50}/P_{10} ratio increases by factors ranging from 2.0 to 3.6 across worker types, while the P_{90}/P_{50} ratio changes by factors ranging from 0.98 to 1.20. Thus, most of the increase in the P_{90}/P_{10} ratio comes from lower incomes at the bottom rather than higher incomes at the top. This result reflects the fall in reservation wages. Without self-employment as an outside option, some workers accept wage offers they would have rejected in the benchmark, lowering income at the bottom of the worker distribution. The falling Gini and Theil indices within worker types indicate that the loss of high-income self-employment opportunities outweighs this decline at the bottom in these summary measures.

When unemployed workers and their flow income are included in the distribution, the aggregate result reverses. The aggregate Gini rises to 1.14 times its benchmark value, while the aggregate Theil index rises to 1.31 times its benchmark value. Both measures also rise within three of the four worker types, with the largest increases among unskilled women. Skilled women are the exception, with the Gini and Theil indices remaining close to their benchmark values. Two mechanisms account for the increase at the bottom of the distribution. Removing self-employment raises unemployment and lowers the accepted wages of workers who remain employed. Self-employment therefore lowers inequality in the broader distribution, even though it raises inequality within worker types when the comparison is restricted to current workers.

Over entire careers, the inequality results are less ambiguous. Except for skilled women, all five measures of lifetime inequality in Panel B of Table 8 rise when self-employment is removed. The aggregate lifetime Gini increases by a factor of 1.23, while the Theil index increases by a factor of 1.49. Income differences also increase more in the lower half of the distribution. The aggregate P_{50}/P_{10} ratio rises by a factor of 1.59, compared with 1.30 for the P_{90}/P_{50} ratio. High income from a successful self-employment spell represents only part of a worker's career income, while longer unemployment spells and lower accepted wages reduce income over multiple periods. Skilled women are the exception. For this group, all lifetime inequality measures remain between 0.95 and 0.99 of their benchmark values, consistent with their relatively low reliance on self-employment.

In sum, the effect of self-employment on inequality depends on the population and time horizon considered. Among current workers, self-employment creates high-income business opportunities that raise inequality within worker types. At the same time, it supports reservation wages, provides a route out of unemployment, and sustains labor-market participation among workers with limited access to wage jobs. Once unemployed workers are included, and especially when income is measured over entire careers, these effects at the lower end of the distribution dominate. Self-employment therefore raises lifetime income and lowers inequality in the broader population, with the largest effects among unskilled women, who rely on it most as a route into work.

6 Conclusions

This paper evaluates self-employment over entire careers rather than at a single point in time. We estimate an on-the-job search model in which workers can move among unemployment, wage employment, and self-employment and can choose whether to participate in the labor market. The estimated model shows that self-employment does not isolate workers from wage jobs. Self-employed workers receive wage offers about as frequently as unemployed workers. Workers whose first job is in self-employment are initially much less likely to hold a wage job than those who begin in wage employment, but this difference closes after roughly eight years for unskilled women, 12 to 15 years for men, and almost 30 years for skilled women. Moreover, workers who begin in self-employment and those who begin unemployed follow nearly identical paths into wage employment. Starting in self-employment therefore produces a long-lasting initial disadvantage, but not permanent exclusion from wage work.

Self-employment nevertheless often reflects limited access to acceptable wage jobs rather than a preference for independent work. Termination rates imply long-duration self-employment spells, even though potential self-employment income is lower than offered wages for every worker group. Workers spend 79 percent of their self-employment time in activities they would leave for a wage job paying their group's average wage. Despite this difference in current income, starting a career in self-employment changes mean lifetime income by no more than 3 percent across worker groups. Mobility also reduces inequality within groups. The standard deviation of lifetime income is about 40 percent of its mean, compared with 65 to 69 percent for income measured in the cross section. These results show why current income alone provides an incomplete assessment of the career value of self-employment.

The counterfactual without self-employment quantifies the value of this option. Removing self-employment pushes marginal participants out of the labor force, raises the aggregate unemployment rate from about 5 to 13 percent, and lowers mean accepted wages by one-fifth because the loss of the outside option reduces reservation wages. Lifetime income falls by 17 percent on average and by almost one-half among unskilled women. The inequality results depend on the population and time horizon considered. Among current workers, removing self-employment lowers inequality because the worker distribution loses high-income business owners. Once unemployed workers are included, the aggregate Gini rises by 14 percent. Over entire careers, the Gini rises by 23 percent and the Theil index by 49 percent. Self-employment therefore supports participation, reservation wages, lifetime income, and equality at the lower end of the income distribution, especially among unskilled women.

These findings point to limited access to acceptable wage offers as the binding constraint, particularly for unskilled women, who wait more than a year for a wage offer while unemployed. Policies that increase the arrival rate or quality of wage offers could reduce low-income self-employment by allowing workers to leave voluntarily for better jobs. By contrast, the counterfactual cautions against policies that raise the cost of self-employment without expanding access to wage work. Within the model, removing this option increases non-participation and unemployment and lowers accepted wages and lifetime income among the workers who rely on it most.

Our analysis has four limitations. First, the model is stationary and in partial equilibrium. Wage-offer distributions and wage-offer arrival rates remain fixed in the counterfactual, so the results do not include firms' responses in wages, vacancies, or job creation. Second, the model combines formal businesses and informal own-account activities in a single self-employment state and therefore cannot separate their roles. Third, the estimated model understates the standard deviations of wages and self-employment incomes by 20 to 37 percent in most worker groups, which suggests that the estimated inequality responses are conservative. Fourth, individuals make the participation decision once, so the model does not capture movements into and out of the labor force over the life cycle. The conclusions should therefore be interpreted as equilibrium responses within the estimated worker-side model, rather than as predictions that incorporate firms' responses or differences between formal and informal self-employment.

Declaration of generative AI and AI-assisted technologies in the writing process

During the preparation of this work the authors used Claude (Anthropic) in order to proof-read and improve the writing in English, and Elicit in order to search for relevant literature. After using these tools, the authors reviewed and edited the content as needed and take full responsibility for the content of the published article.

Replication Package

The replication package for this paper is available at https://github.com/mauriciotejada/self_employment_lifetime.

References

Albrecht, James, Lucas Navarro, and Susan Vroman, "The Effects of Labour Market Policies in an Economy with an Informal Sector," *Economic Journal*, 2009, *119* (2000), 1105–1129.

Amarante, Verónica, "Income Inequality in Latin America: A Factor Component Analysis," *Review of Income and Wealth*, 2016, *62* (S1), S4–S21.

Bargain, Olivier and Prudence Kwenda, "Earnings Structures, Informal Employment, and Self-Employment: New Evidence from Brazil, Mexico, and South Africa," *Review of Income and Wealth*, 2011, *57* (S1), S100–S122.

Blattman, Christopher, Nathan Fiala, and Sebastian Martinez, "Generating Skilled Self-Employment in Developing Countries: Experimental Evidence from Uganda," *The Quarterly Journal of Economics*, 2014, *129* (2), 697–752.

Bobba, Matteo, Luca Flabbi, and Santiago Levy, "Labor Market Search, Informality, and Schooling Investments," *International Economic Review*, 2022, *63* (1), 211–259.

_ , _ , _ , and Mauricio Tejada, "Labor Market Search, Informality, and On-the-Job Human Capital Accumulation," *Journal of Econometrics*, 2021, *223* (2), 433–453.

Boeri, Tito, Giulia Giupponi, Alan B. Krueger, and Stephen Machin, "Solo Self-Employment and Alternative Work Arrangements: A Cross-Country Perspective on the Changing Composition of Jobs," *Journal of Economic Perspectives*, 2020, *34* (1), 170–195.

Bosch, Mariano and Julen Esteban-Pretel, "Job creation and job destruction in the presence of informal markets," *Journal of Development Economics*, jul 2012, *98* (2), 270–286.

_ and William F. Maloney, "Comparative analysis of labor market dynamics using Markov processes: An application to informality," *Labour Economics*, 2010, *17* (4), 621–631.

Breza, Emily and Supreet Kaur, "Labor Markets in Developing Countries," *Annual Review of Economics*, 2025, *17*, 747–776.

Cagetti, Marco and Mariacristina De Nardi, "Entrepreneurship, frictions, and wealth," *Journal of Political Economy*, 2006, *114* (5), 835–870.

Canavire-Bacarreza, Gustavo Javier, "Rekindling the Reduction of Poverty and Inequality in Ecuador," Technical Report, World Bank Group, Washington, D.C. 2025. English.

Donovan, Kevin, Will Jianyu Lu, and Todd Schoellman, "Labor Market Dynamics and Development," *The Quarterly Journal of Economics*, 2023, *138* (4), 2287–2325.

Esteban-Pretel, Julen, Ryo Nakajima, and Ryuichi Tanaka, "Are contingent jobs dead ends or stepping stones to regular jobs? Evidence from a structural estimation," *Labour Economics*, aug 2011, *18* (4), 513–526.

Falco, Paolo and Luke Haywood, "Entrepreneurship versus Joblessness: Explaining the Rise in Self-Employment," *Journal of Development Economics*, 2016, *118*, 245–265.

Flabbi, Luca and James Mabli, "Household Search or Individual Search: Does it Matter? Evidence from Lifetime Inequality Estimates," Technical Report, Georgetown University, Department of Economics WP 12-02 2010.

_ **and Marco Leonardi**, "Sources of earnings inequality: Estimates from an on-the-job search model of the US labor market," *European Economic Review*, aug 2010, *54* (6), 832–854.

Flinn, Christopher J, "Labour market structure and inequality: A comparison of Italy and the US," *Review of Economic Studies*, aug 2002, *69* (3), 611–645.

_ , "Minimum wage effects on labor market outcomes under search, matching, and endogenous contact rates," *Econometrica*, 2006, *74* (4), 1013–1062.

_ **and James J. Heckman**, "New methods for analyzing structural models of labor force dynamics," *Journal of Econometrics*, 1982, *18* (June 1981), 115–168.

Gindling, T.H. and David Newhouse, "Self-Employment in the Developing World," *World Development*, 2014, *56*, 313–331.

Gouriéroux, Christian and Alain Monfort, *Simulation-Based Econometric Methods* CORE Lectures, Oxford University Press, 2002.

Jolivet, Grégory, Fabien Postel-Vinay, and Jean-Marc Robin, "The empirical content of the job search model: Labor mobility and wage distributions in Europe and the US," *European Economic Review*, 2006, *50* (4), 877–907.

Jovanovic, Boyan, "Job Matching and the Theory of Turnover," *Journal of Political Economy*, 1979, *87* (5), 972–990.

La Porta, Rafael and Andrei Shleifer, "Informality and Development," *Journal of Economic Perspectives*, 2014, *28* (3), 109–126.

Maloney, William F., "Informality revisited," *World Development*, 2004, *32* (7), 1159–1178.

Martínez A., Claudia, Esteban Puentes, and Jaime Ruiz-Tagle, "The Effects of Micro-entrepreneurship Programs on Labor Market Performance: Experimental Evidence from Chile," *American Economic Journal: Applied Economics*, 2018, *10* (2), 101–124.

Meghir, Costas, Renata Narita, and Jean marc Robin, "Wages and informality in developing countries," *American Economic Review*, 2015, *105* (4), 1509–1546.

Moore, Mark A., Anthony E. Boardman, and Aidan R. Vining, "Social Discount Rates for Seventeen Latin American Countries: Theory and Parameter Estimation," *Public Finance Review*, January 2020, *48* (1), 43–71.

Narita, Renata, "Self-Employment in Developing Countries: A Search-Equilibrium Approach," *Review of Economic Dynamics*, 2020, *35*, 1–34.

Tejada, Mauricio M., "Lifetime inequality measures for an emerging economy: The case of Chile," *Labour Economics*, 2016, *42*, 1–15.

World Bank, "Regional Jobs Update: Insights from Labor Force Surveys," Technical Report, World Bank, Washington, D.C. 2025. Confirm exact report title and number (document P506095).

A Data Appendix

This appendix describes the construction of the estimation sample from the ENEMDU microdata.

We construct two datasets from the *Encuesta Nacional de Empleo, Desempleo y Subempleo* (ENEMDU), collected by Ecuador's National Institute of Statistics and Censuses (INEC). The pooled cross-sectional dataset combines the 2022, 2023, and 2024 annual surveys and provides the labor-market state distribution, income distributions, and spell durations. Its unit of observation is an individual in a survey round. The matched-panel dataset combines eight quarterly panels covering 2021–2022 and 2022–2023. Each panel observes the same individual at an initial date (t) and approximately 12 months later ($t + 12$). These matched observations provide the annual transitions across labor-market states and the income growth of workers who remain in the same working state.

We calculate all statistics separately for four groups defined by gender and schooling. Workers with at most completed secondary education are classified as unskilled ($j = U$), while those with completed tertiary education are classified as skilled ($j = S$). We combine this classification with gender to obtain the four worker types used throughout the paper: unskilled men, skilled men, unskilled women, and skilled women.

We classify each individual as out of the labor force (o), unemployed (u), wage-employed (e), or self-employed (s). Individuals who neither work nor actively search for work are out of the labor force, while those without work who actively search are unemployed. Among workers, employees in salaried or dependent positions are wage-employed. Own-account workers and employers are self-employed. ENEMDU also records formal-sector status, but we do not use it to define the labor-market states. Self-employment therefore includes both formal and informal activities.

Labor income is measured as real hourly labor income. We divide total monthly labor income by usual monthly hours worked across all jobs and deflate nominal income using the national consumer price index, with December 2023 as the base period. For self-employed workers, labor income is net of business costs. Spell lengths are measured in months. For wage-employed and self-employed workers, duration is ongoing tenure in the current activity. For unemployed workers, duration is the elapsed time spent searching for work.

We restrict both datasets to individuals aged 25–55 who live in urban areas. We drop observations for which the available information does not permit us to assign a labor-market state or worker type. We also drop wage-employed and self-employed respondents with missing or zero hourly income or missing tenure and unemployed respondents with missing search duration. In the matched-panel dataset, an individual is dropped if the required labor-market or income information is missing in either period. The transition moments condition on labor-market participation and cover movements among u, e, and s.

The cleaned datasets provide 31 estimation moments for each worker group: the four labor-market state proportions; the mean, standard deviation, and fifth percentile of income in e and s; the mean and standard deviation of durations in u, e, and s; the nine transition probabilities among u, e, and s; mean durations in e and s below and above the within-group median income; and mean monthly income growth among workers who remain in e or s. All estimation moments are unweighted. The cross-sectional sample contains 86,323 unskilled men, 49,145 skilled men, 84,752 unskilled women, and 60,331 skilled women.

Table 1: Descriptive Statistics of the Cross-Section

Statistic	Men		Women	
	Unskilled	Skilled	Unskilled	Skilled
Labor Market States (Proportions)				
Non-participation	0.049	0.066	0.415	0.168
Unemployment	0.034	0.051	0.034	0.056
Employment	0.563	0.588	0.236	0.566
Self-Employment	0.354	0.295	0.315	0.209
Labor Income (US dollars per hour)				
Mean of Wage in Employment	3.170	6.218	2.861	5.665
Std of Wages in Employment	1.897	4.544	1.698	3.499
Mean of Earnings in Self-Employment	2.863	4.741	2.309	3.825
Std of Net Earnings in Self-Employment	3.596	6.985	2.854	4.223
Durations (years, except unemployment in months)				
Unemployment	2.991	4.307	3.562	4.43
Employment	10.278	8.184	7.052	8.713
Employment (with wages below the median)	9.842	5.23	6.583	5.274
Employment (with wages above the median)	11.123	9.074	9.308	10.245
Self-Employment	12.507	9.183	9.290	7.265
Self-Employment (with earnings below the median)	11.661	7.937	9.027	6.160
Self-Employment (with earnings above the median)	13.684	9.791	9.653	7.814
Number of Observations	86323	49145	84752	60331

Table 2: Yearly Transition Matrix

From / To	Unemployment	Employment	Self-Employment
Men, Unskilled			
Unemployment	0.242	0.495	0.263
Employment	0.027	0.845	0.128
Self-Employment	0.019	0.199	0.782
Men, Skilled			
Unemployment	0.304	0.417	0.279
Employment	0.026	0.885	0.089
Self-Employment	0.031	0.189	0.78
Women, Unskilled			
Unemployment	0.299	0.321	0.379
Employment	0.036	0.846	0.118
Self-Employment	0.024	0.113	0.863
Women, Skilled			
Unemployment	0.352	0.379	0.268
Employment	0.021	0.929	0.05
Self-Employment	0.054	0.196	0.75

Table 3: Estimated Parameters

Parameter	Men				Women			
	Unskilled		Skilled		Unskilled		Skilled	
	Value	Std.Err.	Value	Std.Err.	Value	Std.Err.	Value	Std.Err.
λ^u	0.129	0.0055	0.175	0.0038	0.067	0.0098	0.162	0.0011
λ^e	0.056	0.0025	0.053	0.0024	0.02	0.0007	0.086	0.0033
λ^s	0.119	0.0076	0.17	0.0063	0.254	0.0067	0.147	0.0232
δ^u	0.032	0.0026	0.032	0.0026	0.059	0.0046	0.052	0.0042
δ^e	0.066	0.0007	0.071	0.0025	0.064	0.007	0.016	0.0027
δ^s	0.078	0.0086	0.104	0.0002	0.137	0.0145	0.183	0.011
η_0^e	0.001	0.0001	0.005	0.0003	0.003	0.0002	0.003	0.0004
η_0^s	0.0	0.0002	0.0	0.0	0.0	0.0002	0.002	0.0002
η_1^e	0.056	0.0081	0.057	0.0053	0.071	0.0075	0.087	0.007
η_1^s	0.049	0.0088	0.06	0.0024	0.053	0.0053	0.058	0.004
μ_e	-0.375	0.0257	0.54	0.0309	-0.251	0.0162	0.33	0.0262
σ_e	0.768	0.0082	0.711	0.0124	0.622	0.0074	0.704	0.0123
μ_s	-0.815	0.0321	-0.578	0.0278	-0.886	0.0325	-0.825	0.0303
σ_s	0.887	0.0007	0.956	0.0071	0.9	0.0096	1.063	0.0107
b	0.096	0.0188	0.132	0.0119	0.04	0.0163	0.107	0.0121
γ	1.463	0.0114	0.658	0.038	0.434	0.0076	0.473	0.0123
Loss	33099		20477		18486		23244	

NOTE: Standard errors are computed using the bootstrap method. In the estimation procedure, 15,000 individuals are simulated over a 50-year period. The discount rate ρ is set to 0.1.

Table 4: Implied Values

Labor Market State	Men		Women	
	Unskilled	Skilled	Unskilled	Skilled
Average arrival of job offers (months)				
Unemployed	7.76	5.713	14.886	6.189
Employees	17.935	18.874	49.318	11.656
Self-Employed	8.388	5.875	3.933	6.823
Average arrival of self-employed opportunities (months)				
Unemployed	31.7	31.162	16.81	19.204
Employees	15.175	14.182	15.543	63.569
Self-Employed	12.857	9.575	7.309	5.47
Average arrival of termination shocks (years)				
Employees	15.514	9.243	10.031	9.279
Self-Employed	21.233	17.343	18.673	12.266
Offered wages (U.S. dollars per hour)				
Mean	0.923	2.21	0.944	1.781
Standard Deviation	0.828	1.791	0.649	1.427
Potential self-employment income (U.S. dollars per hour)				
Mean	0.656	0.886	0.618	0.771
Standard Deviation	0.717	1.082	0.69	1.115

Table 5: Moments Fit

Statistic	Men				Women			
	Unskilled		Skilled		Unskilled		Skilled	
	Model	Data	Model	Data	Model	Data	Model	Data
np	0.047	0.049	0.061	0.066	0.425	0.415	0.179	0.168
u	0.038	0.036	0.06	0.054	0.048	0.059	0.066	0.068
e	0.575	0.592	0.651	0.63	0.415	0.403	0.669	0.681
s	0.387	0.372	0.289	0.316	0.537	0.538	0.265	0.251
$E[w\|e]$	3.325	3.17	6.553	6.218	3.01	2.861	5.712	5.665
$SD[w\|e]$	1.87	1.897	3.383	4.544	1.25	1.698	2.75	3.499
$P5[w\|e]$	0.811	0.863	2.437	1.556	0.671	0.728	2.365	1.61
$E[y\|s]$	3.07	2.863	5.09	4.741	2.934	2.309	4.224	3.825
$SD[y\|s]$	2.281	3.596	4.758	6.985	2.209	2.854	4.168	4.223
$P5[y\|s]$	0.336	0.501	0.305	0.733	0.296	0.298	0.23	0.459
$E[t\|u]$	6.634	2.991	7.873	4.307	7.196	3.562	8.767	4.43
$SD[t\|u]$	6.811	3.556	8.173	4.093	7.347	3.819	8.808	4.22
$E[t\|e]$	9.721	10.278	6.943	8.184	7.603	7.052	7.813	8.713
$SD[t\|e]$	10.386	8.715	7.706	7.47	8.417	7.321	8.467	7.643
$E[t\|s]$	9.606	12.507	7.431	9.183	8.489	9.29	5.405	7.265
$SD[t\|s]$	10.901	9.618	10.157	8.021	10.297	9.08	7.852	7.123
$Pr[u\|u]$	0.885	0.884	0.891	0.902	0.894	0.901	0.9	0.913
$Pr[e\|u]$	0.078	0.077	0.07	0.057	0.053	0.045	0.056	0.047
$Pr[s\|u]$	0.037	0.039	0.039	0.041	0.053	0.054	0.043	0.04
$Pr[u\|e]$	0.005	0.004	0.008	0.004	0.008	0.005	0.007	0.003
$Pr[e\|e]$	0.991	0.983	0.987	0.988	0.985	0.984	0.991	0.993
$Pr[s\|e]$	0.004	0.012	0.004	0.008	0.007	0.01	0.002	0.004
$Pr[u\|s]$	0.004	0.003	0.004	0.004	0.004	0.003	0.006	0.008
$Pr[e\|s]$	0.008	0.02	0.014	0.018	0.008	0.01	0.01	0.018
$Pr[s\|s]$	0.988	0.978	0.982	0.978	0.987	0.986	0.984	0.974
$E[t\|e, w < w^{median}]$	6.306	9.842	4.767	5.23	5.406	6.583	5.281	5.274
$E[t\|e, w > w^{median}]$	13.138	11.123	9.121	9.074	9.801	9.308	10.345	10.245
$E[t\|s, y < y^{median}]$	4.711	11.661	2.254	7.937	3.572	9.027	1.831	6.16
$E[t\|s, y > y^{median}]$	14.501	13.684	12.607	9.791	13.405	9.653	8.979	7.814
$E[\Delta w\|e]$	0.003	0.003	0.002	0.004	0.001	0.004	0.003	0.003
$E[\Delta y\|s]$	0.008	0.007	0.042	0.007	0.016	0.007	0.034	0.006

Table 6: Income and Welfare by First-Job and Most-Time Self-Employment Status

Worker type	Overall		Start as SE		Most time as SE	
	Mean	Std	Mean	Std	Mean	Std
Panel A. Cross-sectional Income (in Steady State)						
Men, unskilled	3.109	2.051	3.124	2.078	3.613	2.324
Men, skilled	5.74	3.812	5.758	3.858	6.884	4.543
Women, unskilled	2.826	1.849	2.835	1.851	3.025	1.939
Women, skilled	4.937	3.418	4.977	3.418	5.27	3.908
Aggregated	3.977	3.023	4.014	3.049	3.766	2.902
Panel B. Lifetime Income						
Men, unskilled	246.462	104.254	246.579	106.475	241.124	113.27
Men, skilled	502.323	193.458	488.693	195.902	496.759	250.193
Women, unskilled	233.4	90.376	239.591	94.314	231.299	94.279
Women, skilled	428.928	179.044	420.125	189.039	411.327	202.876
Aggregated	335.019	181.458	335.084	180.842	283.207	161.118
Panel C. Proportion of SE jobs with low value with respect of E jobs						
Men, unskilled	0.759	0.323	0.756	0.313	0.65	0.33
Men, skilled	0.827	0.297	0.828	0.288	0.645	0.356
Women, unskilled	0.629	0.317	0.615	0.312	0.585	0.306
Women, skilled	0.849	0.268	0.837	0.269	0.764	0.289
Aggregated	0.793	0.405	0.775	0.418	0.668	0.471
Panel D. Welfare (in Steady State)						
Men, unskilled	31.398	–	31.497	–	34.784	–
Men, skilled	57.62	–	57.706	–	66.757	–
Women, unskilled	28.496	–	28.544	–	29.847	–
Women, skilled	49.616	–	49.773	–	52.265	–
Aggregated	40.025	–	40.285	–	37.034	–

NOTE: "Start as SE" restricts to individuals whose first job is self-employment; "Most time as SE" restricts to those who spend the majority of their labor-market spells in self-employment. Cross-sectional statistics are computed at the steady state; lifetime statistics are over the full simulated career.

Table 7: Counterfactual: Labor-Market Outcomes without Self-Employment

Statistic	Men		Women		Aggregated
	Unskilled	Skilled	Unskilled	Skilled	
Labor Market States (Proportions)					
Non-participation	0.08 (1.73)	0.08 (1.33)	0.63 (1.49)	0.21 (1.18)	–
Unemployment	0.11 (3.29)	0.15 (2.72)	0.19 (4.12)	0.1 (1.58)	0.13 (2.83)
Employment	0.89 (1.55)	0.85 (1.29)	0.81 (1.98)	0.9 (1.33)	0.87 (1.51)
Accepted Wage for Employees (in US$ per hour)					
Mean	2.74 (0.82)	5.73 (0.87)	1.69 (0.56)	4.86 (0.85)	3.6 (0.78)
Std	1.57 (0.84)	2.89 (0.9)	0.84 (0.67)	2.69 (0.97)	2.58 (0.91)
Mean Durations (U in months, E in years)					
Unemployment	24.19 (3.41)	20.87 (2.63)	27.45 (3.78)	13.84 (1.55)	22.71 (2.89)
Employment	10.02 (1.03)	7.01 (1.01)	7.38 (0.97)	6.85 (0.87)	8.16 (1.01)
Lifetime Income (in US$ per hour)					
Mean	203.57 (0.83)	454.06 (0.9)	125.97 (0.54)	390.44 (0.91)	278.34 (0.83)
Std	96.17 (0.92)	185.38 (0.96)	56.48 (0.62)	155.82 (0.87)	179.9 (0.99)

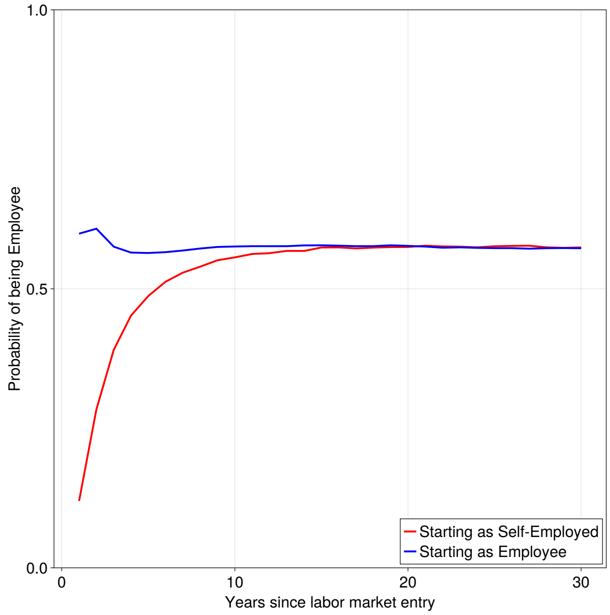
NOTE: Each cell shows the counterfactual (no self-employment) value and, in parentheses, its ratio to the benchmark value.. "–" marks statistics that are undefined when self-employment is shut down.

Table 8: Counterfactual: Income Inequality without Self-Employment

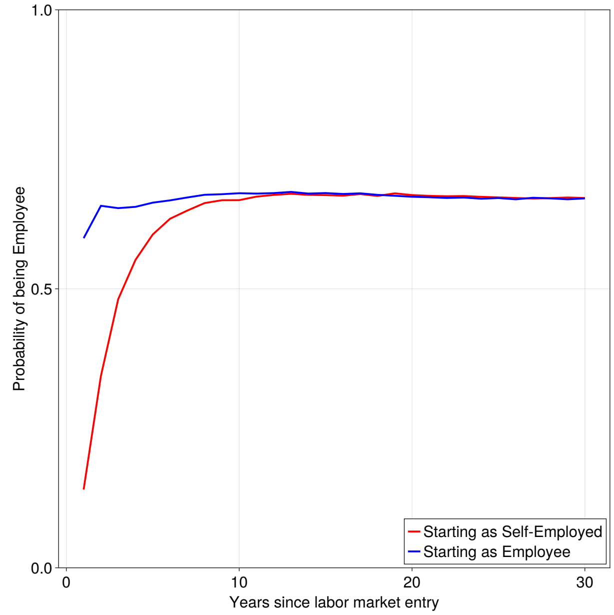
Statistic	Men		Women		Aggregated
	Unskilled	Skilled	Unskilled	Skilled	
Panel A. Cross-Sectional Income					
Gini (excl. U)	0.28 (0.89)	0.25 (0.8)	0.25 (0.81)	0.28 (0.89)	0.36 (1.0)
Gini (all)	0.35 (1.05)	0.36 (1.03)	0.38 (1.14)	0.35 (0.99)	0.44 (1.14)
Theil (excl. U)	0.13 (0.78)	0.1 (0.63)	0.1 (0.64)	0.13 (0.76)	0.21 (0.98)
Theil (all)	0.23 (1.16)	0.25 (1.15)	0.29 (1.38)	0.22 (0.98)	0.34 (1.31)
P_{90}/P_{10}	62.57 (3.66)	89.21 (2.44)	85.78 (3.15)	98.25 (1.97)	118.06 (3.65)
P_{90}/P_{50}	4.76 (1.07)	4.65 (1.03)	5.23 (1.2)	4.61 (0.98)	7.27 (1.31)
P_{50}/P_{10}	37.6 (3.62)	54.83 (2.44)	53.05 (3.13)	60.32 (2.0)	65.79 (3.54)
Panel B. Lifetime Income					
Gini	0.23 (1.12)	0.2 (1.07)	0.23 (1.22)	0.2 (0.99)	0.34 (1.23)
Theil	0.09 (1.25)	0.07 (1.14)	0.09 (1.41)	0.07 (0.95)	0.18 (1.49)
P_{90}/P_{10}	4.32 (1.22)	3.73 (1.13)	4.45 (1.36)	3.44 (0.98)	8.3 (1.63)
P_{90}/P_{50}	2.95 (1.11)	2.61 (1.05)	2.86 (1.16)	2.56 (0.98)	4.55 (1.3)
P_{50}/P_{10}	2.79 (1.18)	2.53 (1.11)	2.97 (1.32)	2.33 (0.99)	5.15 (1.59)

NOTE: Each cell shows the counterfactual (no self-employment) value and, in parentheses, its ratio to the benchmark value. "Excl. U" excludes unemployed workers from the cross-sectional income distribution. Inequality ratios P_{90}/P_{10}, P_{90}/P_{50}, and P_{50}/P_{10} are computed as the mean of the top tail divided by the mean of the bottom tail (not strict quantile ratios).

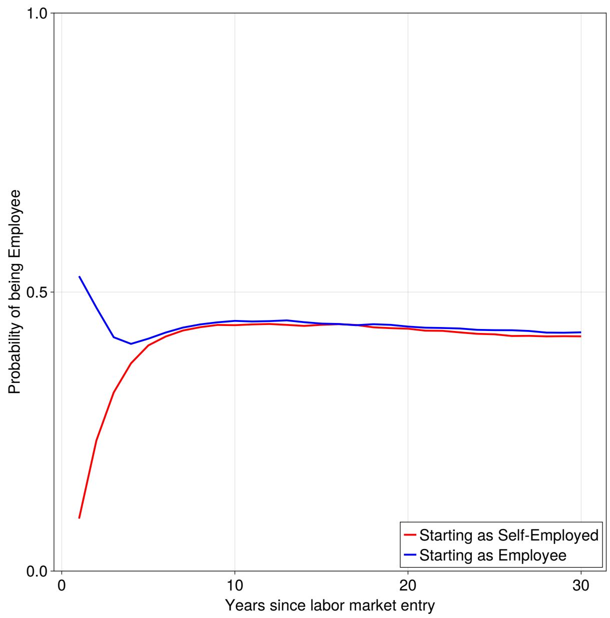
Figure 1: Likelihood of Wage Employment Conditional on Starting Job



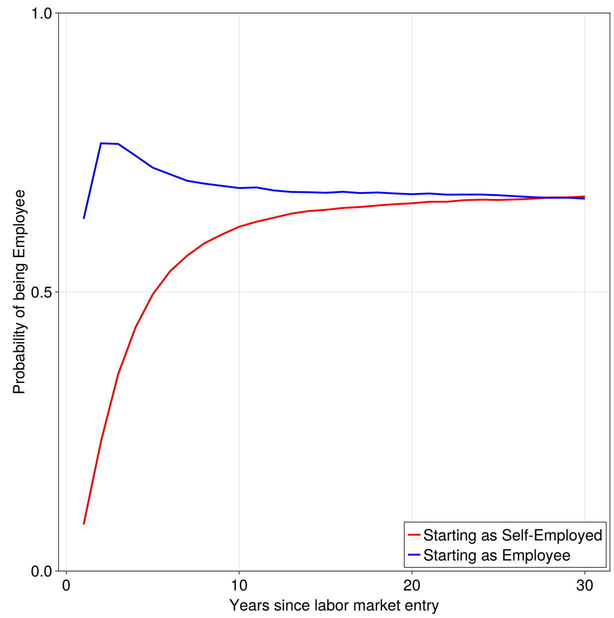
(a) Unskilled Men



(b) Skilled Men

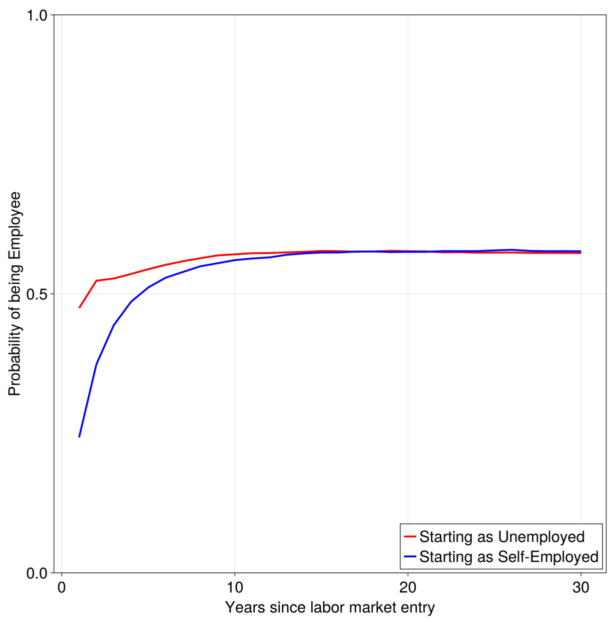


(c) Unskilled Women

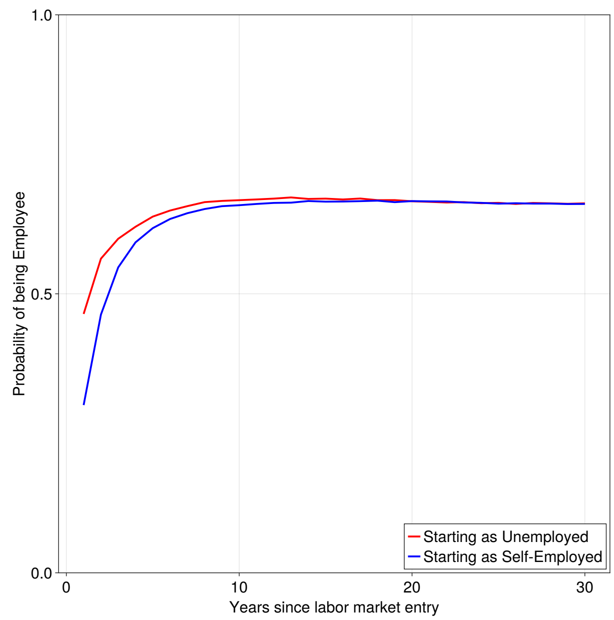


(d) Skilled Women

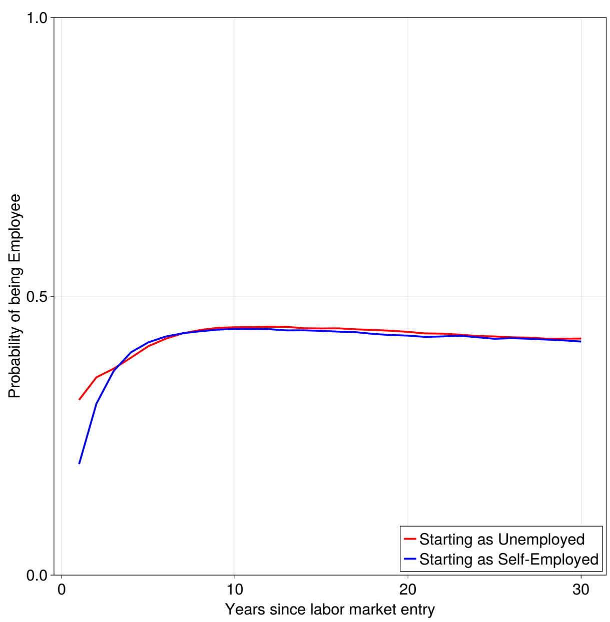
Figure 2: Likelihood of Wage Employment when Starting Career as Unemployed or Self-Employed



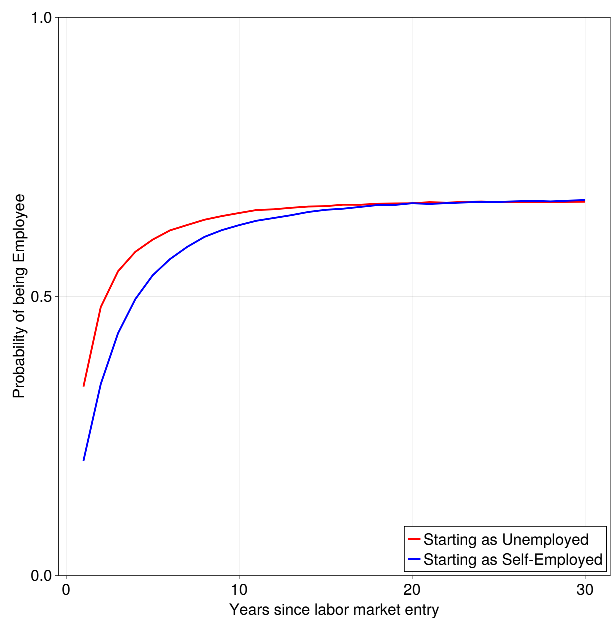
(a) Unskilled Men



(b) Skilled Men



(c) Unskilled Women



(d) Skilled Women